The pricing of G7 sovereign bond spreads –

the times, they are a-changin [1]

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Abstract

Against the background of the current debate about fiscal sustainability in several advanced economies, this paper estimates determinants of G7 sovereign bond spreads, using high-frequency proxies for market expectations about macroeconomic fundamentals and allowing for time-varying parameters. The paper finds substantial asymmetry in the importance of country fundamentals and considerable time variations in the pricing of risks. There has been a reduced pricing of several risk factors in the years preceding the financial crisis, and either an over-pricing of risk or the pricing of a re-denomination risk of euro area bonds during the European sovereign debt crisis.

JEL-codes: E43, E44, F34, G15

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1. Introduction

Since the onset of the European sovereign debt crisis in early 2010, the economics profession has shown a renewed interest in the pricing of sovereign bonds. A question at the heart of the policy debate is to which extent market prices of sovereign bonds reflect economic fundamentals in an appropriate fashion, or whether swings in risk appetite have led to an under-pricing of risk prior to the global financial crisis, and possibly an over-pricing of risk during the European sovereign debt crisis (Aizenman et al. 2013).

One distinguishing feature of the current situation compared to previous crises is that fiscal sustainability concerns are not so much an issue for emerging market economies, but are instead mainly present for advanced economies. And even if the focus of the discussion currently rests on the euro area, fiscal sustainability concerns have also arisen in many other advanced economies, inside and outside Europe. Even several countries with long-standing excellent credit ratings have been affected. For instance, the United States lost their AAA rating (which they had held for 70 years) by Standard and Poors in August 2011, on concerns about the government's budget deficit and rising debt burden. Subsequently, also France was downgraded from AAA to AA+ by Standard and Poors in January 2012. Similarly, Japan was downgraded by Moody's in August 2011, from Aa2 (the third-best rating) to Aa3.

As a consequence, we need to understand better the pricing mechanisms in sovereign debt markets in advanced economies. The earlier literature has typically studied emerging markets (Edwards 1984, 1986; Uribe and Yeu 2006), and most of the literature on advanced economies has dealt with euro area countries, in the uprun to European economic and monetary union (EMU) and in its early years (Favero et al. 1997, Codogno et al. 2003) as well as more recently during the sovereign debt crisis (Bernoth and Erdogan 2012; Borgy et al. 2011). Surprisingly little is known, however, on the pricing of sovereign debt in other advanced economies, and the pricing of low-yielding debt in particular, with few exceptions like Gruber and Kamin (2012) which focus on a panel of OECD and G7 countries.

From an econometric point of view, the analysis of yield spreads between high-yielding and low-yielding bonds is obviously highly promising, given that there is typically also more variability in the data that facilitates the identification of possible determinants of yield spreads. In the light of the recent developments, however, it has become important to broaden the perspective to other advanced economies, and to study what determines the spreads between low-yielding bonds. This is what we will do in the current paper, by studying the sovereign bond markets of the G7 countries over the last two decades.

As mentioned above, a key aspect of the current discussion is to what extent market prices reflect fundamentals, and how this has changed over time. To get at the evolution of pricing patterns, the model employed in this paper allows for time variation in the coefficients, which evolve as random walks, and stochastic volatility in the error term. In order to estimate the role of macroeconomic fundamentals, the paper differs relative to the existing literature in two important ways.

First, market prices are likely to reflect *expectations* about the evolution of fundamentals much more than past realised values (Laubach 2009). Therefore, all fundamental variables used in the current paper (namely debt to GDP ratios, current account, real GDP growth, unemployment and inflation) are forward-looking. The use of Consensus Economics data allows us to have a set of expectations by market participants (as opposed to forecasts by other institutions, which have often been used in previous studies) for all these variables for the G7 countries at a monthly frequency (thus avoiding the interpolation of annual or semi-annual forecasts or of quarterly realised macroeconomic data, as often done in the existing literature).[4] By using expectations data, we furthermore avoid the complication of real-time versus ex post vintage data, given that expectations data are not revised.

The second innovation of the paper is that we allow a relaxation of a commonly imposed assumption – when analysing the determinants of sovereign bond spreads, the existing studies tend to use *relative* variables (i.e. the difference between macroeconomic fundamentals in a given country and the benchmark country). This approach imposes an untested restriction on the coefficients of the econometric model, namely that the fundamentals in both countries are equally important in determining the spread, and it turns out that relaxing this restriction leads to a much better understanding of the underlying pricing mechanisms.

[4] Ciarlone et al. (2008) and Ejsing et al. (2012) use the GDP growth expectations from Consensus Economics, and Monfort and Renne (2011) those for long-term interest rates. The only other paper that employs the whole set of Consensus Economics forecasts for macroeconomic and fiscal variables is Nickel et al. (2011), which studies Eastern European countries and Turkey.

There are two key findings of this paper. First, for a spread of any country relative to a safe haven government bond (such as the U.S. or German bonds), the countries' macroeconomic fundamentals are bound to be considerably more influential determinants of the spread than the fundamentals of the benchmark country. The closer the two bonds are to being substitutable, the more symmetric is the impact of the respective fundamentals. Second, there are considerable time variations in the role of the various determinants. For instance, during the dot-com bubble, expectations of U.S. GDP growth lowered U.S. yields, whereas no such effect is found for the other time periods. Similarly, we find that several risk factors have not been priced in the years preceding the financial crisis. This pattern is particularly pronounced for the determinants of the Italian-German and the French-German spreads, i.e. for spreads of the euro area member countries, where macro fundamentals, general risk aversion and liquidity risks used to be priced in the uprun to monetary union and following the outbreak of the financial crisis, but not in the first years of monetary union.

These findings as well as some counterfactual experiments support the belief that swings in risk appetite have led to an over-pricing of risk during the European sovereign debt crisis or the pricing of a risk that government bonds of some euro area countries might get re-denominated in other currencies than the euro.

The paper proceeds as follows: Section 2 provides an overview of the related literature. The data and the econometric methodology are explained in Section 3. The results are presented and discussed in Section 4, and subjected to several robustness tests. Section 5 tries to get at the hypotheses of under-pricing and over-pricing of risk. Section 6 concludes.

2. Literature review

There is a large literature on the pricing of sovereign bonds to which this paper connects. Much of the earlier literature has analysed spreads of government bonds in *emerging market economies* relative to some "safe" bond, typically those issued by the U.S. treasury. Early contributions are Edwards (1984, 1986), who studies determinants of interest rate spreads charged for bank loans to developing countries and for bonds issued by their sovereigns, and finds that international financial markets had often not anticipated future payment difficulties of the debtor countries. Uribe and Yue (2006) show, inter alia, that macroeconomic fundamentals affect emerging markets' spreads, which in turn exacerbates their business-cycle fluctuations. The importance of macroeconomic fundamentals is also confirmed by Duffie et al. (2003) and Hilscher and Nosbusch (2010), whereas Diaz Weigel and Gemmill (2006) report that they explain only relatively little of the variance of developing countries' bonds, with the bulk being explained by global and regional factors.

A second strand of the literature on sovereign bond spreads deals with *the European case*. In the uprun to EMU, much attention was devoted to the role of exchange rate expectations in determining European spreads (which are typically defined relative to Germany), as for instance in Favero et al. (1997). For the first years of EMU, the convergence of long-term government bond rates of euro area countries was a widely studied phenomenon (see, e.g., Ehrmann et al. 2011), while the importance of international risk factors in determining the (small) spreads has been highlighted by Codogno et al. (2003) as well as by Manganelli and Wolswijk (2009). Also market liquidity has been identified as another important factor during the tranquil early years of EMU (Gomez-Puig 2006), even if not necessarily for all euro area countries (Favero et al. 2010). Moving into the financial crisis, a number of studies noted the increased importance of macroeconomic fundamentals, such as openness and the terms of trade (Maltritz 2012) or the debt burden of countries (Bernoth and Erdogan 2012, Bernoth et al. 2012, Borgy et al. 2011)). Also announcements of bank rescue packages, which transferred risk from the private sector to the government, have been found to have had a substantial impact on euro area spreads during the global financial crisis (Attinasi et al. 2010). The increased importance of fundamentals coincided with a reduced role for global factors in determining spreads, as investors obviously discriminated more across countries (Barrios et al. 2009). Furthermore, Beber et al. (2008) have shown that for the bond market in the euro area countries, investors care about credit quality and liquidity, but with variations over time.

A related set of papers is concerned with the determinants of yield spreads in *monetary unions other than EMU*. Comparing the pricing of sovereign credit risks for U.S. states with those of euro area countries, Ang and Longstaff (2013) provide evidence that there is much less systemic risk among U.S. than among euro area sovereigns. An analysis of bond yield spreads of German, Spanish and Canadian sub-national governments by Schuknecht et al. (2009) reveals that markets tend not to price the fiscal burden of these governments if these are part of a fiscal transfer arrangement, suggesting that the credibility of non-bail out clauses matters for the pricing of risk.

A very recent literature tackles the important question whether there has been *contagion in the sovereign debt crisis*. The overall picture that emerges is that there is compelling evidence for the presence of contagion. For instance, Amisano and Tristani (2012) model sovereign yield spreads in the context of a Markov switching approach, which allows a country's probability of jumping to a crisis state to depend on the occurrence of a crisis in other countries, and find this to be the case. In the context of a global VAR (GVAR) framework, Favero (2012) looks at impulse responses of local spreads to shocks in the spreads of other euro area countries. Interestingly, he replaces the usual measures of distance in the GVAR models, like trade or financial integration, by differences in fiscal fundamentals. Favero and Missale (2012) stress the time-varying importance of the global risk factor, and the fact that fiscal fundamentals matter more when global risk is priced more strongly, pointing to contagion driven by shifts in market sentiment. Claeys and Vasicek (2012), De Santis (2012) and Missio and Watzka (2011) report that rating announcements have generated contagious effects in the euro area. Calice et al. (2013) show that the liquidity of the sovereign CDS market has spilled over to sovereign bond spreads in several countries, including Greece, Ireland and Portugal.[5] Finally, Zhang et al. (2011) develop a measure of conditional probability of default on the debt of a given country, dependent on the default of another country.

