# PART IV

# The Predictability of Bond Returns

# 9

# International Bond Risk Premia

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## 9.1 INTRODUCTION

The endeavor to understand bond returns and the term structure of interest rates has generated an extensive literature, ranging from papers on return predictability and affine term structure models to theoretical contributions in the form of equilibrium models. While most of the empirical literature focuses on U.S. data, a large body of work applies an international perspective. This chapter reviews the relevant literature while also providing empirical evidence on international bond risk premia and the link to the macroeconomy.

The expectations hypothesis of interest rates is commonly expressed in three forms: long-term interest rates equal average expected future short-term rates plus a risk premium, forward rates equal expected future interest rates plus a risk premium, and the expected one-period return on bonds equals the one-period interest rate plus a risk premium. All formulations of the expectations hypothesis restrict risk premia on government bonds to being zero or at least constant over time.[1] As such, the expectations hypothesis leaves no room for time-varying risk premia or compensation for risk.

An extensive literature has empirically examined whether risk premia on government bonds are indeed constant over time. The standard way of testing the expectations hypothesis is by running predictability regressions in which future realized bond returns in excess of the risk-free rate are regressed on predictive variables such as the spread between yields on long- and short-term bonds or the spread between forward and spot rates. A statistically significant slope coefficient in such a regression implies that future bond excess returns are predictable or, equivalently, that bond risk premia are time varying. Hence, excess return predictability and time-varying risk premia are two sides of the same coin, both violating the expectations hypothesis.

For academic researchers, the interest in predictability lies mainly in understanding why investors' required risk compensation should vary over time. Such time variation may arise through either rational or irrational channels. Asset pricing theory suggests that rational variation in risk premia and therefore predictability may arise from either time-varying economic risks or time-varying risk aversion among investors. This suggests that future excess returns can be predictable even if markets are efficient. For example, academic evidence suggests that investors require higher risk premia on nominal bonds during economic

[1]The pure expectations hypothesis sets risk premia to zero. Cox, Ingersoll, and Ross (1981) demonstrate that various formulations of the pure expectations hypothesis are inconsistent with one another when interest rates are random due to Jensen's inequality. Campbell (1986) demonstrates that this is generally not the case for the less restrictive form of the expectations hypothesis in which risk premia are nonzero but constant, which is what the empirical literature is mostly concerned with. We will consider alternative ways of empirically evaluating the expectations hypothesis.

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recessions, when economic uncertainty and investors' risk aversion arguably increase. However, rather than reflecting rational variation in risk premia, predictability may also arise from market inefficiencies and irrational investor behavior. For example, empirical evidence in behavioral finance suggests that predictability of bond returns may reflect expectational errors made by investors whose expectations deviate from rational expectations.

For investors, predictability of returns is naturally attractive from an asset management perspective. Rather than implying that future returns are known with certainty, predictability of returns implies that a statistically significant return-forecasting variable can turn the odds slightly in favor of the investor compared with making a 50:50 bet. Furthermore, diversifying across several predictive variables that are less than perfectly correlated is likely to yield even greater benefits in terms of risk-adjusted returns. In practical implementation, the distinction between in-sample and out-of-sample predictability is particularly important. While evidence suggests that several variables do predict returns in sample, significant out-of-sample predictability is typically harder to come by.

The empirical analysis presented in this chapter builds on Dahlquist and Hasseltoft (2013), who found time-varying risk premia across four international bond markets. Both local and global factors were found to drive international bond risk premia, the global factor being closely linked to U.S. bond risk premia and to international business cycles. In this chapter, we expand the cross section of countries examined and refine the analysis of bond risk premia and economic growth.

We document that the standard Fama and Bliss (1987, Fama–Bliss (FB)) regressions fail to predict excess returns across a broad set of countries; in contrast, Cochrane and Piazzesi (2005, Cochrane–Piazzesi (CP)) regressions seem to capture substantial predictability. Furthermore, we construct a global return-forecasting factor that, jointly with local factors, predicts excess returns across countries. International bond risk premia seem to be driven by both local and global factors, the global factor being closely linked to U.S. bond risk premia. The fact that global and U.S. risk premia are closely related is consistent with the literature on real integration in which, for example, Kose, Otrok, and Whiteman (2003) identify a common factor in world business cycles that is highly correlated with U.S. output growth.

Consistent with asset pricing theory, we find that a rise in global bond risk premia is contemporaneously negatively correlated with global economic growth, a 1% rise in global risk premia being associated with a drop of 0.5–1% in annualized economic growth in three large economic areas, that is, the European Monetary Union (EMU), OECD, and United States. In addition, we find that the global factor positively predicts economic growth, a one-percentage-point increase in global bond risk premia being associated with a 0.5–1.3% annualized increase in future economic growth. Hence, a rise in global bond risk premia is contemporaneously associated with bad economic times but signals improved times ahead.

Though bond risk premia are positive on average, we find that bond risk premia turn negative during some periods. Negative bond risk premia suggest that nominal bonds act as hedging assets rather than risky assets. To further analyze the riskiness of bonds, we calculate rolling betas of nominal bonds with equity and find that bond betas turned negative around year 2000 in all countries, bottoming around the time of the financial crisis. These findings are consistent with what Campbell, Sunderam, and Viceira (2013) document for U.S. data. In addition, we find that movements in bond betas are highly correlated across countries, again suggesting that global and systematic factors may drive changes in the riskiness of nominal bonds.

We also analyze whether global bond risk premia display any state dependency with respect to economic recessions and expansions. Evidence from equity markets suggests that traditional return predictors such as price–dividend ratios mainly predict returns in bad economic times (e.g., Henkel, Martin, and Nardari, 2011). We test this by predicting international bond excess returns using our global factor, conditioning on the state of the economy. We define recessions outside the United States using data from the Economic Cycle Research Institute while using NBER-dated recessions for the United States. Unlike in equity markets, we find no evidence of state dependency in bond markets, meaning that our global factor predicts returns across business cycles.

The chapter proceeds as follows. Section 9.2 reviews the literature on bond return predictability and term structure modeling. Section 9.3 presents our notation and data on international bond markets. Section 9.4 presents and discusses the unconditional properties of international bond risk premia. Section 9.5 examines conditional bond risk premia by running various predictability regressions. Section 9.6 interprets the findings and links international bond risk premia to global economic growth and business cycles. Section 9.7 concludes the chapter.

## 9.2 LITERATURE REVIEW

As far back as the 1960s, Kessel (1965) studied whether the expectations hypothesis of interest rates could explain the level and cyclical behavior of interest rates, and Van Horne (1965) studied the risk premium component embedded in forward rates. More recently, the expectations hypothesis has typically been tested by running predictability regressions in which excess bond returns are regressed on one or several predictive variables. The seminal contributions of FB (1987) and Campbell and Shiller (1991) use the spread between forward and spot rates and the slope of the spot yield curve, respectively, as forecasting variables. They find significant evidence of time-varying risk premia in U.S. bond markets, soundly rejecting the expectations hypothesis. CP (2005) identify a factor in the form of a linear combination of U.S. forward rates that predicts excess returns with roughly double the explanatory power of traditional regressions. Interestingly, the CP factor is largely unrelated to the traditional level, slope, and curvature factors that drive variation in interest rate levels (e.g., Litterman and Scheinkman, 1991). This raises the question of how an interest rate factor can contain important information about risk premia while having a small effect on the cross section of yields. Duffee (2011) analyzes this question and documents a hidden factor in yields that has offsetting effects on risk premia and expectations of future short-term interest rates, leaving the level of yields largely unaffected. Hence, it seems that certain information about future bond returns is not easily summarized by current interest rates. This possibility is supported by Ludvigson and Ng (2009) and Cooper and Priestley (2009), who identify U.S. macro factors that add incremental predictive power beyond the yield-based factors.

In contrast to the fairly robust U.S. evidence against the expectations hypothesis, the international evidence is mixed. For example, Ilmanen (1995) constructs a proxy for risk aversion in the form of the inverse of the real stock market level and finds that it significantly predicts international bond returns, while Hardouvelis (1994) and Bekaert and Hodrick (2001) both find it difficult to reject the expectations hypothesis based on international data. In addition, Dahlquist and Hasseltoft (2013) find weak international evidence against the expectations hypothesis when running FB regressions. However, they document substantial international predictive power when running local and global CP (GCP) regressions, suggesting that there is significant time variation in international bond risk premia after all.

A different stream of literature centers on affine term structure models. The list of relevant contributions in this stream is too long to cite in full but includes Duffie and Kan (1996), Dai and Singleton (2000, 2002), Duffee (2002), Ang and Piazzesi (2003), CP (2008), and Joslin, Priebsch, and Singleton (2014). The affine models typically use a set of latent variables, such as the level, slope, and curvature factors of interest rates, to price bonds, sometimes complementing them with macroeconomic variables. Focusing on international yields, Diebold, Li, and Yue (2008) build on Nelson and Siegel (1987) and document global yield curve factors that appear to be linked to global macroeconomic factors such as inflation and real activity. Wright (2011) decomposes a cross section of international yields into term premia and expected future short-term rates. He finds that term premia have declined globally, in line with the level of yields, since the 1990s and links this to a drop in inflation uncertainty and changes in monetary policy. Bauer, Rudebusch, and Wu (2014) address Wright's (2011) estimation and demonstrate that correcting for small-sample bias in the estimation of affine term structure models changes the conclusion regarding international term premia. Rather than mirroring the long-term downward trend in interest rates, adjusting for small-sample bias renders international term premia countercyclical and consequently more consistent with asset pricing theory.

