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ARTICLE *in* RESEARCH IN INTERNATIONAL BUSINESS AND FINANCE · FEBRUARY 2006

DOI: 10.1016/j.ribaf.2005.04.001 · Source: RePEc

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Research in International Business and Finance 20 (2006) 45–61
RESEARCH IN INTERNATIONAL BUSINESS AND FINANCE
www.elsevier.com/locate/ribaf

An investigation of bond term premia in international government bond indices

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Received 25 January 2005; received in revised form 6 April 2005; accepted 12 April 2005
Available online 7 July 2005

Abstract

In this study we use the monthly excess holding period yields (EHPY), and their volatility for five government bond indices markets, in order to test the popular expectations theory of the term structure and to assess whether there are patterns in risk premia, which are common across these markets. For this reason, a GARCH-M model is used. The empirical results derived indicate that in most cases the Expectation Hypothesis is rejected and short medium-term government bond yields display serial correlation and co movement in the majority of the countries. Additionally, the term premia at the long end of the maturity structure are time varying in the USA, the UK and Canada. The risk factor is statistically significant as an explanatory variable for risk premia, only for 7 years maturity bonds in the USA and the UK.

JEL classification: E43; C22; C52

Keywords: GARCH-M; Excess returns; Term premium

1. Introduction

One of the main characteristics of financial time series is the fact that their volatility changes over time. This feature has to be taken into consideration in various areas of the

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doi:10.1016/j.ribaf.2005.04.001

financial literature, like investment theory, Capital Asset Pricing Model (hereafter CAPM), options pricing, etc. As financial market conditions have become more volatile in recent years, deviations of bond market efficiency and the term premium behavior has increasingly concerned researchers and investors.

In this paper we investigate the behavior of term premia in various medium and long-term government bond indices. We test internationally the expectations theory of the term structure of interest rates and assess whether there are common patterns in excess holding period yields in different markets.

The expectations theory of the term structure starts from the observation that investment strategy of rolling over short-term bonds is an alternative to holding a long-term bond. According to the theory the expected rates of return on these alternative investment strategies differ by a constant premium. There are several reasons for exploring this subject from the point of view of the concerns of researchers, central banks and bond portfolio managers:

(a) In the world capital markets the value of fixed income securities exceeds the total value of equity. Due to the size and importance of the bond market, there will continue to be substantial fixed income research. The bond market indices can provide accurate and timely measurement of the risk/return of bond assets and the characteristics of the market. Given this economic dominance, it is difficult to understand why there has not been many studies concern and analyze bond market indices. Part of the reason for a lack of analysis of bond indices is the very short of their history. International government bond indices were initially created in the middle of 1980s. The amount of money invested in bond index funds grew from $3 billion in 1984 to over $100 billion in 1995 (Fabozzi, 1999).
(b) Second, the bond indices have several significant uses, including acting as performance benchmarks, as means to determine risk/return, as inputs into the asset allocation decision and as means to test term structure theories more efficiently for many countries.
(c) Another reason is the increased internationalization of security markets and the implied interdependence between financial markets. These interdependencies mean that interest rate cannot always be viewed as being completely determined in domestic money and capital markets.

This paper directly identifies the presence of time varying term premium and its behavior within some countries over a specific maturity band. The methodology of Engle et al. (1987) is utilized and maximum likelihood estimation of the GARCH-M model is used to assess the presence of ARCH effect in the data set at hand. Our purpose is to show if there are common patterns in term premia across different countries as well as to test if the conditional variance influences significantly the behavior of the term premia at very high frequencies. Several general regression models are tested in order to better express the time varying term premium. This is important because results that contradict the efficient market hypothesis can generally be explained by the presence of time varying risk premium. However, the results are not universally significant and depend on bond maturity and the characteristics of each national market.

Our empirical results provide additional international evidence against the expectations theory as a model of the behavior of medium- and long-term government bond yields.

Additionally, the empirical results indicate the existence of a time varying term premium at the long end of the term structure in the USA, the UK and Canada. Generally, the risk factor is statistically significant as an explanatory variable for risk premia only for 7 years maturity bonds in the USA and the UK.

The structure of the paper is as follows. In Section 2, we provide a brief review of the recent relative literature in modeling term premia. In Section 3, we present the dataset used in our empirical analysis. In Section 4, we introduce the econometric model used and the methodology followed in order to provide an explanation of the determinants of risk premia by first using the traditional CAPM and then using variants of the GARCH-in-mean model. In Section 5, all empirical results are presented and discussed. The final section concludes the paper.

2. Literature review

Research in testing Asset Pricing Theories for bonds can be divided into two basic categories. The first involves the study of changing risk premiums over time using economic variables (Fama and French, 1989; Stambaugh, 1988). The second type of research involves tests of the Cox et al. (1985) valuation model and its generalizations (Longstaff and Schwartz, 1992).

Our study belongs to the first category, which can be further divided into two subcategories according to the data structure of short- and long-term bonds. The vast majority of research deals with short-term instruments (see among others, Levy and Brooks, 1989; Backus et al., 1989; Richardson et al., 1992; Alles, 1995; Lee and Hoje, 1996). Previous works concerning long-term bonds are limited (Shiller et al., 1983; Fama, 1984; Hall and Miles, 1992; Bansal and Coleman, 1996; Bekdache and Baum, 1998).