Beyond studies focusing on emerging market economies on the one hand and the euro area on the other hand, there are surprisingly few contributions to this literature. Some studies use large *international panels*, and mostly analyse the importance of common factors in the pricing of sovereign bonds (such as Martell 2008) or CDS markets (such as Longstaff et al. 2011). Some papers use such a panel to construct a comparator group for the euro area countries during the sovereign debt crisis: Aizenman et al. (2013), Beirne and Fratzscher (2013) as well as De Grauwe and Ji (2012) show that during the sovereign debt crisis spreads in the most affected countries of the euro area were

[5] Fontana and Scheicher (2010) study the relative pricing of euro area sovereign CDS and the underlying government bonds, and find that market integration for bonds and CDS varies across countries, such that price discovery takes place in the CDS markets for some countries, and in the bond market for others.

considerably higher than those of comparable countries outside the euro area. The study with the closest match in terms of country coverage is Gruber and Kamin (2012), which applies a panel regression technique to model the dynamic of long-term bond yield in the G7 countries. They show that the fiscal position has a significant impact on the yields. Dungey et al. (2000) also focus on advanced economies and use a factor model to the spreads between Australia, Canada, Germany, Japan and the UK relative to the United States over the period 1991 to 1999. Their analysis shows that the common factors exhibit long swings and help explaining the strong persistence observed in the spreads. Australian and Canadian spreads are mostly affected by the world factor, whereas Germany, the UK and in particular Japan show strong individual country effects.

In the current paper, we will try to shed light on the determinants of sovereign bond spreads of G7 countries, with a particular view towards the role of global risk aversion and macroeconomic fundamentals. To study their evolution over time, we will allow for time-varying parameters, an approach that has been applied in several recent contributions to this literature. While Bernoth and Erdogan (2012) estimate a non-parametric panel, Aßmann and Boysen-Hogrefe (2012) and Pozzi and Wolswijk (2012) employ ARCH-type models. All three papers are concerned with the spreads of euro area countries, and find overwhelming evidence for the presence of time variations, with an increasing reaction to country-specific factors during the financial crisis. These results point to a need to allow for time variation in parameters, an avenue that we also follow in the current paper.

3. Data and methodology

Our dataset covers the period from May 1993 until December 2011. The frequency, the set of countries, and the sample period reflect the availability of forecast data for macroeconomic fundamentals from Consensus Economics.[6] The dataset comprises 224 observations for each country.

[6] While forecasts for several macroeconomic variables are available also for other countries, and over longer periods, forecasts for the fiscal position at monthly frequency are only provided for the G7 countries for a sufficiently long time-span.

Our data for government bond yields are based on 10-year benchmark bonds as calculated by Thomson Reuters and provided by Datastream. Summary statistics and a graphical representation are provided in Table 1 and Figure 1. From the figure, it is immediately apparent that there is substantial comovement of the G7 yields. This is particularly the case for five of the seven countries, whereas yields for Italy used to be considerably higher at the beginning of the sample and increased relative to the others towards the end, i.e. during the sovereign debt crisis. For most of EMU, however, also Italian yields were at levels similar to the other countries', and co-moving very strongly. The second exception relates to Japanese yields, which are obviously considerably lower than those of all other countries. Still, however, the comovement is substantial – correlation coefficients of Japanese yields with those of the other G7 countries range from 0.77 (with the United States) to 0.90 (with Italy). This is in line with the findings of the earlier literature that much of the movements in yields are explained by a global factor.

Table 1 and Figure 1 around here

Table 1 furthermore reports summary statistics for spreads, defined once against the yields of the United States, and once against those of Germany. We consider spreads against Germany particularly appropriate when studying Italian and French yields, given that these countries have shared a common currency with Germany for most of our sample period. For an international investor, the relevant question therefore is likely to be whether, conditional on investing in bonds denominated in euro, to invest in France or in Germany on the one hand, or to invest in Italy or in Germany on the other hand. The issue is less clear cut if it comes to the British yields – while the UK is part of the European Union (EU), it is not part of the euro area, and hence does not share the same currency as Germany. Still, we decided to study British-German spreads to start with, also in order to be able to compare the Italian and French spreads, which as we will see are heavily affected by the sovereign debt crisis, to another EU country. For all other countries, spreads against the yields of the United States are most likely the relevant benchmark. Given the strong comovement of German and U.S. yields (with a correlation coefficient of 0.92), the different benchmark should not make a large difference. Furthermore, we will cross-check all our results against selecting the alternative benchmark country in defining the yield spreads.

Looking at the summary statistics of spreads in Table 1, it is apparent that the spreads on average are rather small (with the exception of Japan), and that their variability is not particularly large (with the exception of Italy). This is especially the case if we were to compare these spreads with those of emerging market countries or some euro area countries like Greece, Portugal or Ireland. This notwithstanding, we think that the determinants of the G7 bond yield spreads deserve further study.

When studying spreads, the underlying idea is that these are defined against a benchmark that is close to risk-free. Accordingly, the spread should be based on the pricing of risk of a possible investment relative to the risk-free rate (or its proxy). The literature usually distinguishes four types of potential determinants – exchange rate risk, liquidity risk, credit risk and general risk aversion. We will now explain how we control for each of these factors.

For a proxy of exchange rate risk, we follow the approach suggested by Favero et al. (1997) and subsequently applied in several other studies (e.g. Bernoth and Erdogan 2012, Gómez-Puig 2006) and subtract the difference between the 10-year swap rate in the currency of denomination of the corresponding bond and the 10-year swap rate in the currency of the benchmark bond (U.S. dollars or D-Mark/Euro, respectively). These data are also provided by Datastream. Of course, for the time of EMU, this proxy for exchange rate risk is equal to zero for the French-German and Italian-German spread. When entering this proxy for exchange rate risk as an explanatory variable into the econometric model, it turns out that the estimated parameters for this variable are extremely tightly estimated, statistically not significantly different from one, and show very little time variation. Accordingly, we decided to impose the correction for exchange rate risk ex ante by directly subtracting the swap rate differential from the spreads, as this saves estimating an extra coefficient.

To control for liquidity risk, we follow the literature (e.g. Gómez-Puig 2006) and include the overall outstanding amounts of public debt as provided by the Bank for International Settlements. We add the domestic and the international total outstanding amounts, and subtract those with a remaining maturity below one year. Given that these data are available at the quarterly frequency, we linearly interpolate them to the monthly frequency.[7]

A second block of explanatory variables relates to general risk aversion. A conventionally used measure in the literature (e.g., Codogno et al. 2003, Bernoth and Erdogan 2012) is the corporate bond yield spread in the United States. Our proxy is given by the spread between Moody's Seasoned Baa and Aaa Corporate Bond Yield as provided by the Board of Governors of the Federal Reserve System. The underlying idea is that this spread is positively correlated with risk aversion, as in a more risk-averse environment, less secure corporates are expected to pay an increased premium relative to the safer corporates. As a second proxy for general risk aversion, we use the VIX, i.e. a measure of the implied volatility of S&P 500 index options, which is a well-established proxy in the literature (among many others see, e.g., Longstaff et al. 2011).

Finally, we include various macroeconomic fundamentals, which are our proxies for credit risk. The main novelty in this paper is the use of Consensus Economics forecasts[8], which allow us to construct expectations for several variables at the monthly frequency (and thus spare us the need to interpolate lower frequency forecasts). As we are interested in the role of expectations for market prices, another advantage is that many of the forecasters polled by Consensus Economics are financial institutions. This makes us believe that the forecasts are more likely to reflect market expectations than forecasts by public institutions, such as the often-used fiscal forecasts by the European Commission.

The survey is conducted at the beginning of each month. This generates a useful implicit lag structure in the model, where the average of daily bond yields is assumed to be affected by the forecasts, and not vice versa.[9]

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[7] An alternative, also used by Gómez-Puig (2006), would have been to use bid-ask spreads. Their advantage is the availability at higher frequencies, and their immediate comparability across markets, whereas nominal amounts have to be converted into the same currency. While Gómez-Puig (2006) has shown that both proxies are similar, we strictly prefer the outstanding amounts for our sample, given that bid-ask spreads during the sovereign debt crisis have grown tremendously, which suggests that they are not an objective proxy for the liquidity of a market, but might be endogenous to an increased pricing of liquidity risk.

[8] See http://www.consensuseconomics.com/.

[9] Such an implicit lag structure has also been used in Leduc et al. (2007) to impose identification restrictions in a VAR. Note that there is no implicit lag structure for the other variables in the model.

One important feature of these forecasts is their variable forecast horizon. With the exception of interest rate forecasts, the respondents are asked to provide expectations over the current and the next calendar year. This implies that over the course of a year, the forecast horizon shrinks. Of course, this is not desirable for our purposes, as we would expect a fixed forecast horizon to be most relevant for the pricing of sovereign bonds of a fixed maturity. Following Dovern et al. (2012), we therefore construct such fixed-horizon forecasts by constructing a weighted average of the two forecasts provided, thus yielding forecasts for one-year ahead. To give one example – for forecasts provided in October of a given year, we approximate the one-year ahead forecast by weighting the current-year forecast by 3/12, and the next-year forecast by 9/12, respectively.

Of the various forecasts that are available, we decided to focus on consumer price inflation (% change p.a.), real GDP growth (% change p.a.), unemployment (% of labour force), the current account balance (nominal values), and most importantly the budget balance for the fiscal year (nominal values). With this choice, we try to capture the various dimensions that might be at play – real as well as nominal developments, the labour market, the international performance of a country, and the fiscal position. For robustness, we have also incorporated interest rate forecasts, but did not find that these mattered or changed our results. For parsimony, we decided not to include these in the estimation.