The endeavor to understand movements in interest rates and bond risk premia has produced a large number of equilibrium models, with early contributions such as Cox, Ingersoll, and Ross (1985), Dunn and Singleton (1986), and Backus, Gregory, and Zin (1989). Building on Campbell and Cochrane's (1999) consumption-based framework of habit formation in which investors' risk aversion rises in bad economic times but the amount of economic risk is constant, Brandt and Wang (2003), Wachter (2006), and Buraschi and Jiltsov (2007) all find that habit formation can generate time-varying bond risk premia. In contrast, Hasseltoft (2012) and Bansal and Shaliastovich (2013) build on the long-run risk framework of Bansal and Yaron (2004), in which investors' risk aversion is constant but the amount of economic risk varies, and demonstrate that time-varying macroeconomic volatility can generate bond return predictability similar to that observed in the data. Another area of relevant theoretical models is the disaster risk models. Barro (2006) builds on Rietz (1988) and demonstrates that the presence of low-probability but large economic disasters, such as the Great Depression, can account for the equity risk premium and the volatility of stock returns. Gabaix (2012) argues that a disaster risk model in which the severity of disasters varies over time can explain several asset pricing puzzles, including the predictability of bond excess returns. In this model, inflation is assumed to increase in periods of economic disaster, leading to low bond returns in bad times and therefore making nominal bonds risky assets. Time variation in the severity of disaster risk allows for time-varying risk premia and therefore predictable excess returns.

Moreover, work has been done to explain interest rates and bond returns using theories of portfolio rebalancing and preferred habitats (see Culbertson, 1957; Tobin, 1958; and Modigliani and Sutch, 1966, for early contributions). The portfolio rebalancing theory suggests that changes in the supply of government bonds affect interest rates. For example, increasing the supply increases the interest rate risk for investors, who consequently demand higher expected returns, leading to higher interest rates. The preferred-habitat theory moves away from the common assumption of a representative agent and models a segmented market with various investor clienteles along the yield curve, each having its own demand effect on the term structure. Krishnamurthy and Vissing-Jorgensen (2011, 2012) note that changes in the supply of government debt affect the level and slope of the yield curve. Greenwood and Vayanos (2010) provide empirical support for the preferred-habitat view in the forms of the U.S. Treasury buyback program of 2000–2001 and the U.K. pension reform of 2004. Furthermore, Greenwood and Vayanos (2014) find that an increased supply of long-term government debt steepens the yield curve and predicts excess returns on long-term bonds.

Finally, there are several extensive reviews of the term structure of interest rates. Shiller (1990) reviews fundamental concepts and theories of interest rates. Dai and Singleton (2003) and Piazzesi (2010) relate fixed-income pricing to term structure models, while Gürkaynak and Wright (2012) and Duffee (2013) review the link between the term structure of interest rates and the macroeconomy. While the empirical analysis in these reviews centers on U.S. data, less is known about international term structures, which are the focus of this chapter.

## 9.3 NOTATION AND INTERNATIONAL BOND MARKET DATA

### 9.3.1 Notation

We use standard notation throughout the chapter. The annual return on an $n$-period bond is defined as $r^n_{t+1} = p^{n-1}_{t+1} - p^n_t$, where $p$ denotes the log bond price and the maturity, $n$, and time, $t$, are defined in terms of years. (We will consider data sampled on annual as well as monthly bases. For convenience, we keep the notation for time in years even when we are sampling data on a monthly basis.) The log yield of a bond is computed as $y^n_t = -p^n_t/n$. The annual return in excess of the 1-year yield is defined as $rx^n_{t+1} = r^n_{t+1} - y^1_t$. We define the 1-year forward rate between periods $n-1$ and $n$ as the differential in log bond prices, $f^n_t = p^{n-1}_t - p^n_t$. We consider returns for several countries but typically suppress the country indexation, $c$, of variables unless we find it is necessary to avoid confusion.

### 9.3.2 International Bond Market Data

We consider four datasets when analyzing unconditional and conditional bond risk premia. First, we use the Dimson, Marsh, and Staunton (DMS) dataset, obtained from Morningstar and used by Dimson, Marsh, and Staunton (2002, 2014). This dataset contains annual bond excess returns on long-term bonds, covering 21 countries over the 1900–2013 period.[2] Taking a long-term perspective is interesting because the international bond market data used in academic studies typically start in the 1970s or later. Second, we use an updated version of the dataset used in Dahlquist and Hasseltoft (2013), which allows us to run predictability regressions starting in the 1970s. This dataset covers four countries, that is, Germany, Switzerland, the United Kingdom, and the United States, and covers the period from January 1975 to December 2013. All yields are based on zero-coupon bonds for maturities of 1–5 years: German yields are obtained from the Bundesbank; Swiss yields are obtained from Jorion and Mishkin (1991) until December 1987, after which yields from the Swiss National Bank are used; U.K. yields are obtained from the Bank of England; and U.S. yields are obtained from the Fama–Bliss discount bond file from the Center for Research in Security Prices.

Third, we consider a shorter-term dataset that contains a richer cross section of countries. We obtain yields on derived zero-coupon bonds for ten countries from Datastream covering the period from December 1999 to December 2013.[3] Fourth, we consider a dataset from Datastream consisting of returns on Citigroup bond indexes for 19 countries. For each country, five indexes cover five maturity buckets, that is, 1–3 years, 3–5 years, 5–7 years, 7–10 years, and 10+ years. In addition, we use several complementary data in our analysis. We collect yields on long-term bonds from Global Financial Data starting in the early 1900s and spanning the same set of countries as in DMS. We obtain annual PPP-adjusted gross domestic product (GDP) data from Datastream, which are used to construct a global return-forecasting factor, and we also consider data on industrial production and economic leading indicators from the OECD. Finally, we relate bond and equity returns as described later.

## 9.4 UNCONDITIONAL RISK PREMIA

We start by describing the unconditional properties of bond excess returns around the world. We do so by first applying a broad long-term approach using the DMS dataset, providing us with an understanding of excess returns on long-term bonds since 1900. We then apply a more refined and detailed approach to how excess returns have evolved across countries and maturities since the 1970s.

### 9.4.1 A Long-Term Perspective

Before reporting and discussing long-term risk premia, it is helpful first to consider some caveats regarding the data. Some countries in particular have undergone extreme episodes since 1900. The German hyperinflation from 1922 to 1923 caused investors in German bonds and bills to lose their entire investment. Austria also experienced extreme volatility in the 1920s and 1930s. Furthermore, international bond markets have undergone periods of turbulence due to two World Wars, several instances of civil war, and regulatory changes or financial repression over the years. These events all represent periods in which yields were not subject to regular market forces. Two examples of many in terms of regulations include those of the United States, where

[2]The DMS bond indexes are equal-weighted portfolios of long-term bonds with a desired maturity of 20 years. If no such bonds are available, either perpetuals or shorter-maturity bonds are used.
[3]The zero-coupon yields provided by Datastream are derived from interbank instruments such as futures and swaps and therefore contain a credit risk component; however, using these yields provides us with a richer cross section of markets. Furthermore, yields from Bundesbank, the Swiss National Bank, and Datastream use annual compounding, while the remaining yields use continuous compounding. We accordingly convert annually compounded yields to continuously compounded yields.

**TABLE 9.1 Means and Standard Deviations of Bond Excess Returns**

| | Mean | Standard deviation | Mean I | Mean II | Mean III |
|---|---|---|---|---|---|
| Australia | 0.83 | 10.16 | 1.51 | −0.34 | 1.34 |
| Austria | 4.35 | 43.02 | 11.11 | −0.98 | 2.91 |
| Belgium | 0.48 | 9.18 | 0.02 | −1.19 | 2.60 |
| Canada | 0.59 | 8.21 | −0.93 | 0.03 | 2.69 |
| Denmark | 0.93 | 8.51 | −0.30 | −0.33 | 3.43 |
| Finland | 0.55 | 5.56 | −0.47 | 0.28 | 1.85 |
| France | 2.81 | 7.51 | −0.43 | 3.14 | 5.72 |
| Germany | 0.77 | 10.29 | −1.66 | 0.90 | 2.94 |
| Ireland | 0.73 | 11.38 | −0.21 | −1.39 | 3.78 |
| Italy | 2.19 | 8.10 | 1.13 | 1.79 | 3.64 |
| Japan | 0.91 | 13.55 | −0.60 | −0.57 | 3.89 |
| Netherlands | 0.89 | 7.47 | 0.05 | −0.15 | 2.78 |
| New Zealand | 0.37 | 6.88 | 0.28 | 0.42 | 0.42 |
| Norway | 0.62 | 8.17 | −0.59 | 1.37 | 1.08 |
| Portugal | 1.73 | 17.47 | 8.52 | −3.85 | 0.53 |
| South Africa | 0.86 | 7.85 | 1.56 | 0.12 | 0.90 |
| Spain | 1.13 | 9.27 | 0.34 | 1.36 | 1.71 |
| Sweden | 0.65 | 8.73 | 0.76 | −1.14 | 2.34 |
| Switzerland | 1.41 | 5.94 | 0.24 | 1.36 | 2.62 |
| United Kingdom | 0.50 | 10.25 | −0.30 | −1.26 | 3.06 |
| United States | 0.98 | 8.22 | −0.51 | 0.08 | 3.36 |

This table presents means and standard deviations of annual log returns on bonds in excess of log returns on bills. The sample period is 1900–2013, yielding 114 annual observations. The right-hand side of the table presents means for three subperiods: I. 1900–1937, II. 1938–1975, and III. 1976–2013. Each subperiod has 38 annual observations. The statistics for Germany exclude the hyperinflationary period 1922–23. Means and standard deviations are expressed in percentage per year. Data sources are given in Section 9.3.2.

long-term rates were capped at 2.5% throughout the 1940s, and of Japan, where bond yields experienced extended periods of tight regulation before the start of financial liberalization in 1975.[4]

Table 9.1 reports the means and volatilities of annual realized bond excess returns for the DMS dataset covering the 1900–2013 period. In addition to the means and standard deviations of the entire sample period, we also report the means for three subsamples: 1900–1937, 1938–1975, and 1976–2013. The table shows that the averages for the entire sample period were positive in all countries, albeit subject to high volatility, making standard errors large (untabulated). Austria experienced the highest annual volatility at 43%, while Finland experienced the lowest at approximately 5%. The last three columns in the table refer to the three subsamples. Historical bond risk premia have increased over time, with most countries exhibiting realized risk premia of 2% or more since 1976. The level of risk premia for a particular subperiod is related to the initial yield of that period, as interest rates tend to exhibit mean reversion over time. This is evident from Figure 9.1, in which the solid line plots the average long-term nominal yield across 21 countries, using data from Global Financial Data. The average nominal yield peaks in 1981, after which there is a continued downward drift with consequently high realized returns. A principal component analysis of the international long-term nominal yields reveals a first factor that explains 76% of the variation, suggesting significant commonality across countries. This commonality has increased over time, with the first factor explaining 50% of the variation in the first half of the sample period and 78% in the second half.