The expectation theory of the term structure is among the most popular models of interest rate behavior. According to this theory, “if all agents are risk neutral and are concerned only with expected return then the expected one-period yield (holding period yield, say over a month) on all bonds, no matter what their maturity, would be equalized and would equal to the known safe rate of return RF_t on one-period asset (e.g. 1 month Treasury Bill)” (Cuthbertson, 1996).

Another attempt to explain the term premium under a constancy condition is that given by the CAPM equilibrium model (Bisignano, 1987). This study provides evidence for the CAPM, as explaining model of the term structure in a series of countries. However several problems with the underlying assumptions of the model can lead to simplify results and possible errors.

However, several papers provide evidence against the expectations theory and against the efficient market theory, under which excess returns are unpredictable (Jones and Roley, 1983; Campbell and Shiller, 1991; Elton et al., 1996; Bekdache and Baum, 1998). For the US data this theory does not perform well at the short end of the maturity spectrum, and this may be due to data deficiencies or to the presence of a time varying term premium. But for longer coupon paying bonds the results are opposite, with the exception of the UK (Simon, 1989; Campbell and Shiller, 1991; Cuthbertson, 1996). But multi-country studies are rare, with some exceptions (Mankiw, 1986; Bisignano, 1987; Hall and Miles, 1992).

Others have examined the theoretical and empirical models that may derive the risk premium. Among the models which may generate a time varying risk premium, attention is required to the equilibrium, dynamic, optimizing models of asset pricing. Lucas (1982) presents such a model in an international context. The theoretical research, followed this seminal paper, has been motivated by the operationalizing of the Euler conditions in the cash in advance model. Starting with the optimization of the first order condition there are three different approaches that have been developed. The CAPM/latent variable approach (Hodrick, 1987), the risk-free premium (Hodrick and Srivastava, 1986; Campbell and Clarida, 1987) and the Autoregressive Conditional Heteroskedasticity (ARCH) framework (Engle, 1982).

The latter, together with the diffusion models with stochastic volatility (Chesney and Scott, 1989; Hull and White, 1987; Wiggins, 1987) is the most popular way of implementing the first order conditions and coping with volatility changes. The most common used ARCH and GARCH models as introduced by Engle (1982) and Bollerslev (1986) take h as a linear function of lagged conditional variances and squared residuals. Theoretically these models are characterized by linearity as they imply an ARMA equation for the squared prediction errors u_t^2 allowing a full study of the distributional properties of u_t as well as easier statistical inference. But although GARCH models take into account that asset returns are fat tailed, normalizing the returns by the conditional variances does not completely eliminate this leptokurtosis (Rabemananjara and Zakoian, 1994). Nelson (1991) first cites a number of drawbacks in the use of GARCH models. The choice of quadratic form of the conditional variance and the parameters' non-negativity are some of these limitations. Unlike GARCH specification, the EGARCH model, specified in logarithms, does not impose the non-negativity constraints on parameters. The EGARCH model has been used to examine the interest rates by Brunner and Simon (1996).

A number of alternative parameterizations for the function h are suggested in the literature. Researchers use the ARCH-M class of models as proposed by Engle et al. (1987). Bekdache and Baum (1998) identify the presence of heteroskedastic errors in models of excess holding period yield and investigate the impact of sampling aggregation on the U.S. Treasury securities. They estimate models of the GARCH-M family to weekly, quarterly and monthly bond data. They report that models of excess returns, incorporating the spread and the conditional heteroskedasticity as explanatory variables, are quite robust to the various data sets and in a narrow national market. Hall and Miles (1992) fit multivariate GARCH-M type models to the monthly series of the major bond markets in an international environment with similar results.

3. Data presentation

Monthly data from five different government bond markets are examined. We use the Datastream Bond Indices calculated according to EFFAS (European Federation of Financial Analyst Societies). More specifically, we examine the benchmark total return indices for a variety of bond maturities of 2, 3, 5, 7 and 10 years for the USA, Germany, Canada, the UK, and France. Benchmark indices are based on single bonds which are usually the most representative bonds available for the given maturity band at each point in time. All

Table 1
Statistics for excess holding period yields (EHPY)

Bond maturity	Statistics	USA	UK	Germany	France	Canada
2 years	Mean	1.08	−0.23	0.62	0.58	1.28
	Variance	50.04	101.87	34.33	46.05	145.96
	t-Test	1.85*	−0.28	1.09	0.96	1.36
3 years	Mean	1.66	0.35	1.13	1.05	1.43
	Variance	115.18	182.28	69.08	91.65	250.46
	t-Test	1.90*	0.32	1.37	1.18	1.17
5 years	Mean	2.23	0.69	1.44	1.53	1.94
	Variance	291.34	418.23	138.72	195.50	417.20
	t-Test	1.65	0.47	1.23	1.11	1.20
7 years	Mean	2.90	1.70	2.20	2.68	3.27
	Variance	475.65	580.20	234.65	328.45	553.81
	t-Test	1.76	0.94	1.50	1.50	1.75
10 years	Mean	3.32	2.42	1.87	3.07	3.29
	Variance	710.36	847.68	414.49	509.47	794.04
	t-Test	1.60	1.09	1.13	1.47	1.48

Note: an asterisk * indicates statistically significant at 5% level of significance.

series cover the period from 1 January 1987 to 1 June 2000, which makes a sample size of 161 observations. For CAPM estimation the sample size is from 1 October 1991 to 1 June 2000 due to the availability of the World Market Index Fund provided by Union Bank of Switzerland.