We use the raw forecasts for inflation, real GDP growth and unemployment. The forecasts of inflation and real GDP growth are also used to construct forecasts of nominal GDP,[10] which are obtained by multiplying the nominal GDP, available in a certain quarter, by the CPI inflation and the real GDP growth rate forecasts. For example, suppose that in June we want to compute the one-year ahead nominal GDP forecast, then we multiply the value of nominal GDP, available in June, by the one-year ahead inflation forecast and the one-year ahead real GDP growth rate forecast. Predictions for the next two months, July and August, are then computed in a similar way, by multiplying the nominal GDP in June by the new inflation and real GDP growth rate forecasts available in July and August respectively. In the first month of each quarter the

[10] Note that the dataset only contains consumer price inflation expectations, but not those for the GDP deflator. Accordingly, our calculation of the expected nominal growth rates has to be seen as an approximation of the actual expectations.

nominal GDP is then updated with the new value available for that quarter and the computation is repeated in a similar fashion. The forecast of the nominal GDP allows us to generate an expectation of the current account to GDP ratio for each country, as well as for the budget balance to GDP ratio.

Our preferred measure for the fiscal position is the debt to GDP ratio, which we calculated based on the prevailing debt levels and the expected budget balances; Paesani et al. (2006) show that the accumulation of government debt affects long-term interest rates. For a measure of the existing debt levels, we used those for the central or federal government debt, to be as close as possible to the definition of the government bond yields, which are similarly reflecting the price of central or federal government debt.

In a nutshell, using the Consensus Economics forecast data, we are able to obtain monthly one-year ahead expectations for the debt to GDP ratio, the current account balance relative to GDP, real GDP growth, unemployment and consumer price inflation. An important caveat with these data deserves mentioning at this point. Obviously, the forecast horizon of one year is relatively short compared to the 10-year maturity of the government bonds we consider. However, this might not be as critical as it looks at first sight. First, it is a well-known fact that forecasts become considerably more uncertain, the further the forecast horizon. Accordingly, the information content in shorter-term forecasts might be superior to the one contained in forecasts with very long horizons; also, the forecasts would certainly converge to their unconditional means after few years, therefore movements of the 10-year maturity are very likely to be related to changes in short-run expectations. Additionally, we would assume that the vast majority of market participants does not hold the bonds to maturity, which makes the expectations for the nearer-term future relatively more important. Still, to test for robustness, we will repeat our econometric analysis using shorter (namely 5-year) maturity bond yields.

In our empirical application, we will test to what extent spreads in a given country are affected by the various determinants. We will run estimations separately for each country, as we are not interested in the average effects, but instead would like to understand better the pricing pattern for each spread individually. Furthermore, we believe that cross-country equality constraints are largely implausible, such that a panel analysis would not be warranted.

Our approach is close to that of D'Agostino, Gambetti and Giannone (2013); they use time-varying coefficient VARs with stochastic volatility to forecast in real time, out-of-sample, inflation, the unemployment rate and the interest rate in the United States. They show that allowing for time variation in the coefficients is crucial for the forecasting accuracy of the model. In this paper we use a similar approach; the estimation technique based on a univariate regression equation with time-varying coefficients and stochastic volatility is very well suited to describe yield spread dynamics.

Let us assume that the endogenous variable is denoted by y_t, while the regressors are denoted by a vector of variables $X_t = [x_{1,t}, ..., x_{k,t}]'$. We assume that y_t admits the following representation:

$$y_t = \theta_{0,t} + \theta_{1,t} x_{1,t} + ... + \theta_{k,t} x_{k,t} + \varepsilon_t \qquad (1)$$

where $\theta_{0,t}$ is a time-varying intercept, $\theta_{i,t}$ are time-varying coefficients, $i = 1, ..., k$ and ε_t is a Gaussian white noise with zero mean and time-varying variance σ_t^2. Let $\theta_t \equiv [\theta_{0,t}, \theta_{1,t}, ..., \theta_{k,t}]'$. The time-varying parameters are postulated to evolve according to:

$$\theta_t = \theta_{t-1} + \omega_t \qquad (2)$$

where ω_t is a Gaussian white noise with zero mean and variance-covariance matrix Ω.

We assume that the standard deviation of ε_t, σ_t, belongs to the class of models known as stochastic volatility and evolves as a geometric random walk:

$$\log \sigma_t = \log \sigma_{t-1} + \xi_t \qquad (3)$$

where ξ_t is Gaussian white noise with zero mean and variance ζ. We assume also that ξ_t, ω_t and ε_t are mutually uncorrelated at all leads and lags.

The model (1)-(3) is estimated using Bayesian methods. A detailed description of the algorithm, including the Markov-Chain Monte Carlo (MCMC) used to simulate the posterior distribution of the hyper-parameters and the states conditional on the data, is provided in the Appendix.

It is worth emphasising that the algorithm used in this paper allows us to compute error bands around the median estimates of the coefficients, thereby providing a very natural way to assess their statistical significance.

As for the specification of the priors, we assume that the priors for the initial states, θ_0, of the time varying coefficients and log standard errors, log σ_τ, are normally distributed. The prior for the hyper-parameters, Ω, is assumed to be distributed as an Inverse Wishart (IW), while the distribution of ζ is assumed to be an Inverse Gamma (IG). More specifically, we have the following priors:

- Time-varying coefficients: $P(\theta_0) \sim N(\hat{\theta}, \widehat{V_\theta})$ and $P(\Omega) \sim IW(\Omega_0^{-1}, \rho)$;
- Stochastic volatility: $P(log\sigma_0) \sim N(log\hat{\sigma}, 1)$;
- $P(\zeta) \sim IG(\zeta_0^{-1}, \upsilon_0)$.

The hyper-parameters are calibrated using a time-invariant OLS regression estimated over the entire sample (T_0). The degrees of freedom for the covariance matrix of the innovations to the drifting coefficients, ρ, are set equal to T_0. The degrees of freedom ν_0, for the prior on the stochastic volatility variance ζ, are set equal to 0.001, while the prior d_0, in the scale matrix ζ_0^{-1}, is set equal to 1. The matrix $\Omega_0^{-1} = \lambda\widehat{V_\theta}$, with λ, the parameter governing the amount of time-variation in the unobserved states, is fixed, in each regression, to track the optimal percentage of residuals outside the confidence band for a given percentile.[11]

An important difference relative to previous studies in the literature is that we will not enter the determinants in relative terms, wherever possible. To give an example, if we are to model the French-German yield spread, we will not include relative debt to GDP ratios, i.e. $\frac{Debt}{GDP}_{FR,t} - \frac{Debt}{GDP}_{GE,t}$. Rather, we will include $\frac{\mathrm{Debt}}{\mathrm{GDP}}_{\mathrm{FR,t}}$ and $\frac{\mathrm{Debt}}{\mathrm{GDP}}_{\mathrm{GE,t}}$ as separate variables.[12] Econometrically, this implies that we do not impose the restriction that the coefficient on French and German variables is identical (with the opposite sign), without actually testing it. Economically, this implies that we allow the price impact of a change

[11] Primiceri (2005) shows that the choice of λ has a substantial impact on the estimation results, hence some words on our approach are in order. Our choice of λ is based on the in-sample accuracy of the fit. Very loose values of λ would imply a large variance of the distribution of the coefficients, and hence a large variance of the distribution of the fitted values. In this case the model would tend to overfit the data, and an overly large percentage of observed data would lie within the confidence bands around the fitted values. The opposite would happen if λ is very tight. Ideally, we would like to observe 1% of the observed data to lie outside the 1% confidence bands, 2% to lie outside the 2% confidence bands and so on. We fix λ as the value that minimizes the distance from the actual percentages from their theoretically expected values.

[12] We have tested all variables for stationarity, and found that for the spreads, the debt to GDP ratios, the current account to GDP ratios and the liquidity measures, the presence of a unit root cannot be excluded, while GDP growth, CPI inflation and the VIX are stationary. Importantly, however, the residual diagnostics show that residuals in the regressions are stationary.

in the French fiscal position to differ from the price impact of a change in the German fiscal position. This is immediately intuitive – for instance, if we believe both the United States and Germany to be safe havens, we would expect that their own macroeconomic fundamentals are considerably less important than the positions of the other countries. A restriction of the coefficient is therefore highly implausible. As we will see, it is indeed empirically not justified, and masks interesting differences in the relative price impact of the respective national variables.[13] Also, we expect the variables of the benchmark country to matter relatively more if the two countries are closer to being substitutes for international investors, which not only speaks against imposing equality restriction for a given country pair, but furthermore against imposing equality restrictions across country pairs by estimating the model in a panel framework.

Using the fundamental variables separately for both countries could potentially create a multicollinearity problem in the estimations, as some variables show a high correlation between the countries. We have therefore compared our results across 1) the unrestricted model reported in this paper, 2) specifications where we only included the variables of a given country, i.e. without including the variable of the benchmark country, 3) specifications where we only included the variable for the benchmark country, and 4) a model with differenced variables. We found that the coefficients in the unrestricted specification are very similar to those found in models 2) and 3), and that results with differenced variables are typically determined by the most significant individual variable. These results therefore make us confident that our approach does not suffer from multicollinearity problems.

The only exception to the principle of using unrestricted country-level variables rather than differences is the liquidity risk proxy, where we enter the size of a given bond market relative to the one of the benchmark country. This is due to the fact that our proxy is measured in nominal terms, and as such is not a meaningful regressor in the model – only the ratio of the outstanding amounts of public debt in a given country relative to the amounts outstanding for the benchmark country constitute a variable that can appropriately proxy the liquidity risk of a given bond.

[13] Some studies do not include the determinants of the benchmark country at all (e.g. Nickel et al. 2011), presumably based on the assumption that the benchmark yield constitutes a risk-free rate that is entirely exogenous to country-specific determinants. This also imposes an untested restriction on the econometric model, namely a zero-restriction – which, as we shall see below, often is not warranted, either.