Nominal interest rates can be decomposed into three components: the real interest rate, expected inflation over the maturity of the bond, and an inflation risk premium. The real interest rate is an interesting and useful concept but is difficult to measure accurately for several reasons, particularly on an ex ante basis. Reliable statistical projections of long-term inflation are difficult to compute, and reliable surveys of long-term expected inflation are scarce. In addition, the inflation risk premium is difficult to measure accurately, and the result depends greatly on the methods used. A direct way of measuring the real interest rate is to use yields on real government bonds, that is, inflation-protected bonds or index-linked bonds. A drawback of this approach is that the available time series are short as, for example, real bonds began to be issued in 1981 in the United Kingdom and 1997 in the United States. The principal and coupon payments on a real bond are typically indexed to a measure of consumer price inflation, protecting the investor against inflation risk. Campbell, Shiller, and Viceira (2009) describe the history of real bonds

[4]See Homer and Sylla (2005) for a long-term review of international bond markets.

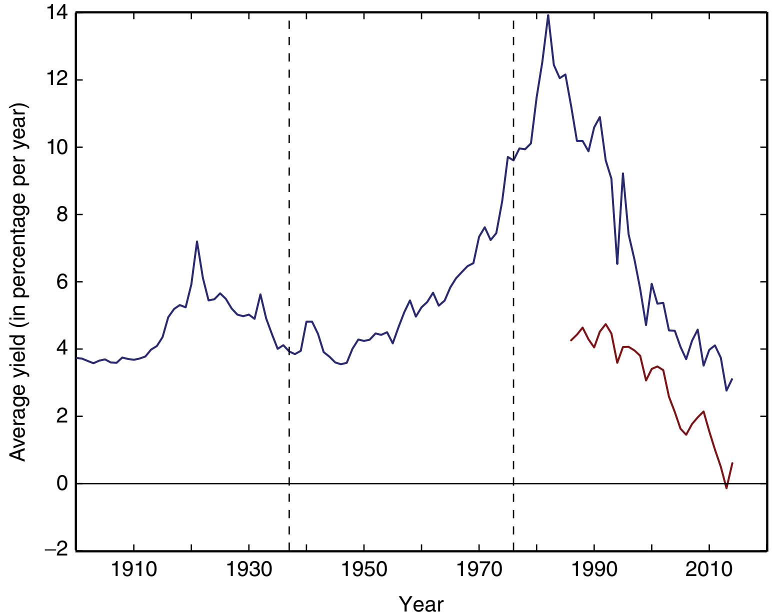


**Figure 9.1** Average long-term international yields. This figure shows global nominal and real interest rates. The black line represents the equal-weighted average of nominal 10-year government yields across 21 countries, excluding Germany for the hyperinflationary period 1922–23. The grey line represents the equal-weighted average of 10-year real interest rates across G7 countries, as computed by King and Low (2014). The dashed vertical lines indicate the subperiods considered in Table 9.1. Source: Global Financial Data and King and Low (2014).

in the United States and United Kingdom and highlight various institutional factors, such as market segmentation and liquidity, that affect yields on real bonds. Pflueger and Viceira (2011) also study nominal and real bonds in the United States and United Kingdom and find the expectations hypothesis to be violated in both markets, implying that excess returns on real bonds are also predictable. Related to the international focus of this chapter is King and Low (2014), who describe the market for real bonds across a number of countries and construct a "world" real interest rate comprising 10-year real yields from the G7 countries. Their estimate of the global real rate starts in 1985 and is plotted as the dotted line in Figure 9.1. The figure shows that both real and nominal rates have been in a downward trend since the 1980s, though the difference between the two rates, referred to as the break-even inflation rate, has varied over time. The drop in break-even inflation over time suggests that inflation expectations have declined from elevated levels in the 1980s to lower levels today. However, it is noteworthy that the break-even inflation rate reflects not only the expected inflation but also the inflation risk premium and potentially a liquidity premium.

### 9.4.2 More Recent Evidence

The sample with yields since 1975 allows us to make a more detailed examination of historical bond risk premia. Table 9.2 reports the first four moments and the Sharpe ratios of annual bond excess returns for Germany, Switzerland, the United Kingdom, and the United States over the 1976–2013 period. Measured risk premia for maturities of 2–5 years are considered. First, average excess returns and volatility tend to be significantly different from zero across countries and increase with maturity. Second, the skewness of excess returns tends to be insignificantly different from zero, except for the German bonds, which exhibited negative skewness throughout the period. Third, excess returns in all countries display kurtosis close to 3, being in the region of the normal distribution. It is noteworthy that the skewness and kurtosis of bond returns differed from those of equity returns during this period, which in each country display large negative skewness and positive kurtosis (untabulated). Fourth, the Sharpe ratios across countries and maturities lie in the 0.30–0.40 region where, for example, the Sharpe ratio of U.S. bonds declines with maturity. The Sharpe ratios tend to be significantly different from zero and larger than the corresponding Sharpe ratios for equities, which are approximately 0.20 in Switzerland and the United States and close to zero in Germany and the United Kingdom.

In line with Litterman and Scheinkman (1991), the variation in yields can be summarized into three principal components, that is, the level, slope, and curvature factors, with the level factor accounting for more than 97% of the variation within each country. Pooling all international yields reveals a level factor that explains 90% of the variation, indicating the existence of a global-level factor in local term structures. A principal component analysis of the pooled excess bond returns results in a common factor that explains 70% of the variation. These observations relate to Driessen, Melenberg, and Nijman (2003), who examine variations in international bond returns and identify a world-level factor that explains approximately 50% of the variation. Furthermore,

**TABLE 9.2 Summary Statistics for Bond Excess Returns**

| | Maturity | Mean | Standard deviation | Skewness | Kurtosis | Sharpe ratio |
|---|---|---|---|---|---|---|
| Germany | 2 | 0.64 | 1.40 | −0.52 | 3.74 | 0.45 |
| | | (0.20) | (0.13) | (0.30) | (0.82) | (0.17) |
| | 3 | 1.25 | 2.61 | −0.51 | 3.30 | 0.48 |
| | | (0.37) | (0.22) | (0.26) | (0.70) | (0.16) |
| | 4 | 1.77 | 3.66 | −0.56 | 3.26 | 0.48 |
| | | (0.52) | (0.31) | (0.25) | (0.67) | (0.16) |
| | 5 | 2.20 | 4.57 | −0.60 | 3.29 | 0.48 |
| | | (0.64) | (0.39) | (0.24) | (0.65) | (0.17) |
| Switzerland | 2 | 0.40 | 1.73 | 0.22 | 4.23 | 0.23 |
| | | (0.25) | (0.19) | (0.31) | (0.68) | (0.14) |
| | 3 | 0.94 | 2.98 | 0.22 | 3.77 | 0.31 |
| | | (0.42) | (0.31) | (0.28) | (0.61) | (0.14) |
| | 4 | 1.45 | 4.00 | 0.10 | 3.53 | 0.36 |
| | | (0.56) | (0.39) | (0.27) | (0.58) | (0.14) |
| | 5 | 1.81 | 4.85 | −0.02 | 3.38 | 0.37 |
| | | (0.67) | (0.46) | (0.28) | (0.54) | (0.15) |
| United Kingdom | 2 | 0.54 | 1.76 | 0.43 | 4.78 | 0.31 |
| | | (0.23) | (0.19) | (0.37) | (0.95) | (0.13) |
| | 3 | 1.07 | 3.21 | 0.38 | 4.32 | 0.33 |
| | | (0.40) | (0.30) | (0.30) | (0.66) | (0.12) |
| | 4 | 1.51 | 4.54 | 0.35 | 4.35 | 0.33 |
| | | (0.55) | (0.42) | (0.32) | (0.59) | (0.12) |
| | 5 | 1.93 | 5.72 | 0.31 | 4.41 | 0.34 |
| | | (0.69) | (0.54) | (0.33) | (0.58) | (0.12) |
| United States | 2 | 0.73 | 1.81 | −0.19 | 3.76 | 0.40 |
| | | (0.26) | (0.17) | (0.29) | (0.49) | (0.15) |
| | 3 | 1.27 | 3.40 | −0.28 | 3.64 | 0.37 |
| | | (0.48) | (0.32) | (0.27) | (0.45) | (0.15) |
| | 4 | 1.78 | 4.73 | −0.25 | 3.52 | 0.38 |
| | | (0.67) | (0.44) | (0.26) | (0.40) | (0.15) |
| | 5 | 2.04 | 5.85 | −0.26 | 3.45 | 0.35 |
| | | (0.82) | (0.54) | (0.26) | (0.40) | (0.15) |

The table presents summary statistics (i.e., means, standard deviations, skewness, kurtosis, and Sharpe ratios) for 1-year returns on bonds with maturities of 2, 3, 4, and 5 years in excess of the return on 1-year bonds (i.e., $rx_{t+1}^{n}$ with $n = 2$, 3, 4, and 5) for Germany, Switzerland, the United Kingdom, and the United States. The sample period is from January 1976 to December 2013, yielding 456 monthly return observations. Central moments are estimated using Hansen's (1982) generalized method of moments, while the standard errors of the reported statistics are derived using the delta method. The standard errors, accounting for serial correlation up to 12 lags as in Newey and West (1987), are reported within parentheses. Means and standard deviations are expressed in percentage per year. Data sources are given in Section 9.3.2.