Yields to maturity and holding period returns (HPY) are converted into annual equivalents. The risk-free rate for each country is the 1-month Eurocurrency rate measured on the first day of each month, which is also provided by Datastream. The excess holding period yield (EHPY) for a particular month is measured as the return from buying a bond at the beginning of the month and selling it at the month's end, having collected coupons that are payable within the holding period, minus the risk-free rate. Table 1 gives some descriptive statistics for each series of EHPY. For each bond maturity in the first column, there are three corresponding rows giving the mean, the variance and a *t*-test for EHPY. The *t*-statistic refers to an autocorrelation adjusted test in case the mean of the series is not zero (Harvey, 1993). In our case there are no significant results.[1]

Looking at the above table we conclude that risk premia and their variances increase with maturity. This seems sensible and consistent with the liquidity preference hypothesis since agents who hold longer maturity bonds in order to obtain higher average excess holding period yield, also experience increased risk. The variability of returns is higher in the USA, the UK and Canada compared to the other countries allowing speculation. This can be explained by high level of interest rates in these countries coupled with the size and the liquidity of the markets. In all cases, except for the UK 2-year bond, the risk premia are

[1] It is worth mentioning that one possible explanation for the lack of significance may be in the aggregation of our data set. For this reason, contacting a joint test of statistical significance, we end up again to statistically insignificant results.

positive, though statistically insignificant over the whole period. However by investigating further, there is some evidence of statistically significant EHPY over sub-periods which imply time variation of bond premium. This is consistent with previous studies like for instance, Lauterbach (1989).

Over the decade 1980–1990 fixed income investors benefited from both the decline in nominal interest rates which resulted from falling inflation rates and provided capital gains to bond holders while increasing real interest rates at the same time. However the decline in nominal interest rates over the period 1992–1998 was small compared to the decreases which occurred between 1982 and 1992.

The relatively high degree of comovement of long rates between the USA, Canada and the UK coupled with the global nature of the 1988 and 1994 bond market reversals, has lead to speculation that international factors considered also very important compared to domestic determinant of long rates (see Fig. 1). The IMF's World Economic Outlook (1999) documents an average growth rate (1.9%) of world output in the 1990s that is below the average growth rate of the 1980s (3.1%). This decrease of economic growth has hit USA as well as Europe. In order to face this problem United States and Europe commonly adopted the policies of deregulation and inflation-fighting.

4. The econometric methodology and estimation

The first section attempts to utilize the well-known CAPM framework, to explain the behavior of term premium, on government bond indices of varying maturities for the five countries. This is the first attempt in order to get some understanding of the behavior of risk premium. The second trial, with a general form of a GARCH-in-mean regression model, allowing time varying risk, is presented next.

Let us first define the excess holding period yield at time t (EHPY_t) on an n-period bond as the difference between the holding period return on the bond (HPY_t) and the one period risk-free interest rate (RF_t), that is:

$$\mathrm{EHPY}_t = \mathrm{HPY}_t - \mathrm{RF}_t \tag{1}$$

The spread for an n-period bond (S_t) at time t is defined as the difference between the yield to maturity (YTM_t) of the security at time t and the risk-free interest rate (the latter will be the Eurocurrency rate), that is:

$$S_t = \mathrm{YTM}_t - \mathrm{RF}_t \tag{2}$$

Firstly, we test the Expectation Hypothesis (EH) across the national markets. According to the Pure Expectation Hypothesis (PEH) and under the assumption of rational expectations:

$$\mathrm{HPY}_{t+1} = E_t\mathrm{HPY}_{t+1} + u_{t+1}, \quad \text{where } E_t(u_{t+1}|\Omega_t) = 0 \tag{3}$$

the term premium is zero for all maturities (Cuthbertson, 1996):

$$\mathrm{HPY}_{t+1} - \mathrm{RF}_t = u_{t+1}, \quad \text{where } E_t(u_{t+1}|\Omega_t) = 0 \tag{4}$$