To get a first impression of the relevance of this assumption, and to identify which variables might be particularly important in determining spreads, we ran an initial estimation using simple OLS and neglecting the time-variation of parameters. For ease of illustration, we ran the regression in a panel framework (with spreads defined relative to the United States), allowing for country-fixed effects. Table 2 provides the results, in panel (1) using the restricted approach where all variables are entered in relative terms, and in panel (2) leaving all variables unrestricted. The relative size of markets as measured by the liquidity ratio, does not seem to matter much on average (where our hypothesis would have been that larger ratios lower the spreads). The proxies for general risk aversion should show up with a positive sign if more risk aversion increases spreads. Here, the VIX clearly dominates the Baa-Aaa spread. The estimated parameters for the macroeconomic fundamentals are typically in line with the expectations: we would assume that spreads increase with higher relative debt, unemployment or inflation, and decrease with higher relative current account balances and GDP growth. With the exception of inflation, these hypotheses are confirmed when looking at panel (1). When we relax the parameter restriction and re-estimate the model in panel (2) with the individual variables, it is clearly evident that the parameters for the domestic variables and those for the United States are not equal in magnitude and of opposite signs. The last columns labelled "p-value" puts this hypothesis to a statistical test, and clearly shows that the restrictions are typically rejected by the data. The parameters often are of a very different magnitude (e.g. on unemployment, where an increase in the U.S. unemployment rate lowers spreads by much more than an increase in the domestic unemployment rate would increase spreads). At times, only one of the two variables exerts a statistically significant effect – for instance, the evolution of domestic debt and domestic inflation matters, whereas the corresponding U.S. figures are not estimated to be statistically significant.

Table 2 around here

Table 2 also allows us to select the most important variables in our subsequent models. These will not be estimated in a panel, and will include time-varying parameters, such that the specification of a parsimonious model is advisable. Based on the finding that the Baa-Aaa spread is clearly dominated by the VIX as a measure of general risk aversion, in what follows we will no longer include the Baa-Aaa spread.[14] Furthermore, we will also exclude the unemployment variables, which are capturing the business cycle in a similar fashion to GDP growth. Sensitivity tests show that our results are robust to the exclusion of these variables.

Having specified the data and the econometric model, we will now move on to discuss the econometric results.

4. The time-varying determinants of sovereign bond yield spreads

Italy

As mentioned above, the country which saw its yields co-move the least with those of the other G7 countries was Italy. In the early years of the sample, i.e. in the uprun to EMU, its yields were considerably higher. Subsequently, under EMU, yields converged, whereas during the sovereign debt crisis, Italian yields again substantially exceeded those of the other G7 countries. This suggests that the Italian case could be particularly interesting to study, and as such should give us a feeling for whether or not the econometric models provide us with reasonable results. We will therefore first focus on discussing the Italian results, which are provided in Figure 2. Spreads are defined relative to Germany. The bold solid lines in the figure show the time-varying parameter estimates, the posterior median values, together with the 68% posterior error bands (the dotted lines), associated with the distribution of the parameters. The dashed line shows the OLS estimate.

Figure 2 around here

The first observation that emerges from looking at Figure 2 is that the assumption of parameter constancy is clearly rejected. For a number of variables, there are substantial time variations that are both large economically and statistically significant. Furthermore, time variation is present for all three risk factors, general risk aversion, liquidity risk and credit risk. Let us take these three in turns.

[14] The VIX and the Baa-Aaa spread are likely correlated leading to a multicollinearity problem in the estimations. When entering each of these variables separately, we still find the VIX to be the more relevant variable, as judged by R^2, AIC as well as BIC criteria.

Liquidity risk clearly affects Italian spreads towards the end of our sample: the larger is the amount of outstanding debt of Italian relative to German bonds, the smaller the spread. However, this was not always the case – for most of the sample, there is no significant pricing of this risk. Only starting in early 2008, liquidity risk is reflected in the Italian spreads, with increasing magnitudes. At the end of the sample, we estimate an effect in the order of 30 basis points for each 10% change in the liquidity ratio.

A similar, albeit much stronger, picture emerges for general risk aversion as proxied by the VIX. When risk aversion is high, spreads are relatively large. However, there is a clear U-shape in the parameter estimates, with large and significant impacts at the beginning of the sample, small and largely insignificant effects in the early years of EMU, and a steep increase that starts in 2009, i.e. during the global financial crisis, and continues until the end of our sample. This is in line with the hypothesis that there has been very little pricing of risk in the uprun to the global financial crisis. The coefficient stands at around 0.03 at the end of the sample, which implies a 24 basis point increase in spreads following a one-standard deviation in risk aversion. Obviously, the increase in the VIX during the financial crisis was much larger than that – it rose by roughly 15 points, from an average of around 15 in the years prior to the crisis (starting in 2005) to an average of around 30 during the financial crisis. This would imply an increase of Italian spreads by around 45 basis points.

The most striking picture emerges, however, when looking at the effect of the expected Italian debt to GDP ratio. An increase in the ratio by 10% led to an increase in the spread by around 60 basis points at the beginning of the sample, whereas the effect reduced to virtually zero in the early 2000s, and has risen to 100 basis points at the end of the sample.

If we compare the coefficients on Italian debt to those on German debt, it becomes apparent that the imposition of an equality restriction is by no means justified. Higher debt in Germany lowers spreads, as expected, but there is far less time variation than for Italian debt, magnitudes are considerably smaller, and throughout the entire sample period, the effect is not statistically significant.

As to the other macroeconomic fundamentals, the effect of the Italian current account balance to GDP ratio is as expected – a larger balance reduces the spread, especially in the early part of the sample. Also here, the German position matters much less, with typically substantially smaller and mostly insignificant coefficients. Expected GDP growth is largely unimportant, for both countries. Expected Italian inflation used to affect spreads prior to monetary union, with higher inflation leading to rising spreads, but that effect disappeared with monetary union, and has not re-emerged since. German expected inflation was unimportant throughout the sample.

Another important issue to note is that the model performs in a highly satisfactory manner. The dotted grey line in Figure 2 provides the results from a simple OLS model, and shows that these are often statistically considerably different from those stemming from the model that explicitly takes into account the time variation. Furthermore, as we will see in Section 5, the residuals generated by our econometric model are considerably less persistent than those from a simple OLS model.

To summarise the results, it is apparent that in the time period prior to the introduction of the euro, a large number of (primarily Italian) macroeconomic fundamentals mattered for determining Italian spreads, whereas in the early years of monetary union, this effect entirely disappeared. With the financial crisis, liquidity risk and general risk aversion started to be priced substantially more, but in terms of macroeconomic fundamentals, only the evolution of expected Italian debt started to reappear as an important determinant. These results clearly corroborate the hypothesis that there was only very little pricing of risk in the early years of this century, whereas there is much more of it currently. Another key finding from these results is the fact that the macroeconomic fundamentals of Italy and Germany always mattered very disproportionately for the Italian-German spreads, with Italian fundamentals being much more important than their German counterparts.

France

Let us turn next to the French-German spreads (provided in Figure 3), given that these are easily comparable to the Italian spreads just discussed. Very similar findings result. The coefficients for the pricing of liquidity risk show an inverse U-shape, whereas those related to the VIX and to French debt are U-shaped. Like in the Italian example, this suggests that there has been a pricing of liquidity risk, general risk aversion and macroeconomic fundamentals prior to monetary union, which disappeared in the first years after the introduction of the euro, only to re-emerge with the financial crisis. While the pattern is similar, the overall magnitudes are substantially smaller – they are roughly half the magnitude of the coefficients estimated for Italy with regard to the liquidity ratio and debt, and around a third with regard to the VIX.

Figure 3 around here

Another similarity to the Italian example is that macroeconomic fundamentals other than French debt matter(ed) very little, and if so, mainly in the early years of our sample. An interesting difference emerges with regard to the effect of German debt – here, we find that an increase in German debt lowered French spreads towards the end of the sample. Even though the coefficient on German debt is still considerably smaller than the one of French debt (-2 compared to 5 at the end of the sample), this finding might already suggest that there is a more even impact of both countries' fundamentals if the rating of both government bonds is similar, thus making them substitutes in the eyes of potential investors. We will get back to this issue when discussing the British-German and the German-U.S. spreads.

United Kingdom

Results for the British-German spread are reported in Figure 4. The differences with the Italian and French spreads are striking. While there is still a U-shaped pattern for the VIX, the coefficients for the British debt show very little time variation, and are statistically significant throughout the sample, suggesting that the time variation we saw for Italian and French spreads was most likely a euro area-specific phenomenon. Interestingly, the magnitude of the coefficient stands at around 1.5, and is therefore substantially smaller than what we observe for the Italian and French spreads during the sovereign debt crisis.

Similarly to the French spreads, we find significant coefficients for German debt, with the exception of a short period in the early 2000s. The coefficients on our proxy for liquidity risk have the expected sign, and are significant for large parts of the sample. Also here, the magnitude of the coefficient suggests that there is much less pricing of liquidity risk than for the Italian and French spreads during the sovereign debt crisis.

Figure 4 around here

With regard to the other credit risk proxies, we do find some of these to matter, but several coefficients have a counterintuitive sign – for instance, increasing expectations of current account balances in the UK as well as in Germany seem to raise the spreads (whereas we would have expected that only expectations for Germany do so), and the coefficients on GDP growth are both counterintuitive. At the same time, they are small in magnitude.

Canada

When analysing the Canadian spreads (in Figure 5), it is again apparent that the results, and in particular the time variations, look remarkably different from those found for the Italian and French spreads. A first important result is that the signs of the coefficients are generally in line with our hypotheses. With regard to time variations, the figures reveal two interesting periods. The first is located at the end of the 1990s and the beginning of the 2000s, and presumably related to the events surrounding the dot-com bubble in the United States, and the belief that potential output of the U.S. economy had permanently increased due to a productivity growth arising from the IT revolution (Jorgenson, 2001). During this period, GDP growth in the United States increased Canadian spreads.[15] As will be seen later, this seems to be an effect genuinely arising from the pricing of U.S. government bonds, as the same pattern is also found for the German-U.S. and the Japanese-U.S. spreads. Accordingly, we believe that enhanced GDP growth expectations for the United States helped keeping the yields for U.S. government bonds low, rather than increasing those of Canada or of the other countries. In parallel, an improvement in the Canadian current account (much of which is against the United States) helped to lower the spreads.