Perignon, Smith, and Villa (2007) find that U.S. bond returns share one common factor with German and Japanese bond returns and link this to changes in the level of interest rates. Also related are the findings of Jotikasthira, Le, and Lundblad (2015), who study the high correlation of yield changes across countries and argue that a world inflation factor drives risk compensation for long-term bonds. Moreover, the yields in our sample can be well summarized as all exhibiting upward-sloping yield curves, yield levels being highly persistent, short-term rates being more volatile than long-term rates, and yield levels being highly positively correlated both within and across countries. These summary statistics for yields are not tabulated.

## 9.5 CONDITIONAL RISK PREMIA

Having described unconditional bond risk premia, we next consider conditional bond risk premia. As discussed earlier, an extensive literature has documented time-varying risk premia in international bond markets using a broad set of predictive variables. We focus on regressions that use yield-based predictive variables. Using several datasets, we run FB regressions, CP regressions, and regressions using a global factor constructed as the GDP-weighted average of local CP factors.

Our predictability regressions consider annual excess returns sampled on a monthly basis. Hence, there is a data overlap, inducing serial correlation in error terms. All regression results are therefore reported together with Newey and West (1987) standard errors, which account for conditional heteroskedasticity and serial correlation, with the number of lags set equal to the forecasting horizon in months. We do not consider bootstrapped standard errors in this chapter but instead refer the reader to Dahlquist and Hasseltoft (2013), who ran similar regressions. They computed bootstrapped confidence intervals for the $R^2$ values stemming from the predictive regressions, using the stationary block bootstrap simulation of Politis and Romano (1994) with the optimal block size of Politis and White (2004) and Politis, White, and Patton (2009). They found that incorporating bootstrapped confidence bounds in the analysis did not change the main findings. We do not address further econometric challenges such as small-sample biases due to persistence in yields and measurement errors (see, e.g., Bekaert, Hodrick, and Marshall, 1997, and Bekaert and Hodrick, 2001).

### 9.5.1 Local Predictors of Returns

We start by predicting excess returns using FB regressions, doing so using two datasets, one starting in 1975 and the Datastream sample starting in 1999. Annual excess returns are regressed on the differential between forward and spot rates, as follows:

$$rx_{t+1}^n = a^n + b^n FB_t^n + \epsilon_{t+1}^n, \tag{9.1}$$

where $a^n$ and $b^n$ are parameters, $FB_t^n$ equals the forward–spot spread $f_t^n - y_t^1$, and $\epsilon_{t+1}^n$ is an error term. The first columns in Table 9.3 present the results for the 1975 sample. Consistent with earlier results in the literature, we find statistically significant coefficients for the United States with $R^2$ values ranging from 6% to 12%. Switzerland too has highly statistically significant slope coefficients. However, the results for Germany and the United Kingdom are less encouraging, suggesting that the international support for FB regressions is mixed. This is consistent with earlier findings in the literature that it is more difficult to reject the expectations hypothesis for samples outside the United States. We later consider alternative ways of evaluating the expectations hypothesis, namely, by using predictability regressions as in Campbell and Shiller (1991) and Backus et al. (2001).

The first columns in Table 9.4 present the results for the sample starting in 1999. We report the results for bonds with maturities of 2 and 5 years; the results for bonds with 3- and 4-year maturities are similar and typically lie between those for maturities of 2 and 5 years. These results paint a different picture, as only one country, Japan, is subject to statistically significant slope coefficients across maturities. There could be two reasons for the difference in results: either the sample period of 1999–2013 is too short to permit detection of predictability or the forward–spot spread lost its predictive power during this period. To shed further light on this, we next run CP regressions for the two datasets and compare the outcomes to those of the FB regressions.

We follow CP (2005) and construct local CP factors for each country for the two datasets. A CP factor is constructed by regressing average excess returns across maturities at each time, $t$, on the 1-year yield and four 1-year forward rates, as follows:

$$\overline{rx}_{t+1} = \gamma_0 + \gamma_1 y_t^1 + \gamma_2 f_t^2 + \gamma_3 f_t^3 + \gamma_4 f_t^4 + \gamma_5 f_t^5 + \bar{\epsilon}_{t+1} \tag{9.2}$$

where $\overline{rx}_{t+1} = \sum_{n=2}^{5} rx_{t+1}^n/4$. Let the right-hand side variables, including the constant term, for each country be collected in vector $f_t$ and let the corresponding estimated coefficients be collected in vector $\hat{\gamma}$. A local CP factor, $CP_t$, is then given by $\hat{\gamma}' f_t$. The CP factors as of date $t$ are later used to predict future excess returns.[5] By construction, the CP factor at time $t$ contains some forward-looking information, as the entire sample period is used in the construction. Dahlquist and Hasseltoft (2013) report that a recursively constructed CP factor for the United States, using only information up to time $t$, has a positive correlation of 0.85 with the full-sample factor. This is consistent with results presented in the appendix to CP (2005), which documents that full-sample forecasts of bond returns and real-time forecasts are quite similar.

The average pairwise correlation of CP factors in the 1975 sample is 0.28, suggesting a positive comovement of bond risk premia across countries, a comovement that has increased over time. In the first half of the 1975–2013 period, the average pairwise correlation was 0.13, while it was 0.58 in the second half. Figure 9.2 plots the local CP factors for the four countries starting in 1975. The gray areas in the graphs represent economic contractions as measured by the NBER for the United States and by the Economic Cycle Research Institute for the remaining countries.[6] Overall, the CP factors and therefore bond risk premia tend to increase during bad economic times, while there are large variations outside recessions.

[5]Cochrane and Piazzesi (2005) note that the estimated $\gamma$ values are tent shaped. We also find a tent-shaped pattern for the United States, using the sample starting in 1975. However, the patterns differ for the remaining countries. Dai, Singleton, and Yang (2004) emphasize that different ways of smoothing yield curves give rise to different patterns, yields that are choppy and less smoothed producing patterns that are more tent shaped. While the U.S. yields we use are unsmoothed yields, yields for the remaining countries are smoothed, so the patterns differ. However, including only the 1-year yield, the 3-year forward rate, and the 5-year forward rate on the right-hand side tends to produce tent shapes for smoothed yields as well, without substantially changing the dynamics of the CP factor.

[6]Recession dates for the United States produced by the Economic Cycle Research Institute line up closely with the ones from NBER.

**TABLE 9.3 Fama–Bliss and Cochrane–Piazzesi Regressions, 1975–2013 Sample**

| | Maturity | $FB_t^n$ | $R^2$ | $CP_t$ | $R^2$ |
|---|---|---|---|---|---|
| Germany | 2 | 0.37 | 0.03 | 0.40 | 0.07 |
| | | (0.37) | | (0.16) | |
| | 3 | 0.59 | 0.05 | 0.83 | 0.09 |
| | | (0.42) | | (0.30) | |
| | 4 | 0.73 | 0.05 | 1.22 | 0.10 |
| | | (0.47) | | (0.42) | |
| | 5 | 0.84 | 0.06 | 1.55 | 0.10 |
| | | (0.52) | | (0.53) | |
| Switzerland | 2 | 0.83 | 0.21 | 0.56 | 0.30 |
| | | (0.22) | | (0.11) | |
| | 3 | 0.91 | 0.14 | 0.92 | 0.27 |
| | | (0.29) | | (0.17) | |
| | 4 | 1.00 | 0.12 | 1.17 | 0.24 |
| | | (0.38) | | (0.21) | |
| | 5 | 1.15 | 0.11 | 1.35 | 0.22 |
| | | (0.49) | | (0.25) | |
| United Kingdom | 2 | 0.34 | 0.03 | 0.41 | 0.08 |
| | | (0.25) | | (0.15) | |
| | 3 | 0.49 | 0.03 | 0.85 | 0.10 |
| | | (0.31) | | (0.27) | |
| | 4 | 0.54 | 0.03 | 1.23 | 0.10 |
| | | (0.36) | | (0.41) | |
| | 5 | 0.53 | 0.02 | 1.50 | 0.10 |
| | | (0.40) | | (0.53) | |
| United States | 2 | 0.74 | 0.08 | 0.43 | 0.17 |
| | | (0.29) | | (0.09) | |
| | 3 | 1.07 | 0.11 | 0.85 | 0.19 |
| | | (0.35) | | (0.18) | |
| | 4 | 1.25 | 0.12 | 1.26 | 0.21 |
| | | (0.42) | | (0.25) | |
| | 5 | 1.00 | 0.06 | 1.46 | 0.19 |
| | | (0.48) | | (0.32) | |

This table presents the results of Fama and Bliss (1987) and Cochrane and Piazzesi (2005) regressions, corresponding to regression equations 9.1 and 9.3. Estimates of constant terms are not tabulated. The sample period is from January 1975 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to 12 lags, within parentheses. Adjusted $R^2$ values are also reported. Data sources are given in Section 9.3.2.

We next predict annual excess returns using the local CP factors, as follows:

$$rx_{t+1}^n = a^n + b_{CP}^n CP_t + \epsilon_{t+1}^n \tag{9.3}$$

The last columns in Table 9.3 present results for the 1975 sample and constitute strong evidence of predictability. The slope coefficients are highly significant across all countries, with $R^2$ values varying between 7% and 30%. For countries in which the FB regressions indicate significance, the explanatory power of the CP regressions is roughly doubled. For countries in which the FB regressions indicate no evidence of predictability, the CP regressions indicate strong evidence in favor of predictability. A potential explanation for this difference in results is that the CP regressions use more information from the yield curve than do the FB regressions.

We consider the time variation in bond risk premia using fitted values from the estimated regression in 9.3. In untabulated results, we find that the bond risk premia occasionally turn negative, particularly during the period after year 2000.[7] Negative bond risk premia imply that nominal bonds in fact are hedges against bad economic times as opposed to being risky assets, such as equity. We analyze this further by computing rolling 60-day betas of nominal bonds with equity for each country. We do so

[7] A fall in term premia on long-term bonds during the 2000s, albeit not to negative levels, has been documented by, for example, Backus and Wright (2007).