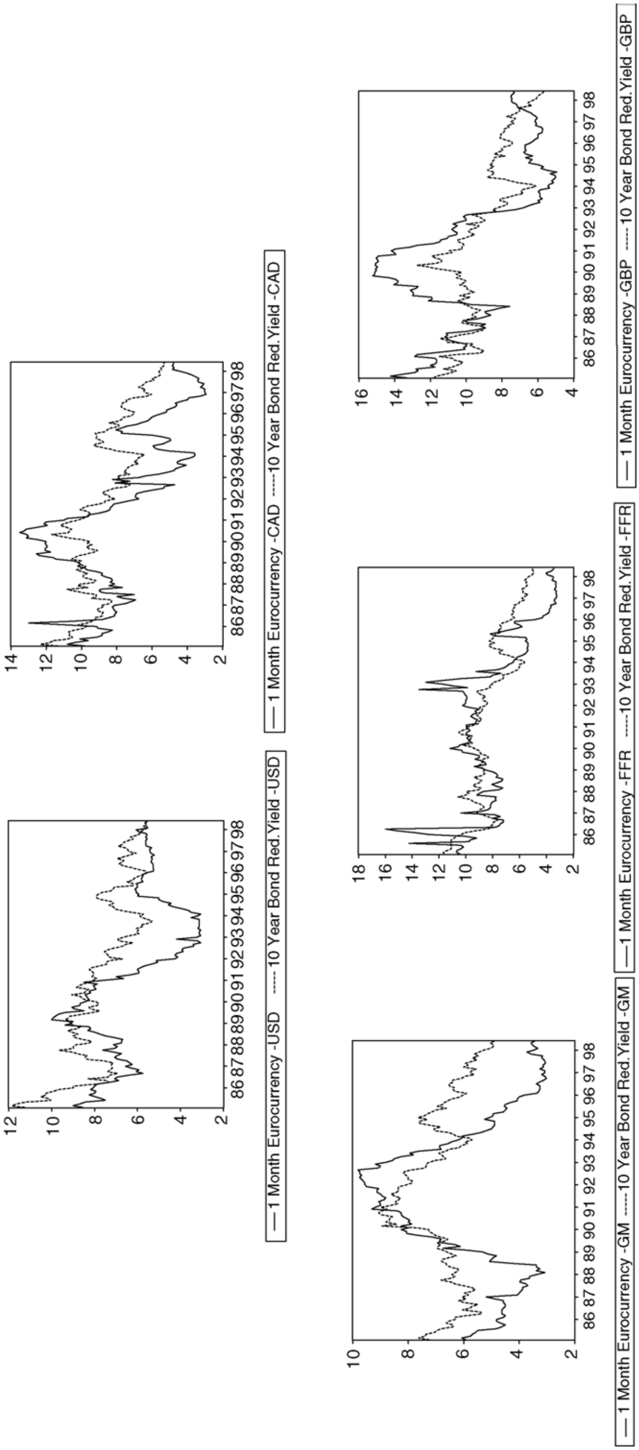

Fig. 1. Long and short term securities yields.

Hence, according to the PEH the excess holding period yield should have a zero mean, be independent of all information at time t and be serially uncorrelated. According to a weaker version, which we shall refer to from now on as the Expectation Hypothesis (EH), the term premium is constant. A standard test of the expectation theory is to regress EHPY on some chosen economic variables and test whether it has the estimated coefficient equal to zero and simultaneously check for serial correlation. Campbell (1995) and Mankiw (1986) among others have found that excess returns depend on variables such as the lagged values of the spread between long rates and short rates. Our approach is to adopt a model of the form:

$$\mathrm{EHPY}_t = c + a\, S_{t-1} + u_t \tag{5}$$

where u_t is an error term, and to test whether a equals zero. We also apply a test for serial correlation in u_t's. If either of the null hypotheses is rejected, there is some evidence against the efficiency conditions and the Expectations Hypothesis. Either the government security markets are "inefficient" in some sense or there exists considerable time dependence in risk or liquidity premiums.

4.1. Explaining term premia using the traditional CAPM

According to the traditional Sharpe–Lintner CAPM, risk is associated with the covariability of an asset's return with the average return on a broad market portfolio:

$$E_t\mathrm{HPY}_t^n - \mathrm{RF}_t = a_n + b_n(E_t R_t^m - \mathrm{RF}_t) + u_t^n, \quad \text{where } u_t^n \sim N(0, \sigma^2) \tag{6}$$

where $E_t\mathrm{HPY}_t^n$ is the expected return on bond with maturity n; RF_t the "risk-free" rate of return; $E_t R_t^m$ the expected return on the "world market" bond portfolio and

$$b_n = \frac{\mathrm{cov}(\mathrm{HPY}^n, R^m)}{\mathrm{var}(R^m)}$$

The "market" portfolio is approximated by the Global Bond Mutual Fund, created by the Union Bank of Switzerland. A monthly return is computed in local currency so that the sample period available for estimation of Eq. (6) is October 1991 to June 2000. In Fig. 2 we can see some details about the structure of the index fund.

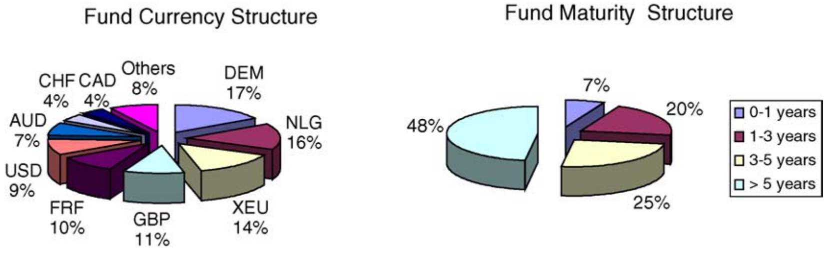

Fig. 2. Structure of index fund.