Figure 5 around here

A second distinct pricing pattern is found for the years 2004-2008. During this period, the liquidity ratio matters (whereas we find a counterintuitive sign for it at the beginning of the sample), and rising debt expectations for Canada increase the spread, as do (counter intuitively) expectations of increasing U.S. inflation. We rationalise this with the completely divergent debt evolution in Canada and the United States over the time

[15] Interestingly, this period also coincides with the time when markets priced a "scarcity premium" on U.S. long-term government bonds based on concerns that these would become increasingly scarce if the U.S. Treasury would pay down its outstanding debt over the coming decade (Reinhart and Sack 2002).

period prior to the financial crisis, when U.S. federal debt was strongly increasing while Canada's debt showed a moderate decline. As a consequence, the liquidity ratio was on a declining trend, and we conjecture that at some point the increasing gap in the size of the two markets led to an increasing pricing of liquidity risk. This diverging trend was abruptly stopped during the financial crisis, when also Canadian debt started to rise again.

Japan

The pricing pattern of Japanese spreads (shown in Figure 6) is particularly interesting. As mentioned above, like in Canada also here expectations of U.S. GDP growth were raising spreads during the times of the dot-com bubble, whereas expectations of an improved Japanese current account lowered spreads during this period. There are several distinctive features, however, that are worth mentioning. As is well known, the bulk of the Japanese debt is held nationally, such that a number of variables might not be capturing relevant determinants of the yield spreads. As a matter of fact, the effect of the liquidity ratio is insignificantly estimated through nearly the entire sample period. There is also very little effect of Japanese debt on spreads, whereas the coefficient on U.S. debt is large in magnitude and strongly statistically significant. The most intriguing difference compared to all previous results is the finding that increasing inflation expectations in Japan actually *lowered* the spread, consistently so until 2003. This might well be the case in a scenario of deflation, where increasing (decreasing) inflation might actually be good (bad) news to investors. As a matter of fact, inflation expectations were declining in our sample until late 2002, and started rising thereafter.

Figure 6 around here

Germany

The last spread to be studied is the one of German relative to U.S. bonds. Results are depicted in Figure 7, with several interesting findings. First, the effect of U.S. GDP growth expectations increasing the spread during the dot-com bubble emerges also here. Second, and more interesting, the pricing of debt is now highly symmetric – German debt increases the German spread significantly throughout the sample, and U.S. debt decreases the spread, also in a significant fashion. Interestingly, the magnitudes of the parameters are comparable, in the range of 2 (suggesting that a 10% change in debt to

GDP ratios changes the spread by around 20 basis points). This speaks in favour of a symmetric pricing pattern in case government bonds are perceived as reasonably close substitutes by market participants. The third finding is that an increase in general risk aversion, as measured by the VIX, actually lowers spreads, which could come about if times of increasing general risk aversion, Germany's safe haven status improves in relative terms.

Figure 7 around here

To corroborate this finding, it is interesting to see the VIX coefficient estimates when all country spreads are estimated relative to Germany. As we will see in our robustness tests, doing so provides a very clear picture – for all spreads relative to Germany, the coefficient estimates for the VIX are U-shaped, with smaller coefficients during the late 1990s and early 2000s, but strongly increasing coefficients towards the end of the sample, in particular during the financial crisis.

To summarise these results, we find that macroeconomic fundamentals matter in very distinct ways, with those of the benchmark country typically being less relevant. A symmetric pricing pattern is found for the German-U.S. spread, however, suggesting that if bonds are reasonably close substitutes, or can both be considered as safe havens, the underlying fundamentals of both countries start to be of roughly equal importance. This implies that the analysis of spreads with relative variables, where the restriction of symmetry is imposed, is particularly problematic for the spreads of higher-yielding bonds relative to low-yielders.

A second key finding is that time variations are important, as shown for instance by the increasing importance of U.S. GDP growth expectations during the dot-com bubble, but even more importantly by the U-shaped relationship that is found for several risk factors in euro area spreads, suggesting that these risks had been priced prior to monetary union and during the financial crisis, but not in between.

Robustness

We have conducted several robustness checks to investigate the sensitivity of our results to our modelling choices. Given the large amount of material, Table 3 synthesises the results by portraying, for each country and for each variable of the model, the posterior median values of the parameter estimates at the beginning, in the middle and at the end of the sample period, always along with the 68% posterior error bands (in brackets). The first panel repeats the estimates of the benchmark model.

Table 3 around here

The second panel of Table 3 shows how results change if the denominator is changed from Germany to the United States (for Italian, French and British spreads) and vice versa (for Canadian and Japanese spreads). As mentioned previously, in this case we can detect the U-shaped pattern for the coefficient on the VIX that exists for the French, British and Italian spreads relative to Germany also for the spreads of Canada and Japan, suggesting that in times of increasing general risk aversion, German yields are particularly low, thus increasing the spreads of all other countries (note that this even holds against the United States).

A second robustness test, reported in the third panel of Table 3, consists in substituting inflation expectations with those of unemployment. Overall, results are very robust, with few changes in significance or magnitude of coefficients. The estimates for unemployment are also interesting – while there are counterintuitive findings for the Canadian-U.S. spread, it turns out that expectations of higher domestic (foreign) unemployment increase (decrease) the spread, as one should expect, for Italy, France, the United Kingdom, Japan and Germany.

A final test is portrayed in the last panel of Table 3, namely a repetition of the benchmark model estimates for 5-year yield spreads. As can be seen, results are very robust.

5. Testing under-pricing and over-pricing of risk

As mentioned in the introduction to this paper, a question at the very heart of the current policy debate is the extent to which we have seen an under-pricing of risk prior to the global financial crisis, and an over-pricing during the sovereign debt crisis. Providing evidence for these hypotheses is intrinsically difficult. In the context of our time-varying parameter models, it is not sufficient to point to the large swings in the pricing of risk for the euro area countries Italy and France. While such swings go in the direction of these hypotheses (i.e. there has been very little pricing of risk in the first years of monetary union, and there is a lot more at the end of the sample), they are not sufficient – small coefficients in the early years of EMU could have been in line with fundamentals, e.g. if there are non-linearities in credit risks, and if fundamentals have been in a region where reasonable changes in their expected values would not have triggered a re-assessment of credit risk and therefore its price.

To get at this question, we have added non-linearities to the model, such as the squared value of the expected debt to GDP ratio, or interaction terms of the VIX or of the expected debt to GDP ratio with macroeconomic fundamentals. None of these turned out to be important (results not shown for brevity), suggesting either that non-linearities are not relevant in the pricing of risk, at least not during our sample, or alternatively that these are implicitly picked up by the time-variations in our parameter estimates.

Of course, due to its time-varying parameters, our model is extremely flexible, and allows that the pricing of risk differs substantially over time. One possible thought experiment could therefore be to see to what extent *actual* pricing of risk falls outside the bands that are predicted as plausible by such a flexible model, conditional on the evolution of fundamentals. Figure 8 therefore plots the actual spreads (depicted by the solid black lines) against the 68% posterior bands for the fitted values of our model (shown by the grey shaded areas). A number of interesting results emerge from this chart.

First, the overall fit of our models is extremely good. While we would of course expect that 32% of all observations fall outside the grey shaded bands, we note that the magnitude of the deviations is overall very small, or in other words that the residuals from our models are small in magnitude. Second, there are basically only two periods when our models generate large residuals, the first being the period of the scarcity premium on U.S. bonds around the turn of the millennium (as discussed above). This affects all spreads against the USA, which are effectively larger than what is estimated by the models.

The second period is the sovereign debt crisis, where we find the models for all countries but Italy and France to do rather well, whereas we observe massive residuals for these two countries. For Italy, at the very end of the sample, the model estimates that spreads against Germany should be in the order of 2%, whereas actual spreads are 4.5%, i.e. more than double.[16] The gap is also large for France, with an estimated spread of 0.8%, and an actual spread of 1.2% to 1.5%.

Figure 8 around here

How to interpret these results? First, the charts show no evidence of an under-pricing of risk prior to the global financial crisis – if it was there, our models are sufficiently flexible to incorporate this. Second, conditional on fundamentals, and even allowing for substantially elevated pricing of risk during the sovereign debt crisis, our models do a very poor job in explaining the actual spreads of Italy and France at this time. That our models cannot capture this development might either relate to the speed of the increase, which might have been too high to be captured by our time-varying parameters, or alternatively suggest a severe mis-specification of our models, which have omitted some risk factors that get priced currently in the euro area. Given that we rejected the possibility of non-linear terms, we can only conclude that markets are pricing risks that are not modelled here, such as a re-denomination risk for euro area bonds (see also Draghi 2012).

What would a time-invariant OLS regression model have estimated? This question is answered by the dotted line in Figure 8. It is immediately apparent that this model generates much larger residuals, which are furthermore much more persistent. Of course, this is to be expected, given that it is much less flexible in fitting the data. A crucial difference to our model, however, is that the OLS estimates are much more sensitive to the estimation sample. While the time-varying parameter model might give rather conservative estimates of possible over- or under-pricing of risks, its results are more robust over time.