**TABLE 9.4 Fama–Bliss and Cochrane–Piazzesi Regressions, 1999–2013 Sample**

| | Maturity | $FB_t^n$ | $R^2$ | $CP_t$ | $R^2$ |
|---|---|---|---|---|---|
| Australia | 2 | −0.15 | 0.00 | 0.38 | 0.21 |
| | | (0.52) | | (0.19) | |
| | 5 | 0.00 | 0.00 | 1.58 | 0.35 |
| | | (1.30) | | (0.54) | |
| Canada | 2 | 0.41 | 0.12 | 0.43 | 0.26 |
| | | (0.18) | | (0.10) | |
| | 5 | 0.72 | 0.06 | 1.56 | 0.34 |
| | | (0.48) | | (0.21) | |
| Denmark | 2 | 0.24 | 0.03 | 0.39 | 0.11 |
| | | (0.23) | | (0.19) | |
| | 5 | 0.82 | 0.05 | 1.57 | 0.18 |
| | | (0.70) | | (0.57) | |
| Germany | 2 | −0.14 | 0.00 | 0.50 | 0.20 |
| | | (0.58) | | (0.15) | |
| | 5 | 0.60 | 0.02 | 1.42 | 0.15 |
| | | (0.90) | | (0.55) | |
| Japan | 2 | 0.38 | 0.12 | 0.21 | 0.32 |
| | | (0.19) | | (0.06) | |
| | 5 | 1.27 | 0.19 | 1.87 | 0.52 |
| | | (0.56) | | (0.27) | |
| Norway | 2 | −0.35 | 0.03 | 0.48 | 0.03 |
| | | (0.41) | | (0.45) | |
| | 5 | −0.07 | 0.00 | 1.48 | 0.04 |
| | | (0.55) | | (1.26) | |
| Sweden | 2 | −0.22 | 0.01 | 0.53 | 0.28 |
| | | (0.47) | | (0.14) | |
| | 5 | 0.02 | 0.00 | 1.37 | 0.19 |
| | | (0.88) | | (0.42) | |
| Switzerland | 2 | 0.17 | 0.00 | 0.46 | 0.33 |
| | | (0.44) | | (0.11) | |
| | 5 | 0.64 | 0.03 | 1.49 | 0.31 |
| | | (0.64) | | (0.33) | |
| United Kingdom | 2 | 0.35 | 0.01 | 0.33 | 0.06 |
| | | (0.47) | | (0.12) | |
| | 5 | 0.98 | 0.09 | 1.61 | 0.16 |
| | | (0.59) | | (0.48) | |
| United States | 2 | 0.23 | 0.01 | 0.50 | 0.27 |
| | | (0.45) | | (0.13) | |
| | 5 | 0.59 | 0.04 | 1.41 | 0.25 |
| | | (0.60) | | (0.34) | |

This table presents the results of Fama and Bliss (1987) and Cochrane and Piazzesi (2005) regressions, corresponding to regression equations 9.1 and 9.3. Estimates of constant terms are not tabulated. The sample period is from December 1999 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to 12 lags, within parentheses. Adjusted $R^2$ values are also reported. Data sources are given in Section 9.3.2.

by collecting daily returns on 10-year bond indexes and daily returns on broad stock market portfolios for each country from Datastream. Consistent with Campbell, Sunderam, and Viceira (2013), who document negative bond betas during the 2000s in U.S. data, we find that bond betas in Germany, Switzerland, and the United Kingdom also turned negative around year 2000 and reached a trough during the financial crisis of 2008 and 2009. Interestingly, the evolution of bond betas is highly positively correlated across countries, suggesting the existence of systematic and global factors driving the riskiness of nominal bonds. A potential explanation for the time-varying nature of bond risk is the time-varying relationship between inflation and economic growth, as pointed out by Burkhardt and Hasseltoft (2012) and David and Veronesi (2013).

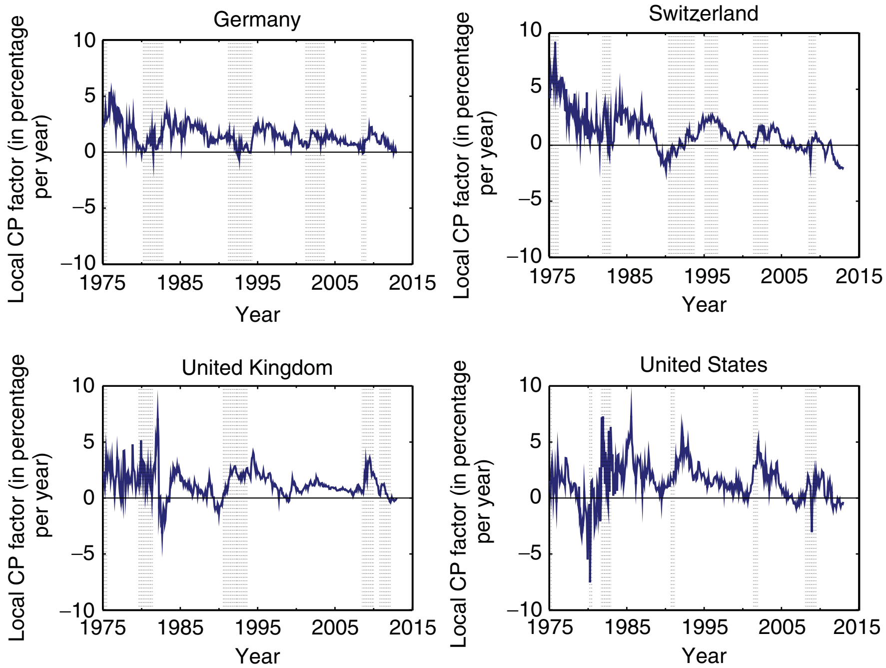


**Figure 9.2** Local CP factors. The figure shows the local CP factors (in % per year) for Germany, Switzerland, the United Kingdom, and the United States. The shaded areas indicate contractions (peak to trough) as dated by the NBER for the United States and by the Economic Cycle Research Institute for the other countries. Source: CRSP, NBER, Economic Cycle Research Institute, and central banks.

We also run the regression in 9.3 using the Datastream sample starting in 1999, with the results being reported in the last columns of Table 9.4. In this broader cross section of countries, we also find support for the CP regressions. Slope coefficients are statistically significant for all countries except Norway, with $R^2$ values as high as 52% in the case of Japan. These results stand in stark contrast to those of the earlier FB regressions, which indicate no predictability for the sample starting in 1999. As the same sample period is used for the FB and CP regressions, the differences in results suggest that the forward–spot spread lost its predictive power over the last 13 years, whereas the CP factors seem to embody predictive information.

In addition to predicting annual excess returns, we also consider the prediction of monthly excess returns (untabulated). For the United States, we collect monthly returns from CRSP for seven portfolios of bonds sorted by maturity. We then construct a monthly CP factor in the same manner as the annual CP factor previously. The results suggest that monthly returns are also predictable but mainly for bonds with maturities greater than 3 years and with $R^2$ values of approximately 1–2%. Since we do not have corresponding monthly returns for the remaining countries, we construct approximate monthly returns using our existing data.[8] As with the U.S. data, we find evidence of monthly predictability mainly for bonds with maturities in excess of 3 years, except in the case of Germany, for which the predictability is weaker. We also analyze the relationship between monthly and annual CP factors within countries, finding that they have correlations of approximately 0.60–0.80. Hence, monthly and annual bond risk premia are positively but not perfectly correlated, suggesting a time variation in the term structure of risk premia.

Following Campbell and Shiller (1991) and Backus et al. (2001), we consider two alternative ways of evaluating the expectations hypothesis. Campbell and Shiller predict the change in long-term rates using the yield spread between long- and short-term bonds, that is, $y_{t+1}^{n-1} - y_t^n = a^n + b^n(y_t^n - y_t^1)/(n-1) + \epsilon_{t+1}^n$, where the expectations hypothesis implies a slope coefficient equal to one. The results for our 1975 sample are similar to those of the FB regressions, as we find statistically significant violations of the expectations hypothesis for Switzerland and the United States but not for the remaining countries. Backus et al. (2001) consider a different form of the expectations hypothesis, namely, that forward rates equal expected future interest rates, that is, $f_t^n = E_t(y_{t+n-1}^1)$, which implies that forward rates are martingales, that is, $f_t^n = E_t(f_{t+1}^{n-1})$. Consequently, they suggest the regression $f_{t+1}^{n-1} - y_t^1 = a^n + b^n(f_t^n - y_t^1) + \epsilon_{t+1}^n$, where the expectations hypothesis implies a slope coefficient equal to one. We run this regression for our 1975 sample and obtain results similar to those of the FB and Campbell–Shiller regressions, except for Germany, for which the expectations hypothesis is violated for 4- and 5-year bonds. Overall, the estimated slope coefficients are less than one and increase with maturity.

[8] Since we only have data on bonds with annual maturities, we approximate the monthly returns by assuming one buys an $n$-year bond and sells it 1 month later at a price computed from the yield of an $n$-year bond but with a maturity equal to $n$ years minus 1 month. The excess return of a bond is over a 1-month rate for the United States but over a 3-month rate for the remaining countries due to data availability.

To summarize our results, the regressions of FB (1987), Campbell and Shiller (1991), and Backus et al. (2001) provide strong evidence against the expectations hypothesis in U.S. data but considerably weaker evidence in international data. In contrast, CP regressions document statistically significant predictability across all countries.