The fund is invested on a broadly diversified basis in the international bond markets and the maturity of the investments in the fund portfolio is actively managed. The investment objective is to achieve the highest possible return based on the performance of the international bond markets. The fund is consistent of high-quality international bonds

While the traditional CAPM still receives considerable empirical attention, the majority of the theorists in this area argue that it does not provide complete description of asset returns (Jensen, 1979). Interpreting the term premium in government bond indices as a simple CAPM covariance factor may be shortsighted and lead us to wrong results because of the unrealistic assumptions of the model about term premium constancy. Risk-free interest rate and World bond market Portfolio is very difficult to be defined. As a result of all the above we are going on further by investigating term premium using a more general model allowing for time variation.

4.2. A GARCH-M model

In this section we investigate the need to extend model (5) to cope with heteroskedasticity and risk as an explanatory variable. For this purpose we introduce more explicitly the statistical models used to explain the excess holding period yield. Appropriate models for returns are models of the GARCH-M family (Engle et al., 1987).

According to this model, the conditional mean is an explicit function of the conditional variance or standard deviation of the forecast errors. In this model an increase in the conditional variance will be associated with an increase or a decrease in the conditional mean depending on the sign of the partial derivative of the function with respect to the variance. The basic idea in this model, as in many theories in financial economics, is the use of a measure of risk as an explanatory variable. To the extent that the conditional variance of an error term is a measure of risk it seems reasonable that the variance should enter the regression function as a measure of the risk premium.

In most applications, a GARCH in-mean term h_t^2 is suggested (Bollerslev et al., 1992; Engle, 1982). In our work, the general form of the model is

$$\mathrm{EHPY}_t = c + a\, S_{t-1} + b\mathrm{EHPY}_{t-1} + \delta h_t^2 + u_t \tag{7}$$

$$u_t|\Omega_{t-1} \sim N(0, h_t^2) \tag{8}$$

$$h_t^2 = \gamma_0 + \sum_{i=1}^{n} \gamma_i u_{t-i}^2 + \zeta' Z_t \tag{9}$$

where X_t and Z_t are vectors of weakly exogenous conditioning variables. Eq. (7) is an ARCH-M regression model for the mean EHPY. Eqs. (8) and (9) give the GARCH structure of its conditional time varying variance where Ω denotes an appropriate information set up to time $t-1$. Eq. (9) is a non-stochastic equation and h_t^2 is the conditional variance of u_t at period t relying on the information available up to period $t-1$.

Bollerslev (1986) extended this formulation imposing smoother behavior on the conditional second moments as function of their own lagged values and the squares and cross-products of lagged forecast errors, as

$$h_t^2 = B_0 + \sum_{i=1}^{n} B_i u_{t-i}^2 + \sum_{i=1}^{P} \Gamma_i h_{t-1}^2 + \zeta' Z_t \tag{10}$$

where B_i and Γ_i are coefficients. Putting all the system parameters into a single vector $\xi = (c, a, \delta, B_0, B_1, B_2, \ldots, B_n, \Gamma_1, \Gamma_2, \ldots, \Gamma_P)$, the log-likelihood conditional on initial values for a sample of T observations is proportional to

$$L(\xi) = \sum_{t=1}^{T} -\log|h_t^2(\xi)| - \sum_{t=1}^{T} u_t^2/h_t^2 \tag{11}$$

Assuming normality of the forecast errors and subject to regularity conditions maximization of (11) will produce maximum likelihood estimates with the usual properties (Goodhart et al., 1994).[2]

5. Empirical results

In Table 2, the estimates of the single-factor CAPM for all the countries with the exception of Canada, indicate that (a) systematic risk measured by betas, on government bond indices increase with maturity; (b) the value of beta is similar across these countries for short-term bond indices; (c) the betas are statistically significant at 5% level of significance; (d) for the UK and the USA 10-year bond indices the systematic risk as measured by beta, does not increase as expected, indicating an anomaly according to the theory, where high beta assets should in equilibrium receive higher risk and higher premium.

In this study, in order to detect a well defined econometric model to test our EH, we start by estimating a simple regression of EHPY on the spread of the previous period as in Eq. (5) in Section 4. We perform tests for serial correlation and non-linearity. If the test statistics are significant, we adjust for the effects by including the lagged EHPY value and the GARCH part. The tests we apply are the Durbin–Watson test for serial correlation, the Box–Pierce test on the squared residuals, an LM test for ARCH(12) (Engle, 1982), in order to investigate the presence of nonlinearities. Table 3 gives the results of these tests. For the D–W test we give the test statistic while for the Box–Pierce test on the squared residuals and the LM test we give the corresponding P-values.