A similar thought experiment is to fix the parameter estimates not at their OLS values, but at the level estimated by the time-varying model at a conveniently chosen point in time. Given the hypotheses that there has been under-pricing and over-pricing of risk

[16] In a recent speech (Annual Meeting of the Italian Banking Association, July 2012) the Governor of Bank of Italy, Ignazio Visco, stated that "The difference between the yields on Italian and German government securities is far greater than could be justified by our economy's fundamentals. It reflects general fears of the monetary union breaking up – a remote possibility but one that is nevertheless influencing the choices made by international investors." See also Di Cesare et al. (2012).

over the recent years, we should try to find a point in time that has been unperturbed by any crises and that has not been a suspect of mis-pricing of risk. One such candidate could be the time around 1994, i.e. after the ERM crisis of 1992/1993, but prior to the uprun to EMU. Figure 9 provides the results of this counterfactual exercise. The grey shaded areas and the black solid line are as in Figure 8, whereas the area indicated by the vertical bars provides the 68% posterior error bands of the counterfactual simulation, with parameters fixed at their June 1994 values.[17] Results are intriguing. We identify the same two periods where the actual spreads deviate substantially and persistently from the estimated spreads. First, the time of the scarcity premium on U.S. bonds, and second, the sovereign debt crisis for Italy and France. However, there is now also a third period – it turns out that relative to the pricing patterns observed in 1994, there is a persistent and non-negligible under-pricing of risk in the early 2000s for all spreads, with the only exception of the German-U.S. spread. We identify this pattern (where the vertical bars lie above the black solid line) for 2002-2004 for Italy, for 2004-2007 for France and the UK, for 2002-2008 for Canada, and for 2002-2006 for Japan.

We are very well aware that these counterfactual simulations and attempts to discover a mis-pricing of risk are fraught with caveats. However, in view of the fact that our model estimates have to be seen as rather conservative (since, due to their time-varying nature, they already incorporate swings in the pricing of risk), we take the fact that they still point to an under-pricing of risk in much of the first decade of this millennium and that they cannot explain the large spreads of Italy and France at the very end of the sample as indicative.

6. Conclusions

Against the background of the current debate about fiscal sustainability in several advanced economies, and the recent history of rating downgrades of countries with long-standing excellent credit ratings, this paper has estimated the determinants of sovereign bond spreads of the G7 countries, using time-varying parameter stochastic volatility models. This is in contrast to the bulk of the existing literature, which has

[17] The counterfactual exercise is performed by computing the fitted values, over the entire sample, conditional on the parameters estimated for June 1994.

typically focused on either emerging market economies or euro area (candidate) countries.

Beyond controlling for exchange rate risk and analysing liquidity risk and general risk aversion, the paper has studied the role of macroeconomic fundamentals in determining yield spreads. In order to do so, it has used high frequency expectations of financial institutions, with the advantage that these data should be close to market participants' views, and that they do not require an interpolation from lower frequencies.

A major difference compared to previous studies has been that this paper allows for asymmetric effects of countries' fundamentals on yield spreads by entering the fundamentals of both countries defining the spread separately. It turns out that this innovation leads to a much better understanding of the determinants of bond yield spreads.

The key findings of this paper are as follows. First, for a spread of any country relative to a safe haven government bond (such as the U.S. or German bonds), the countries' macroeconomic fundamentals are bound to be considerably more influential determinants of the spread than the fundamentals of the benchmark country. The closer the two bonds are to being substitutable, the more symmetric is the impact of the respective fundamentals. Second, there are considerable time variations in the role of the various determinants. For instance, during the dot-com bubble, expectations of U.S. GDP growth lowered U.S. yields, whereas no such effect is found for the other time periods. Similarly, we find that several risk factors have not been priced in the years preceding the financial crisis. This pattern is particularly pronounced for the determinants of the Italian-German and the French-German spreads, i.e. for spreads of the euro area member countries, where macro fundamentals, general risk aversion and liquidity risks used to be priced in the uprun to monetary union and following the outbreak of the financial crisis, but not in the first years of monetary union.

Running counterfactual exercises where we fix the pricing patterns observed in 1994, we identify three periods where actual spreads deviated substantially and persistently from those estimated by our model: the time of the scarcity premium on U.S. bonds (where actual spreads were larger than estimated), the first decade of the millennium (where spreads were lower than suggested by the model), and the sovereign debt crisis (where Italian and French spreads were substantially larger than our model would have predicted). These findings support the belief that swings in risk appetite have led to an under-pricing of risk prior to the global financial crisis, and either an over-pricing of risk during the European sovereign debt crisis or the pricing of a risk that government bonds of some euro area countries might get re-denominated in other currencies than the euro.

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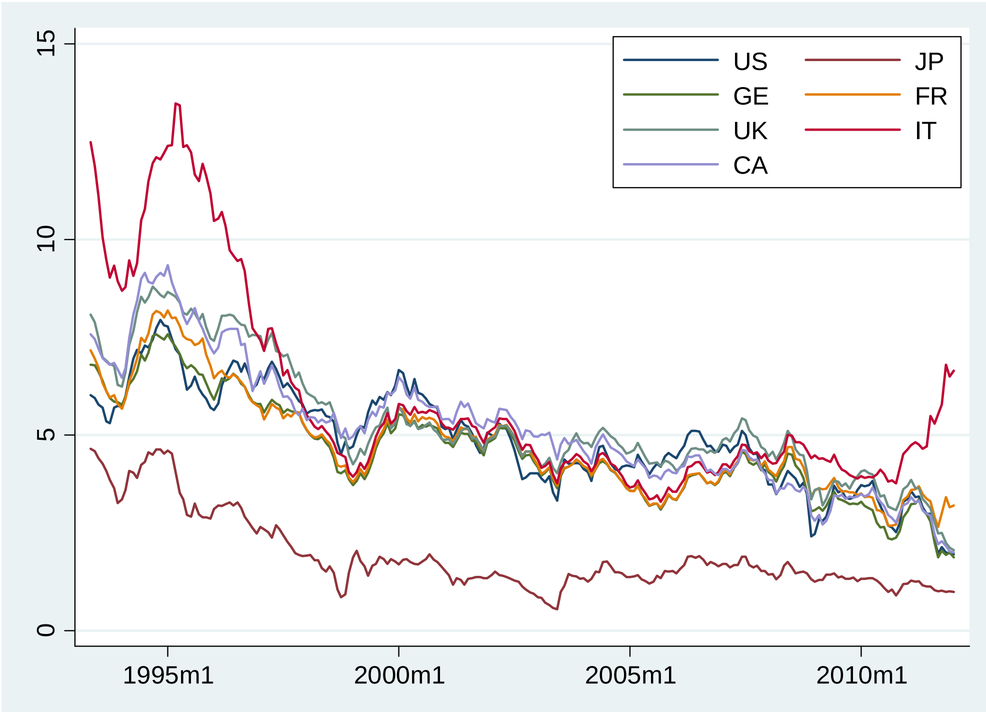
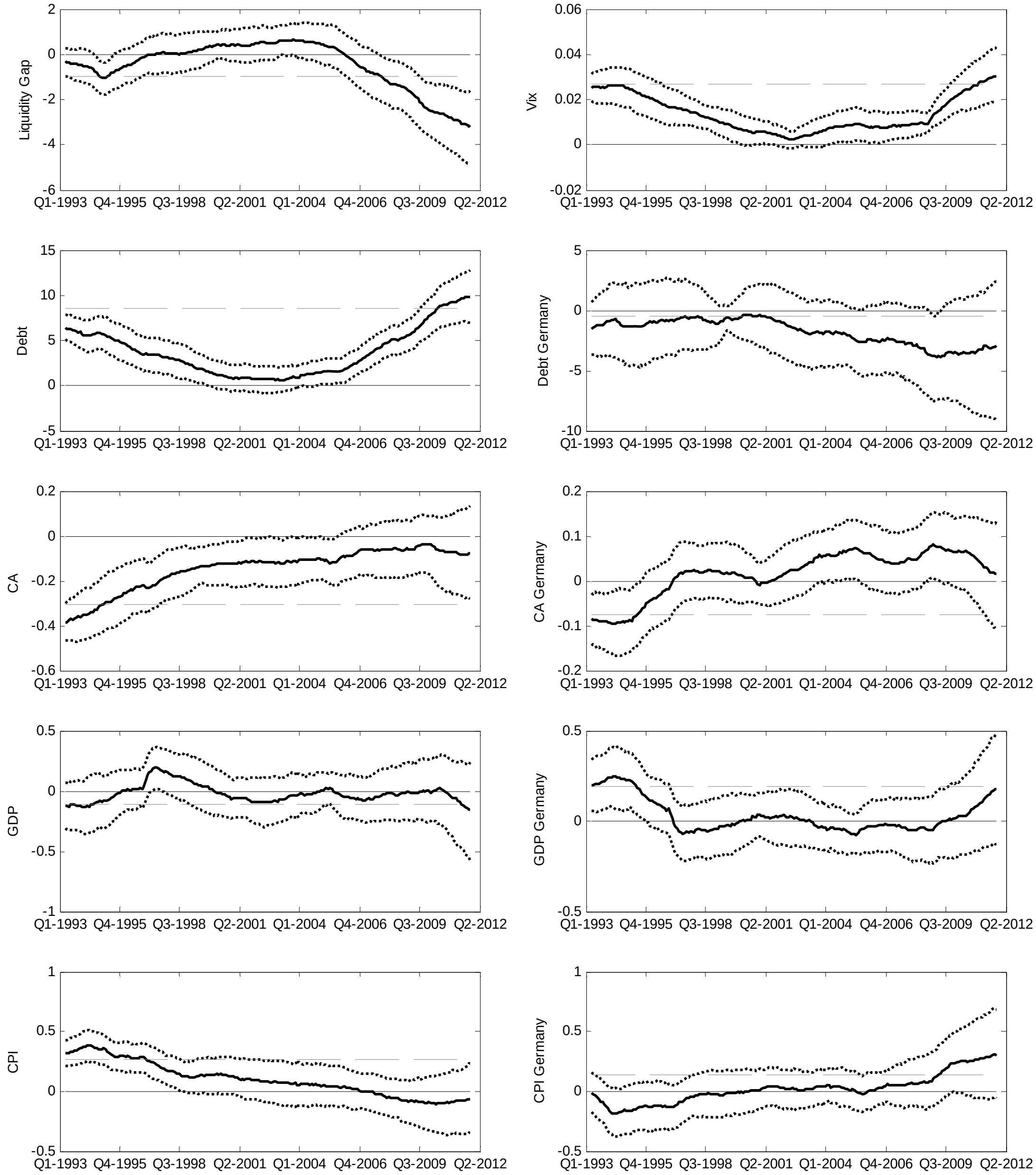
Figure 1: Long-term government bond yields