### 9.5.2 Global Predictors of Returns

Having established time variation in bond risk premia stemming from country-specific factors, we next analyze whether there also exists a common global factor driving risk premia. We construct a GCP factor as the GDP-weighted average of all local CP factors, now denoted with subscript c for each country, at time $t$, as follows:

$$GCP_t = \sum_{c=1}^{C} w_{c,t} CP_{c,t} \tag{9.4}$$

where $w_{c,t} = GDP_{c,t} / \sum_{c=1}^{C} GDP_{c,t}$ and where $C = 4$ for the 1975 sample and $C = 10$ for the 1999 sample. The GDP weights used in 1 year are based on the GDP of the previous year. For example, the average weights for the months in the 1975 sample are based on the GDP in 1974 and are 0.19 for Germany, 0.02 for Switzerland, 0.10 for the United Kingdom, and 0.69 for the United States. Hence, the size-weighted global risk factor is dominated by the United States in this sample. Using GDP weights to construct the global factor as opposed to, for example, using the first principal component of the covariance matrix of local CP factors is appealing as it does not introduce an extra layer of forward-looking information. Figure 9.3 plots the global factor together with U.S. recessions and suggests that the level of global bond risk premia tend to increase during U.S. recessions. The figure also shows that global bond risk premia on occasions have turned negative, in particular around 1980 and during the most recent years.

We run the following monthly regression, predicting annual bond excess returns with the global factor, as follows:

$$rx_{t+1}^{n} = a^{n} + b_{GCP}^{n} GCP_t + \epsilon_{t+1}^{n} \tag{9.5}$$

Table 9.5 reports results for the 1975 sample and shows that the global factor predicts local bond returns across all four countries, with the economic magnitude of the regression coefficients and the explanatory power being similar to when using local CP factors, except in the case of Germany, where the global factor doubles the $R^2$ to 20%. As the local U.S. CP factor receives approximately 70% of the weight of the global factor, our results suggest that U.S. bond risk premia carry economically great predictive power for bond returns outside the United States.

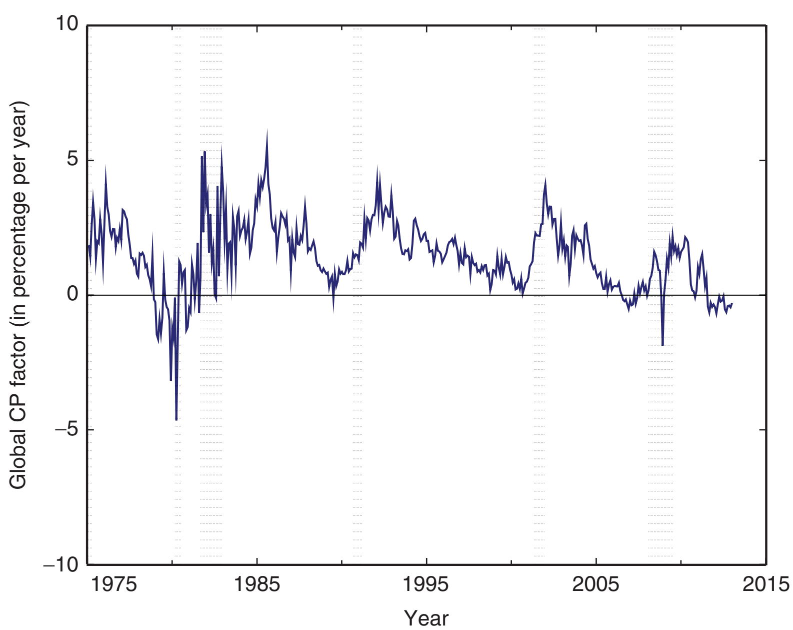


**Figure 9.3** Global CP factor. The figure shows the global CP factor (in % per year). It is a GDP-weighted average of the local CP factors for Germany, Switzerland, the United Kingdom, and the United States. The shaded areas indicate U.S. contractions (peak to trough) as dated by the NBER. Source: CRSP, NBER, and central banks.

**TABLE 9.5 Local and Global Cochrane–Piazzesi Regressions, 1975–2013 Sample**

| | Maturity | $GCP_t$ | $R^2$ | $CP_t$ | $GCP_t$ | $R^2$ | Wald |
|---|---|---|---|---|---|---|---|
| Germany | 2 | 0.50 | 0.21 | 0.15 | 0.50 | 0.22 | [0.00] |
| | | (0.08) | | (0.14) | (0.08) | | |
| | 3 | 0.93 | 0.20 | 0.40 | 0.93 | 0.22 | [0.00] |
| | | (0.15) | | (0.27) | (0.15) | | |
| | 4 | 1.26 | 0.19 | 0.65 | 1.26 | 0.22 | [0.00] |
| | | (0.21) | | (0.40) | (0.22) | | |
| | 5 | 1.55 | 0.19 | 0.87 | 1.55 | 0.21 | [0.00] |
| | | (0.28) | | (0.50) | (0.29) | | |
| Switzerland | 2 | 0.60 | 0.19 | 0.46 | 0.60 | 0.36 | [0.00] |
| | | (0.11) | | (0.13) | (0.13) | | |
| | 3 | 1.05 | 0.20 | 0.72 | 1.15 | 0.34 | [0.00] |
| | | (0.21) | | (0.20) | (0.24) | | |
| | 4 | 1.41 | 0.20 | 0.90 | 1.41 | 0.33 | [0.00] |
| | | (0.29) | | (0.26) | (0.33) | | |
| | 5 | 1.70 | 0.20 | 1.01 | 1.70 | 0.30 | [0.00] |
| | | (0.35) | | (0.30) | (0.40) | | |
| United Kingdom | 2 | 0.40 | 0.08 | 0.34 | 0.40 | 0.13 | [0.00] |
| | | (0.14) | | (0.14) | (0.13) | | |
| | 3 | 0.72 | 0.08 | 0.72 | 0.72 | 0.15 | [0.00] |
| | | (0.26) | | (0.26) | (0.23) | | |
| | 4 | 1.04 | 0.08 | 1.04 | 1.04 | 0.15 | [0.00] |
| | | (0.36) | | (0.38) | (0.32) | | |
| | 5 | 1.35 | 0.09 | 1.24 | 1.35 | 0.15 | [0.00] |
| | | (0.43) | | (0.49) | (0.40) | | |
| United States | 2 | 0.57 | 0.16 | 0.42 | 0.57 | 0.17 | [0.00] |
| | | (0.13) | | (0.46) | (0.12) | | |
| | 3 | 1.11 | 0.17 | 1.12 | 1.11 | 0.19 | [0.00] |
| | | (0.27) | | (0.82) | (0.25) | | |
| | 4 | 1.65 | 0.20 | 1.67 | 1.65 | 0.21 | [0.00] |
| | | (0.38) | | (1.03) | (0.35) | | |
| | 5 | 1.91 | 0.17 | 1.67 | 1.91 | 0.18 | [0.00] |
| | | (0.48) | | (1.25) | (0.45) | | |

The table presents the results of local and global Cochrane and Piazzesi (2005) regressions, corresponding to regression equations 9.5 and 9.6 for maturities $n = 2, 3, 4$, and 5. Estimates of constant terms are not tabulated. When both local and global CP factors are included, the local factor is orthogonalized relative to the global factor. The sample period is from January 1975 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to twelve lags, within parentheses. Adjusted $R^2$ values are also reported. $P$-Values from Wald tests of joint significance are presented in square brackets. Data sources are given in Section 9.3.2.

Table 9.6 presents results for the shorter sample period. Again, we only report results for maturities of 2 and 5 years. Results are consistent with those for the longer sample period in that the global factor predicts local bond returns in all but two countries. In general, the magnitude of the slope coefficients and the explanatory power are somewhat less than when using local CP factors, except in the case of the United States, where in fact the global factor seems to have a somewhat larger effect on bond risk premia.[9]

We also run regressions in which we predict annual excess returns on Citigroup bond indexes using the global factor across 19 countries. For each country, we predict excess returns on five different maturity buckets. In untabulated results, we find estimated slope coefficients to be positive across all countries except Ireland, suggesting that the global factor raises risk premia across countries and maturities. However, the statistical significance is lower in this sample, with only about half the countries having statistically significant regression coefficients.

Finally, we consider predictive regressions with both local and global factors as independent variables, as follows:

$$rx^n_{t+1} = a^n + b^n_{CP}CP_t + b^n_{GCP}GCP_t + \epsilon^n_{t+1} \tag{9.6}$$

[9] We have also predicted returns for the sample starting in 1999 using the global factor computed from the 1975 sample; the slope coefficients are statistically significant, but the explanatory power is somewhat lower.