For values within the 1.76 and 2.24 interval, we do not reject the hypothesis of uncorrelated errors. For values in the interval from 1.71 to 1.76 and respectively from 2.24 to 2.29, we cannot draw conclusions. Outside these boundaries, we reject the null hypothesis. With an asterisk we mark the cases where the null hypotheses are rejected. If a GARCH-M part is included, all diagnostic tests are applied again, and particular emphasis is given on the closeness to normality of the standardized residuals.

[2] The likelihood function was maximized by using the SIMPLEX algorithm in terms of improving the initial starting values and the BHHH algorithm to achieve both convergence and consistency in the estimates of the asymptotic standard errors (Berndt et al., 1974).

Table 2
Single factor CAPM model for holding period yields (HPY)

Bond maturity	Parameters	USA	UK	Germany	France	Canada
2 years	*a*	1.17 (0.05)*	0.61 (0.48)	0.66 (0.19)	1.07 (0.05)*	0.9695 (0.31)
	b	0.06 (0.00)*	0.075 (0.00)*	0.073 (0.00)*	0.10 (0.00)*	0.00 (0.77)
	Adj. R^2	0.07	0.07	0.11	0.16	0.001
3 years	*a*	1.73 (0.05)*	1.31 (0.25)	1.24 (0.07)**	2.00 (0.01)*	1.94 (0.13)
	b	0.08 (0.00)*	0.09 (0.00)*	0.11 (0.00)*	0.15 (0.00)*	0.007 (0.83)
	Adj. R^2	0.08	0.06	0.144	0.17	0.0007
5 years	*a*	2.36 (0.08)**	2.07 (0.21)	1.61 (0.099)**	2.98 (0.01)*	2.374 (0.19)
	b	0.12 (0.00)*	0.14 (0.00)*	0.163 (0.00)*	0.22 (0.00)*	0.007 (0.88)
	Adj. R^2	0.07	0.06	0.14	0.18	0.0002
7 years	*a*	2.87 (0.09)**	2.96 (0.14)	3.22 (0.00)*	4.25 (0.00)*	3.81 (0.07)**
	b	0.14 (0.00)*	0.14 (0.02)*	0.22 (0.00)*	0.24 (0.00)*	0.012 (0.83)
	Adj. R^2	0.064	0.044	0.16	0.13	0.0004
10 years	*a*	2.961 (0.16)	3.27 (0.19)	3.13 (0.04)*	4.81 (0.01)*	3.774 (0.13)
	b	0.114 (0.05)*	0.122 (0.09)**	0.24 (0.00)*	0.29 (0.00)*	0.013 (0.84)
	Adj. R^2	0.03	0.023	0.13	0.12	0.0003

Note: The *P*-values of the *t*-test are in parentheses.
* Statistically significant at 5% level of significance.
** Statistically significant at 10% level of significance

Table 3
Diagnostic tests for residuals of $\text{EHPY}_t = c + a\, S_{t-1} + u_t$

Bond maturity	Statistics	USA	UK	Germany	France	Canada
2 years	DW	1.68*	1.65*	1.42*	1.67*	1.99
	Qsq(4)	0.73	0.09**	0.45	0.72	0.00*
	ARCH(12)	0.87	0.21	0.65	0.7	0.81
3 years	DW	1.71*	1.69*	1.46*	1.65*	2.06
	Qsq(4)	0.30	0.28	0.96	0.95	0.47
	ARCH(12)	0.69	0.13	0.61	0.90	0.44
5 years	DW	1.74*	1.72*	1.63*	1.62*	1.94
	Qsq(4)	0.21	0.13	0.96	0.95	0.19
	ARCH(12)	0.51	0.01*	0.97	0.99	0.63
7 years	DW	1.82	1.84	1.72*	1.71*	1.84
	Qsq(4)	0.28	0.03*	0.74	0.77	0.01*
	ARCH(12)	0.02*	0.00*	0.97	0.21	0.16
10 years	DW	1.87	1.83	1.96	1.77*	1.90
	Qsq(4)	0.03*	0.48	0.36	0.53	0.00*
	ARCH(12)	0.02*	0.24	0.90	0.35	0.02*

* Statistically significant at 5% level of significance.
** Statistically significant at 10% level of significance.

According to this table, there are indications of conditional heteroskedastic errors in the cases of the UK, USA and Canada for longer maturities bonds. We have evidence of serial correlation in Germany and France in almost all cases, and in the USA and the UK over the short and medium-term bonds.

In Table 4 we present the estimation results for models (7)–(9) for all the markets of our analysis. The first column contains the parameters following the notation of relationships (7)–(9). Results for each maturity band are presented in two columns on the upper part of the table. The first column contains the point estimate and the second the corresponding P-value of a test of significance. The lower part of the table gives the P-values of some diagnostic statistics. The first and second rows give the results of the Box–Pierce test on the residuals and the squared residuals. If the mean equation is correctly specified all Q statistics should be insignificant. Similarly, if the variance equation is correctly specified all Q-statistics of the squared residuals should be insignificant. The third row gives the result of the Bera–Jarque test for normality of errors. Lastly, the fourth gives the P-value of the LM test for an ARCH(1) specification of the residuals. The ARCH(1) test shows if the standardized residuals exhibit additional ARCH effects.