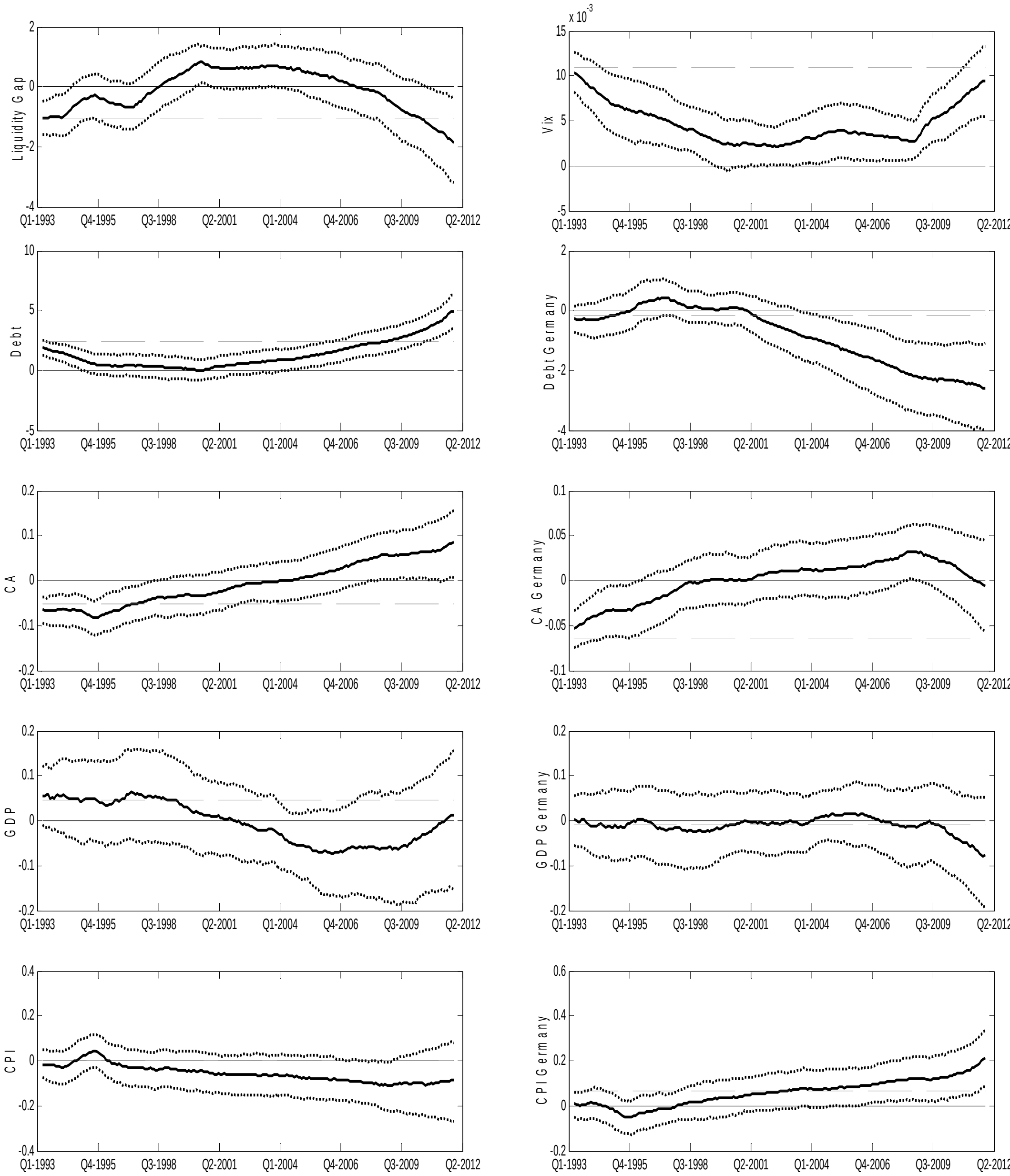
Note: The chart shows the evolution of long-term government bond yields. Data are in per cent.

Figure 2: Determinants of bond yield spreads, Italy versus Germany



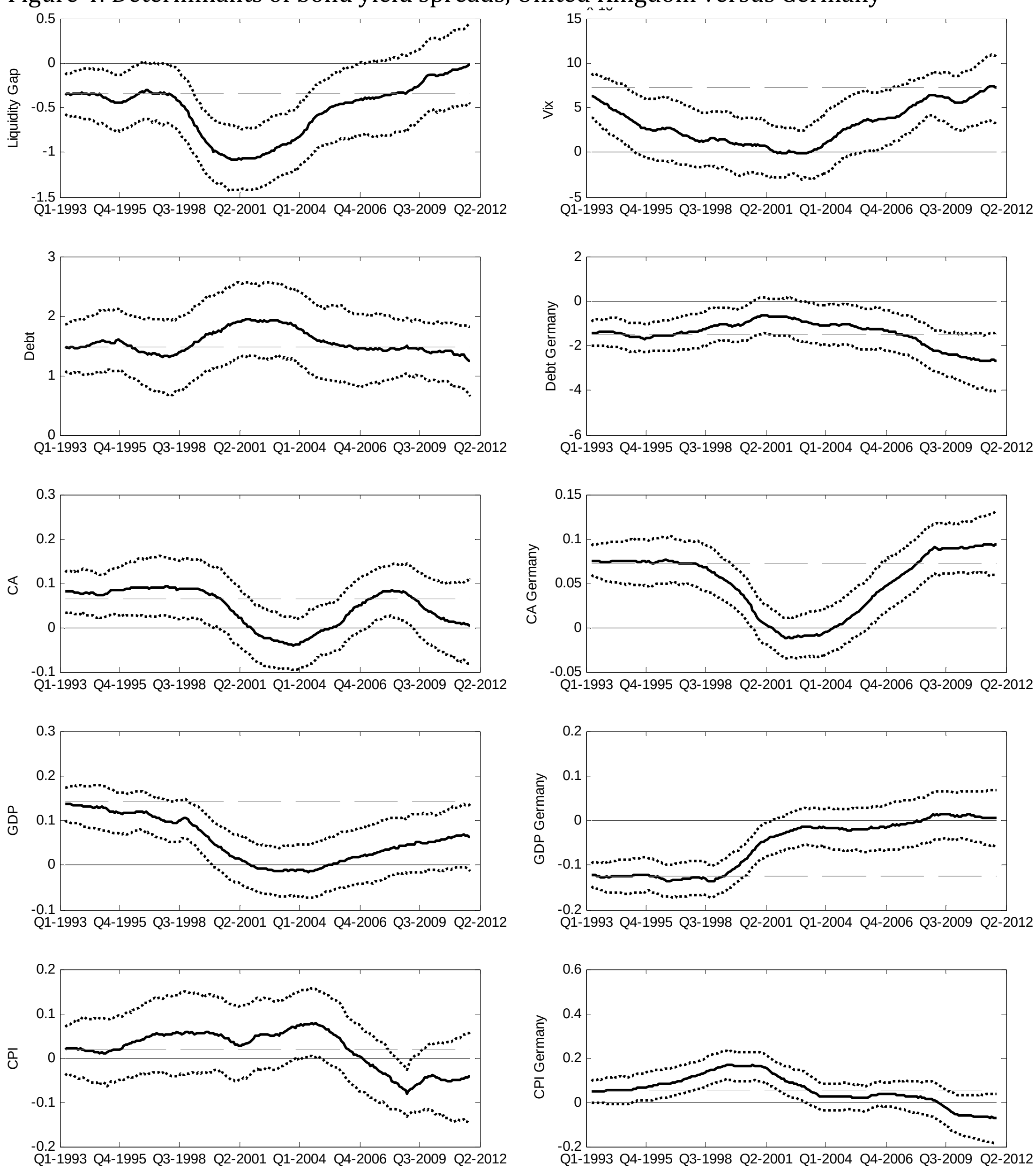
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of Italian and German bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 3: Determinants of bond yield spreads, France versus Germany