**TABLE 9.6 Local and Global Cochrane–Piazzesi Regressions, 1999–2013 Sample**

| | Maturity | $GCP_t$ | $R^2$ | $CP_t$ | $GCP_t$ | $R^2$ | Wald |
|---|---|---|---|---|---|---|---|
| Australia | 2 | 0.27 | 0.04 | 0.38 | 0.27 | 0.21 | [0.03] |
| | | (0.19) | | (0.18) | (0.25) | | |
| | 5 | 0.70 | 0.02 | 1.77 | 0.70 | 0.37 | [0.00] |
| | | (0.62) | | (0.41) | (0.76) | | |
| Canada | 2 | 0.47 | 0.20 | 0.31 | 0.47 | 0.28 | [0.00] |
| | | (0.13) | | (0.14) | (0.13) | | |
| | 5 | 1.75 | 0.28 | 1.11 | 1.75 | 0.37 | [0.00] |
| | | (0.31) | | (0.28) | (0.31) | | |
| Denmark | 2 | 0.28 | 0.06 | 0.39 | 0.28 | 0.11 | [0.03] |
| | | (0.17) | | (0.37) | (0.15) | | |
| | 5 | 1.11 | 0.10 | 1.57 | 1.11 | 0.17 | [0.00] |
| | | (0.53) | | (1.14) | (0.45) | | |
| Germany | 2 | 0.37 | 0.12 | 0.54 | 0.37 | 0.20 | [0.00] |
| | | (0.15) | | (0.26) | (0.13) | | |
| | 5 | 1.04 | 0.09 | 1.56 | 1.04 | 0.15 | [0.01] |
| | | (0.53) | | (0.99) | (0.47) | | |
| Japan | 2 | 0.10 | 0.19 | 0.17 | 0.10 | 0.35 | [0.00] |
| | | (0.04) | | (0.06) | (0.03) | | |
| | 5 | 0.70 | 0.21 | 1.69 | 0.70 | 0.54 | [0.00] |
| | | (0.26) | | (0.24) | (0.20) | | |
| Norway | 2 | 0.12 | 0.00 | 0.52 | 0.12 | 0.02 | [0.24] |
| | | (0.19) | | (0.44) | (0.19) | | |
| | 5 | 0.52 | 0.02 | 1.35 | 0.52 | 0.04 | [0.24] |
| | | (0.54) | | (1.33) | (0.52) | | |
| Sweden | 2 | 0.42 | 0.13 | 0.59 | 0.42 | 0.28 | [0.00] |
| | | (0.14) | | (0.20) | (0.12) | | |
| | 5 | 1.41 | 0.15 | 1.05 | 1.41 | 0.20 | [0.00] |
| | | (0.44) | | (0.60) | (0.41) | | |
| Switzerland | 2 | 0.49 | 0.33 | 0.26 | 0.49 | 0.36 | [0.00] |
| | | (0.13) | | (0.16) | (0.13) | | |
| | 5 | 1.37 | 0.23 | 1.38 | 1.37 | 0.31 | [0.00] |
| | | (0.36) | | (0.48) | (0.35) | | |
| United Kingdom | 2 | 0.44 | 0.12 | 0.00 | 0.44 | 0.12 | [0.00] |
| | | (0.14) | | (0.26) | (0.14) | | |
| | 5 | 1.18 | 0.09 | 1.47 | 1.18 | 0.15 | [0.00] |
| | | (0.47) | | (0.82) | (0.51) | | |
| United States | 2 | 0.77 | 0.32 | –0.89 | 0.77 | 0.34 | [0.00] |
| | | (0.22) | | (0.51) | (0.24) | | |
| | 5 | 2.30 | 0.32 | –4.98 | 2.30 | 0.43 | [0.00] |
| | | (0.55) | | (1.49) | (0.67) | | |

The table presents the results of local and global Cochrane and Piazzesi (2005) regressions, corresponding to regression equations 9.5 and 9.6 for maturities $n =$ 2 and 5. Estimates of constant terms are not tabulated. When both local and global CP factors are included, the local factor is orthogonalized relative to the global factor. The sample period is from December 1999 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to twelve lags, within parentheses. Adjusted $R^2$ values are also reported. $P$-Values from Wald tests of joint significance are presented in square brackets. Data sources are given in Section 9.3.2.

To simplify interpretation of the results, we first orthogonalize the local factors relative to the global factor. More specifically, we regress the local factors on the global factor and treat the residuals as the truly local factors. The last columns in Table 9.5 present the results for the 1975 sample. First, the local and global factors are jointly significant in all countries. Second, for all countries except the United States, the explanatory power of these joint regressions is greater than when using only local factors. A similar picture emerges when we run the regressions for the broader sample of ten countries, as reported in the last columns of Table 9.6. The local and global factors are jointly significant in all countries except Norway. Overall, our regression results suggest that the expectations hypothesis is violated in international data, local bond risk premia being driven by both local and global factors, and that U.S. bond risk premia matter for international risk premia.

We find that the global factor had significant predictive power across all countries leading up to the financial crisis in 2007 but experienced a marked decrease in predictive power over the 2008–2013 period. The deterioration in the predictive power of the CP factor during the financial crisis has also been noted by Sekkel (2011) and Buraschi and Whelan (2012). The underlying reason for this is unclear but may be related to the extraordinary monetary policy measures employed since 2008 and to the large contemporaneous changes in liquidity funding that are not necessarily captured by the CP factor. Analyzing the latter, Fontaine and Garcia (2012) find that tightness of liquidity funding indeed affects risk premia on U.S. bonds. It is also possible that the so-called zero bound on nominal interest rates affected predictability, as the Federal Funds rate and other short-term rates dropped to virtually zero in 2008 and 2009. However, the effect of the zero bound on the predictability of annual excess returns on long-term bonds is less clear and might even be muted, as the dynamics and responsiveness to news of long-term rates remained similar to their previous levels, as documented by Swanson and Williams (2014a,b). However, the fact that the predictive power of the CP factor dropped fairly significantly during and after the Great Recession suggests that there are state dependencies in bond risk premia. We explore one such potential dependency in Section 9.6.2.

In addition, it is noteworthy that all our results pertain to predicting bond excess returns in local currencies. One might wonder whether the predictive power of the global factor carries over to an international bond strategy that takes into account movements in exchange rates. Dahlquist and Hasseltoft (2013) conduct such an analysis by considering the excess return for a U.S. investor who borrows for 1 year in USD, invests in a foreign government bond with maturities of 2–5 years, and then converts the proceeds back into USD after 1 year.

The return on such a strategy can be written as $rx^n_{FX,t+1} = \Delta s_{t+1} + r^n_{t+1} - y^1_{U.S.,t}$ for a foreign currency versus the USD, where $y^1_{U.S.,t}$ is the 1-year yield in the United States. Dahlquist and Hasseltoft (2013) find that the global factor loses most of its predictive power in such a strategy. To understand why, one can decompose the overall return into two parts by adding and subtracting the local short-term rate, $y^1_t$. The first part equals the 1-year FX excess return, $\Delta s_{t+1} + y^1_t - y^1_{U.S.,t}$, and the second part equals the 1-year local bond excess return, $r^n_{t+1} - y^1_t$. While the global factor predicts the second part, that is, the local bond excess return, it cannot predict exchange rate movements. Including FX returns in the overall strategy therefore introduces noise from the perspective of the global factor, with a corresponding drop in predictive power. However, Dahlquist and Hasseltoft (2013) demonstrate that one can construct a global factor based on dollar returns, called FXGCP, which has predictive power similar to that of GCP for local returns. The FXGCP factor has a correlation of approximately 0.50 with the local return GCP factor.

While our study concerns in-sample regressions, results concerning the out-of-sample performance of the FB and CP regressions are mixed. For example, Kessler and Scherer (2009) note that the CP regressions generate positive out-of-sample performance with information ratios reaching 1.5, the best performance being achieved with a short forecasting horizon of 1 month. On the other hand, Thornton and Valente (2012) find that predictive models using forward rates generate no significant economic value for investors. They also note a deterioration in predictive performance during the Great Recession, as we discussed earlier.

Finally, Dahlquist and Hasseltoft (2013) incorporate the global factor into an affine term structure model and analyze how the factor affects the cross section of yields. They find that the yield level is only slightly affected by variations in the global factor. While a rise in the global factor increases yields through higher risk premia, it also decreases yields as it is associated with lower expectations of future short-term rates. Movements in the global factor give rise to two offsetting effects, leaving the effect on yield levels muted. The global factor therefore seems to act as a hidden factor in the term structure, as discussed in the introduction to this chapter.

## 9.6 UNDERSTANDING BOND RISK PREMIA

The fact that bond excess returns contain a strongly predictable component raises the natural question of "Why?" Can we explain the strong evidence on predictability using rational risk-based explanations? The classical risk-based argument suggests that assets whose returns covary positively with investor well-being should be subject to positive risk premia on average. Furthermore, investors' required compensation for risk seems to move countercyclically, being high in bad times and low in good times as documented by, for example, Fama (1986), Stambaugh (1988), and Fama and French (1989).

### 9.6.1 Links to Economic Growth

We study whether the evidence of predictability provided previously can be explained by business cycles and economic growth. Our empirical evidence suggests that a global factor of risk premia exists, and if it is linked to global economic growth, the classical risk-based explanation may be true.

We evaluate this by relating the estimated global factor to global economic growth. More specifically, we consider growth in industrial production in three large economic areas, that is, the EMU, OECD, and United States. We start by regressing annualized economic growth over 1, 2, and 3 years on the change in the global factor across the same time horizons, measuring

**TABLE 9.7 The Global CP Factor and Industrial Production Growth**

| | Horizon | $\Delta GCP_t$ | $R^2$ | $GCP_t$ | $R^2$ |
|---|---|---|---|---|---|
| EMU | 1 | –0.57 | 0.02 | 0.37 | 0.01 |
| | | (0.27) | | (0.39) | |
| | 2 | –0.59 | 0.11 | 0.67 | 0.07 |
| | | (0.19) | | (0.33) | |
| | 3 | –0.48 | 0.18 | 0.83 | 0.22 |
| | | (0.12) | | (0.24) | |
| OECD | 1 | –0.72 | 0.05 | 0.56 | 0.03 |
| | | (0.23) | | (0.30) | |
| | 2 | –0.70 | 0.20 | 0.87 | 0.17 |
| | | (0.18) | | (0.27) | |
| | 3 | –0.58 | 0.30 | 0.96 | 0.35 |
| | | (0.11) | | (0.21) | |
| United States | 1 | –1.02 | 0.09 | 0.96 | 0.08 |
| | | (0.25) | | (0.33) | |
| | 2 | –0.90 | 0.22 | 1.26 | 0.24 |
| | | (0.21) | | (0.31) | |
| | 3 | –0.75 | 0.30 | 1.30 | 0.38 |
| | | (0.14) | | (0.27) | |

The table presents the results of regressions of industrial production growth on the global CP factor. The left-hand side of the table presents contemporaneous regressions run for annualized growth in industrial production on the change in the global CP factor across horizons of 1, 2, and 3 years for the EMU, OECD, and United States. Estimates of constant terms are not tabulated. The right-hand side of the table presents a predictability regression run for annualized growth in industrial production on the global CP factor across horizons of 1, 2, and 3 years for the EMU, OECD, and United States. The sample period is from January 1975 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to and including the forecasting horizon, within parentheses. Adjusted $R^2$ values are also reported. Data sources are given in Section 9.3.2.

the contemporaneous relationship between economic growth and changes in risk premia. If the notion of countercyclical risk premia holds true, we would expect negative and significant slope coefficients. Indeed, we find that to be the case, as reported in the first columns of Table 9.7. A one-percentage-point increase in global bond risk premia is associated with a contemporaneous drop of 0.5–1% in annualized economic growth, all regression coefficients being statistically significant.