Where an underscore (–) appears, no such parameter is included in the model for the particular maturity. There is a suggestion that a GARCH(1,1) is satisfactory for most cases (Bera and Higgins, 1993). We start with a simple model of the above type and when appropriate statistics suggest so, we end up using a different more complicated one.

Finally, in a statistical model, which is an extension of (7) and fulfills all diagnostic tests, we test the Expectations Hypothesis for the coefficient a in Eq. (7) to be equal to zero. Looking at the first row of the table, we conclude that for the USA, Germany bonds (except 5-year bonds) and the French, Canadian bonds (except 2-year bonds) the coefficient a in all cases is statistically significant. Clearly, we can reject the EH for these cases according to this criterion. We further comment, that in the majority of the cases the estimates have negative signs, indicating negative correlation between EHPY and the spread variable. A negative value of α implies that when current bond yields are high relative to short-term rates subsequent holding period returns tend to be low. This result provides evidence against a type of mean reversion in bond yields.

Next, we examine the EH with respect to the second criterion, that of uncorrelated errors. If the term premium varies through time it is possibly be serially correlated. This lead us to reject the EH for the short-term bonds (2-, 3-year bonds) in the USA and the UK, because b is statistically significant, something which agrees with the D–W statistic in Table 3. By proceeding with the same way for the rest of the countries, we conclude that the null hypothesis of serially uncorrelated errors shall be rejected for the German and French 2-, 3-, 5- and 7-year cases. This result indicates a consistent pattern in excess yields of the government bonds in these developed European countries.

For Germany and France over the maturity bands, we conclude that the variance is constant over time. For the USA market, we have constant variance for the 2–5-year bonds, and a low order GARCH accommodate for the rest of the series. Almost the same applies for Canada over 2–7-year bonds. For the UK we have detected heteroskedasticity in 5- and 7-year maturity bonds.

One of the common characteristics across USA and UK bond markets is the significance of the coefficient of the risk factor (δ) for the 5- and 7-year bonds. Another important

Table 4
Estimates in EHPY and its volatility

	2 years		3 years		5 years		7 years		10 years	
	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value
US bond market.										
Regression										
S_{t-1}	−2.05	0.05*	−2.61	0.04*	−2.19	0.18	−1.48	0.10	−1.50	0.03*
EHPY_{t-1}	0.17	0.02*	0.15	0.04*	0.14	0.10	–	–	–	–
GARCH	–	–	–	–	–	–	−0.97	0.03*	−0.11	0.18
Constant	1.99	0.01*	3.35	0.01*	4.38	0.04*	23.68	0.02*	8.08	0.29
GARCH										
α_0	–	–	–	–	–	–	308.61	0.00*	306.50	0.04*
α_1	–	–	–	–	–	–	0.22	0.02*	0.03	0.10
α_2	–	–	–	–	–	–	0.10	0.05*	–	–
β_1	–	–	–	–	–	–	–	–	0.95	0.00*
Diagnostics										
Q(12)		0.97		0.82		0.90		0.75		0.97
Qsq(12)		0.69		0.60		0.49		0.24		0.46
J-B		0.99		0.94		0.85		0.64		0.61
ARCH(1)								0.92		0.66
UK bond market										
Regression										
S_{t-1}	0.80	0.20	0.99	0.17	1.07	0.30	1.06	0.35	0.85	0.45
EHPY_{t-1}	0.17	0.02*	0.15	0.05*	0.087	0.37	–	–	–	–
GARCH	–	–	–	–	1.08	0.00*	1.04	0.01*	–	–
Constant	0.26	0.74	0.82	0.45	−21.64	0.00*	−24.55	0.01*	2.69	0.24
GARCH										
α_0	–	–	–	–	441.61	0.00*	890.27	0.00*	–	–
α_1	–	–	–	–	0.18	0.03*	0.08	0.05*	–	–
α_2	–	–	–	–	–	–	–	–	–	–
β_1	–	–	–	–	0.60	0.00*	0.77	0.00*	–	–
Diagnostics										
Q(12)		0.36		0.34		0.29		0.35		0.35
Qsq(12)		0.31		0.29		0.33		0.11		0.22
J-B		0.34		0.31		0.30		0.28		0.26
ARCH(1)		–		–		0.81		0.78		–
German bond market										
Regression										
S_{t-1}	−0.24	0.068**	1.19	0.08*	1.20	0.12	1.34	0.015*	−0.11	0.85
EHPY_{t-1}	0.26	0.04*	0.25	0.00*	0.17	0.02*	−0.22	0.07**	–	–
GARCH	–	–	–	–	–	–	–	–	–	–
Constant	0.58	0.20	0.40	0.55	0.44	0.67	2.32	0.006*	3.93	0.00*
GARCH										
α_0	–	–	–	–	–	–	–	–	–	–
α_1	–	–	–	–	–	–	–	–	–	–
α_2	–	–	–	–	–	–	–	–	–	–
β_1	–	–	–	–	–	–	–	–	–	–
Diagnostics										
Q(12)		0.55		0.25		0.37		0.97		0.97
Qsq(12)		0.53		0.44		0.88		0.99		0.96
J-B		0.11		0.53		0.15		0.63		0.53
ARCH(1)										