Having established a contemporaneous negative link between growth and bond risk premia, we turn to predictive regressions. We study whether the global factor has predictive power for growth by regressing future 1-, 2-, and 3-year annualized economic growth on the global factor as of time $t$. As reported in the last columns of Table 9.7, we find significant evidence of predictive power. For the OECD and United States, the global factor is a statistically significant predictor across all horizons, whereas for the EMU it is significant only across the 2- and 3-year horizons. A one-percentage-point increase in global bond risk premia is associated with a 0.5–1.3% annualized increase in future economic growth. Our results are consistent with those of Plosser and Rouwenhorst (1994), who document that term structures in Germany, the United Kingdom, and the United States can forecast future economic activity and with those of the NBER working paper version of CP (2005), which demonstrates that the CP factor can predict GDP growth across horizons longer than 1 year.

In addition, we consider the relationship between global bond risk premia and the OECD composite leading indicator. The OECD economic leading indicator is a time series formed by aggregating various economic indicators (e.g., consumer sentiment indicators, business climate indicators, and the purchasing managers' index) for each country to form a composite indicator that anticipates economic movements and turning points. Figure 9.4 plots the lead–lag correlation between the global factor at time $t$ and the OECD indicator at time $t + l$, where $l$ extends from $-36$ months to 36 months. The figure shows a negative contemporaneous correlation between risk premia and the economic leading indicator. Furthermore, a drop in the leading indicator leads a rise in risk premia by 1 year with a correlation of approximately $-0.40$. Moreover, a rise in risk premia signals improved economic conditions starting 1 year hence. Overall, the figure supports our earlier regression findings.

To summarize, our results suggest that a rise in global bond risk premia is associated with a contemporaneous drop in economic growth but signals improved future growth. These results are supportive of the classical risk-based explanations, suggesting a tight link between bond risk premia and business cycles.

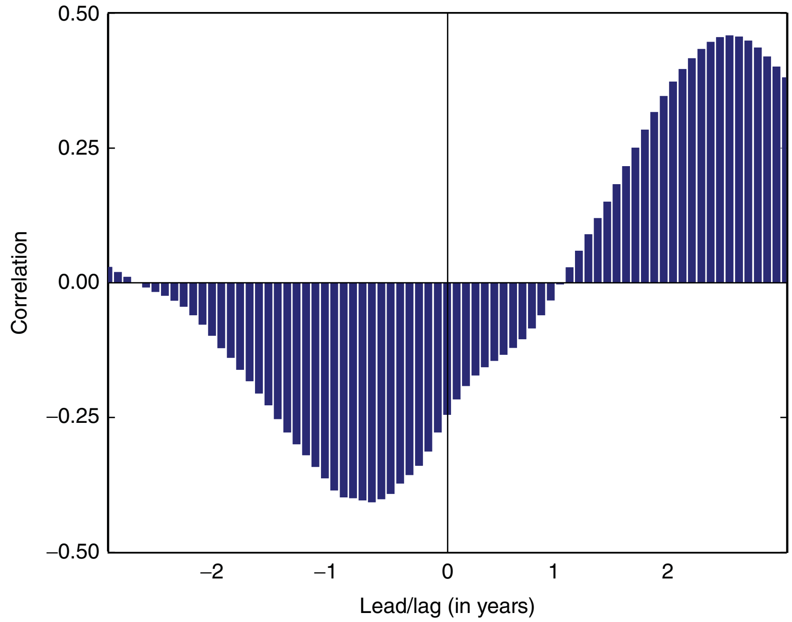


**Figure 9.4** Relationship between global CP factor and OECD leading indicator. The figure shows the lead–lag correlations between the global CP factor and an economic leading indicator covering all 30 OECD member countries. The global CP factor is as of date $t$ and the OECD leading indicator is as of date $t + l$, where $l$ refers to the lead (if negative) or lag (if positive). The OECD economic leading indicator is a time series formed by aggregating various economic indicators (e.g., consumer sentiment indicators, business climate indicators, and the purchasing managers' index) for each country to anticipate economic movements and turning points. Source: CRSP, NBER, and central banks.

### 9.6.2 State Dependency

While an extensive literature has analyzed whether asset returns are predictable, fewer papers have conditioned the predictive regressions on the state of the business cycle. Theory suggests that risk premia on risky assets should be higher in bad economic times than in good times. Focusing on G7 equity markets, Henkel, Martin, and Nardari (2011) combine predictive return regressions with a regime-switching model and indeed find that risk premia in equity markets are higher during recessions. In fact, they find that traditional return predictors, such as dividend yields and term structure variables, are statistically significant only during recessions and are insignificant during business cycle expansions. This suggests a state dependency in risk premia on risky assets.

We build on this notion and analyze whether a similar dependency exists for global bond risk premia. We consider only our sample starting in 1975, as a longer time period is necessary for capturing business cycle variations. We predict local bond returns using our global factor, introducing a dummy variable, $D_t$, that takes the value of one in recessions. We use data from the Economic Cycle Research Institute to define recessions for Germany, Switzerland, and the United Kingdom but use NBER-dated recessions for the United States. Using our dummy variables, we run the following regression:

$$rx^n_{t+1} = a^n D_t + b^n (1 - D_t) + c^n GCP_t D_t + d^n GCP_t (1 - D_t) + \epsilon^n_{t+1} \tag{9.7}$$

We report the results in Table 9.8, which shows statistically significant coefficients for both regimes with similar economic magnitudes. Testing the null hypothesis of equal slope coefficients fails to reject, as shown by the $p$-values. Hence, the global factor seems to drive risk premia during both economic expansions and contractions. Interestingly, this stands in contrast to earlier findings regarding equity markets. The explanation for this may lie either in differences between the dynamics of equity and bond returns or in the dynamics of the typical predictors used in stock and bond markets. We leave this matter for further research.

## 9.7 CONCLUSION AND OUTLOOK

This chapter provides empirical evidence regarding the predictability of international bond risk premia. Our results suggest that the expectations hypothesis is violated in international data, as we document considerable time variation in bond risk premia. While the classical FB (1987) regressions fail to consistently predict excess returns internationally, the well-known CP (2005) regressions do predict excess returns across countries. We also consider a global forecasting factor that we find predicts local bond excess returns significantly, suggesting a commonality in risk premia across international bond markets. The global factor is

**TABLE 9.8 State-Dependent Global Cochrane–Piazzesi Regressions**

| | Maturity | $GCP_tD_t$ | $GCP_t(1-D_t)$ | $R^2$ | Wald |
|---|---|---|---|---|---|
| Germany | 2 | 0.46 | 0.52 | 0.21 | [0.71] |
| | | (0.18) | (0.10) | | |
| | 3 | 0.84 | 0.99 | 0.20 | [0.58] |
| | | (0.27) | (0.19) | | |
| | 4 | 1.17 | 1.36 | 0.19 | [0.58] |
| | | (0.36) | (0.29) | | |
| | 5 | 1.47 | 1.68 | 0.19 | [0.62] |
| | | (0.43) | (0.38) | | |
| Switzerland | 2 | 0.45 | 0.50 | 0.23 | [0.79] |
| | | (0.20) | (0.10) | | |
| | 3 | 0.83 | 0.85 | 0.25 | [0.95] |
| | | (0.37) | (0.18) | | |
| | 4 | 1.15 | 1.16 | 0.25 | [0.98] |
| | | (0.48) | (0.26) | | |
| | 5 | 1.35 | 1.41 | 0.24 | [0.90] |
| | | (0.56) | (0.33) | | |
| United Kingdom | 2 | 0.30 | 0.52 | 0.13 | [0.16] |
| | | (0.17) | (0.16) | | |
| | 3 | 0.56 | 0.95 | 0.13 | [0.20] |
| | | (0.32) | (0.28) | | |
| | 4 | 0.79 | 1.36 | 0.13 | [0.19] |
| | | (0.46) | (0.39) | | |
| | 5 | 0.98 | 1.77 | 0.14 | [0.15] |
| | | (0.57) | (0.46) | | |
| United States | 2 | 0.58 | 0.58 | 0.22 | [0.95] |
| | | (0.11) | (0.16) | | |
| | 3 | 1.22 | 1.12 | 0.23 | [0.57] |
| | | (0.23) | (0.32) | | |
| | 4 | 1.78 | 1.65 | 0.23 | [0.61] |
| | | (0.33) | (0.45) | | |
| | 5 | 1.96 | 1.95 | 0.20 | [0.98] |
| | | (0.51) | (0.56) | | |

This table presents the results of global Cochrane and Piazzesi (2005) regressions with state-dependent coefficients, corresponding to regression equation 9.7. The recession dummy corresponds to economic contractions according to the Economic Cycle Research Institute for Germany, Switzerland, and the United Kingdom and to NBER-dated recessions for the United States. Estimates of state-dependent constant terms are not tabulated. The sample period is from January 1975 to December 2013. Point estimates are reported with Newey and West (1987) standard errors, accounting for conditional heteroskedasticity and serial correlation up to 12 lags, within parentheses. Adjusted $R^2$ values are also reported. $p$-Values from Wald tests of equal coefficients are presented in square brackets. Data sources are given in Section 9.3.2.

closely related to U.S. bond risk premia, which indicates that U.S. risk premia matter for international bond returns. Furthermore, the dynamics of the global factor are consistent with classical asset pricing theory, as they are countercyclical and predict future economic growth. However, the predictive performance of the CP regressions across countries deteriorated starting in the Great Recession. The exact reason for this is unclear, but potential explanations are the extraordinary measures taken by the central banks worldwide, the rise in fiscal policy uncertainty, and the great time variation in funding and liquidity conditions – none of which are likely to be fully captured by typical yield-based predictors. Furthermore, the high levels of government debt around the world are likely to reinforce the roles of fiscal policy, credit risk, and liquidity risk in determining bond risk premia around the world. The good news is that this leaves plenty of opportunities for researchers.

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