Table 4 (*Continued*)

	2 years		3 years		5 years		7 years		10 years	
	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value	Estimate	*P*-value
French bond market										
Regression										
S_{t-1}	−0.22	0.36	−0.37	0.017*	−0.42	0.01*	−0.77	0.02*	−1.18	0.01*
$EHPY_{t-1}$	0.15	0.05*	0.17	0.01*	0.18	0.02*	0.12	0.10	–	–
GARCH	–	–	–	–	–	–	–	–	–	–
Constant	0.40	0.44	0.73	0.32	1.29	0.24	2.98	0.05*	3.39	0.049*
GARCH										
α_0	–	–	–	–	–	–	–	–	–	–
α_1	–	–	–	–	–	–	–	–	–	–
α_2	–	–	–	–	–	–	–	–	–	–
β_1	–	–	–	–	–	–	–	–	–	–
Diagnostics										
Q(12)		0.84		0.34		0.53		0.66		0.26
Qsq(12)		0.42		0.47		0.97		0.80		0.17
J-B		0.36		0.79		0.69		0.36		0.33
ARCH(1)										
Canadian bond market										
Regression										
S_{t-1}	0.59	0.21	0.6032	0.06**	0.461	0.032*	0.817	0.08**	1.682	0.017*
$EHPY_{t-1}$	–	–	–	–	–	–	–	–	–	–
GARCH	–	–	–	–	–	–	–	–	−0.0314	0.05*
Constant	2.44	0.00*	2.033	0.00*	2.95	0.00*	2.95	0.00*	23.62	0.5
GARCH										
a_0	–	–	–	–	–	–	–	–	225.29	0.00*
a_1	–	–	–	–	–	–	–	–	0.0432	0.016*
b_1	–	–	–	–	–	–	–	–	0.647	0.037*
Diagnostics										
Q(12)		0.87		0.96		0.97		0.97		0.698
Qsq(12)		0.82		0.56		0.54		0.66		0.02
J-B		0.44		0.72		0.37		0.49		0.11
ARCH(1)										0.96

* Statistically significant at 5% level of significance.
** Statistically significant at 10% level of significance.

characteristic from our findings is that a GARCH(1,1) model is enough to explain the time variation in the term premium in long-term bonds.[3]

Finally, a joint diagnostic test for nonlinear ARCH effects (following Engle and Ng, 1993) is applied. The test relies on the following regression:

$$\varepsilon_{it}^2 = a + bS_{t-1}^- + b_2 S_{t-1}^- \varepsilon_{i,t-1} + b_3(1 - S_{t-1}^-)\varepsilon_{it-1} + cZ_i + e_i$$

where ε_{it} is the standardized residuals, S_{t-1} the dummy variable taking the value of 1 if ε_{it-1} is negative and 0 otherwise and Z a vector of explanatory variables. The test statistic for

[3] The likelihood ratio (LR) test for GARCH(1,1) against a GARCH(1,2) for h_t was more significant in all cases considered.

this test is the LM statistic $nR^2 \sim \chi^2_\alpha$ with 3 degrees of freedom. None of the test statistics is significant at the usual significance levels implying that the used volatility model is not rejected. Specifically the test values fount were 3.02, 2.72 and 2.56 for the UK (5 years) and the USA (5 and 7 years) respectively.

6. Conclusions

This paper examines the expectations theory of the term structure of interest rates, by testing, whether the term premium of monthly excess holding period yields is constant, in the government bond indices of five different countries. It assess if there are patterns in risk premia which are common across these markets.

Assuming a constant time term premium there are significant evidence against the expectations theory based on the criteria of the yield spread effect. The spread between long and short rates appears to forecast in most countries (exception the UK) the excess holding period yield on government bonds. As it was seen in Table 3, the effects of the lagged error variable in some countries provide evidence against efficiency conditions.

When attempt is made to explain term premium using the CAPM, where systematic risk is defined as the covariance of return with the return on a “market” bond portfolio, we find that risk premium rises with maturity. The value of this systematic risk is similar for the same maturity band across countries. However, all the above does not take into account the time varying term premium. The failure of the expectations theory to international government bond indices data, in some countries, could be explained by the existence of this time varying premium. If this is considered, common patterns in risk premia across some countries may appear.

There is some evidence of inefficiencies across mid-term bonds in the USA (2, 3, 5 years), Germany and France (2, 3, 5, 7 years), where lagged error variable is statistically significant. Evidence of time varying term premium is found for longer maturity bonds, in the USA (7, 10 years), the UK (5, 7 years) and Canada (10 years) as expected from high level of bond yields and high volatility over the period tested. The finding that long-term bond premia are positive correlated across markets suggests that the factors responsible for time variation in term premia, many times in our days have an international character. And probably this variation reflects changes in risk premia that are easily announced with the network technologies in all over the world. However, the risk factor is only in few cases statistically significant for the explanation of EHPY in a GARCH-M model.

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