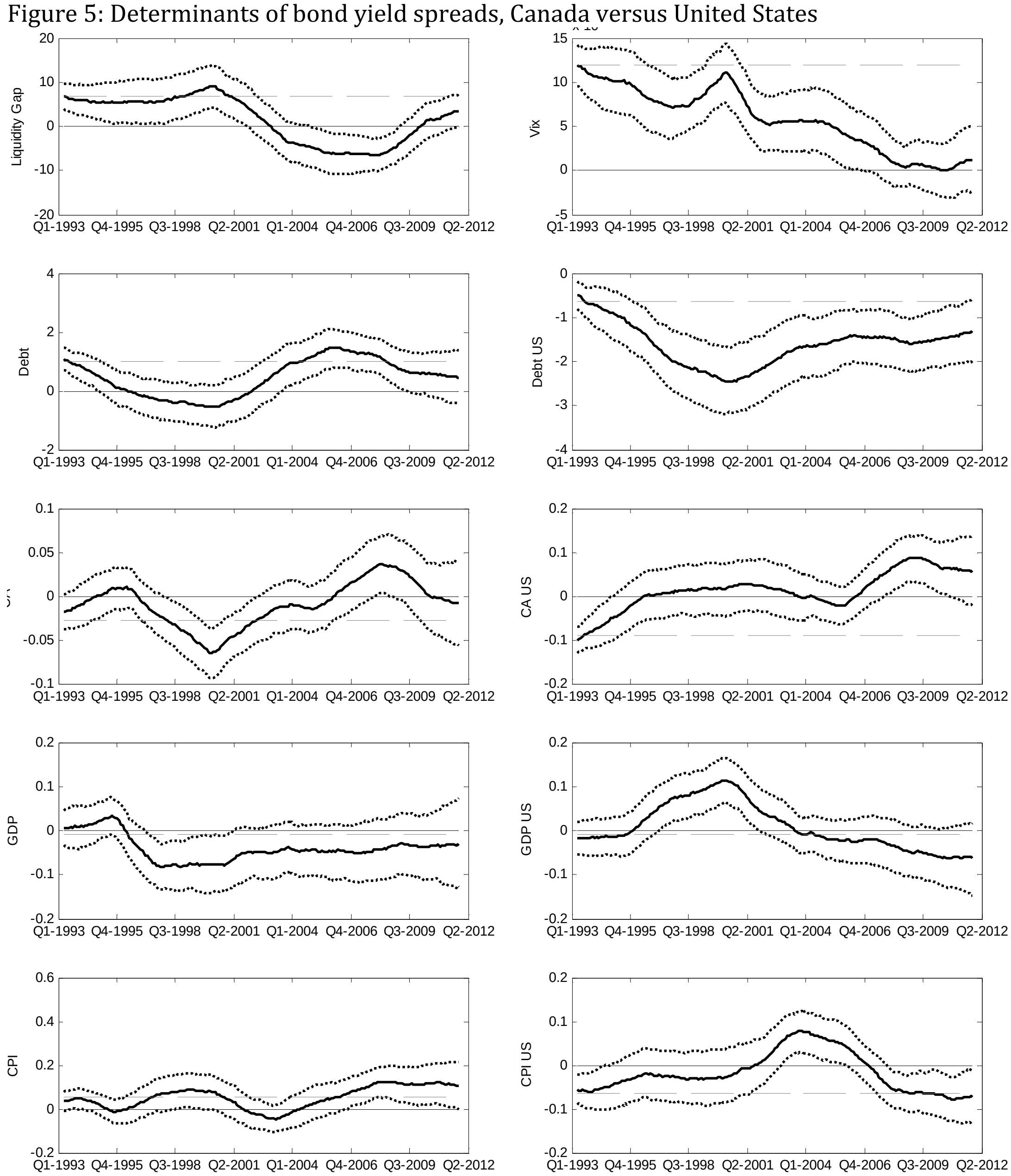
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of French and German bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 4: Determinants of bond yield spreads, United Kingdom versus Germany



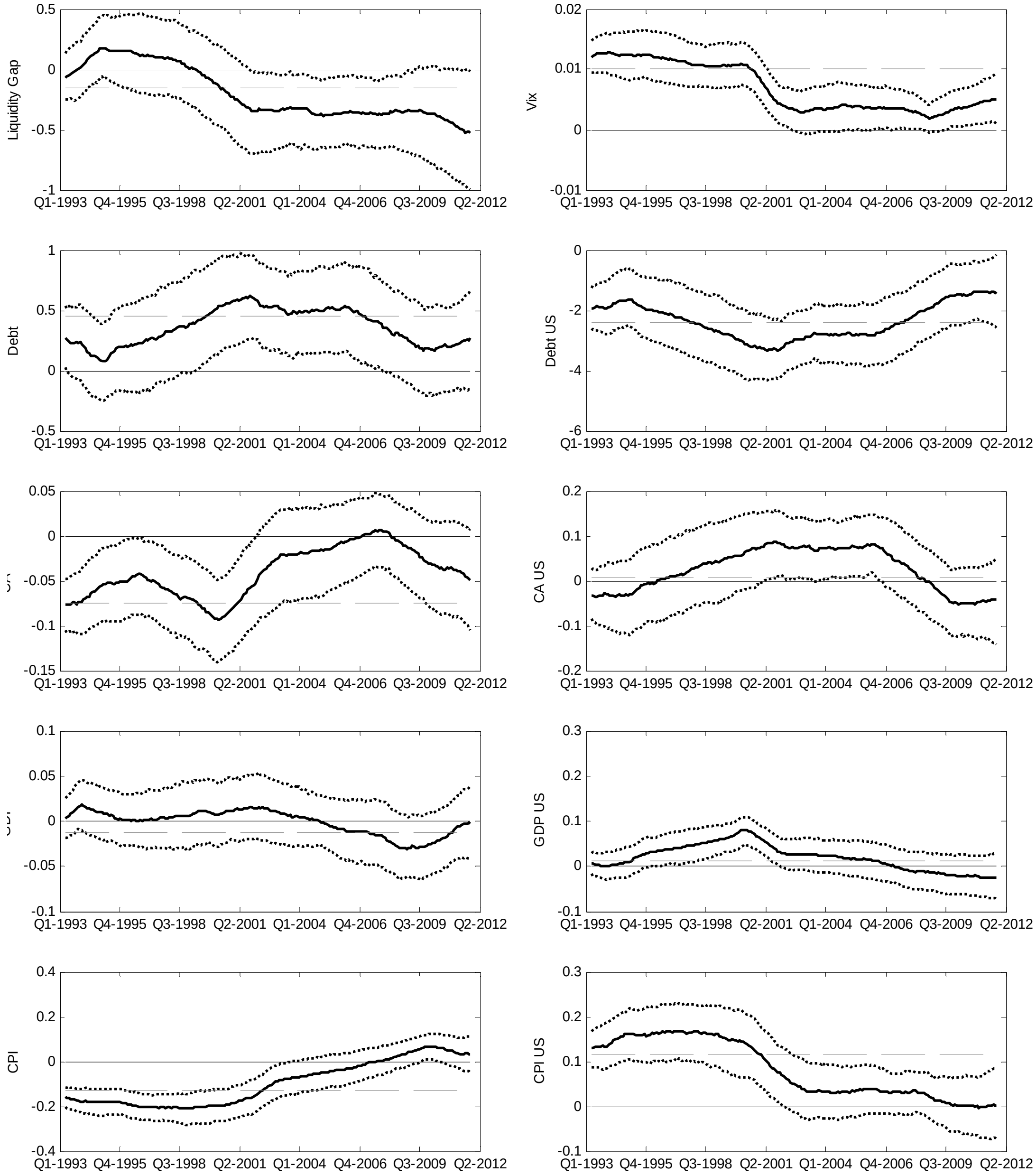
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of British and German bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 5: Determinants of bond yield spreads, Canada versus United States



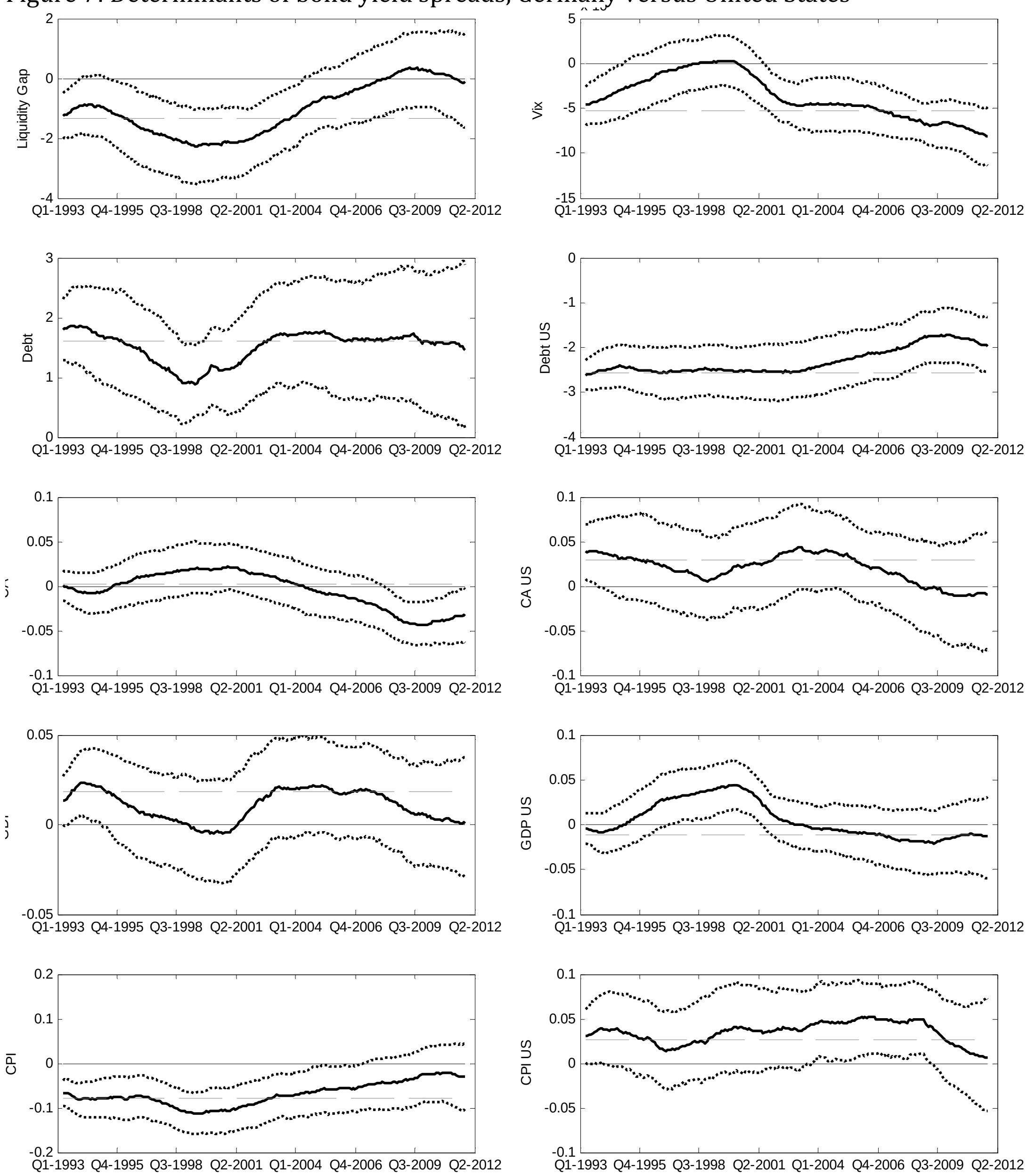
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of Canadian and U.S. bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 6: Determinants of bond yield spreads, Japan versus United States



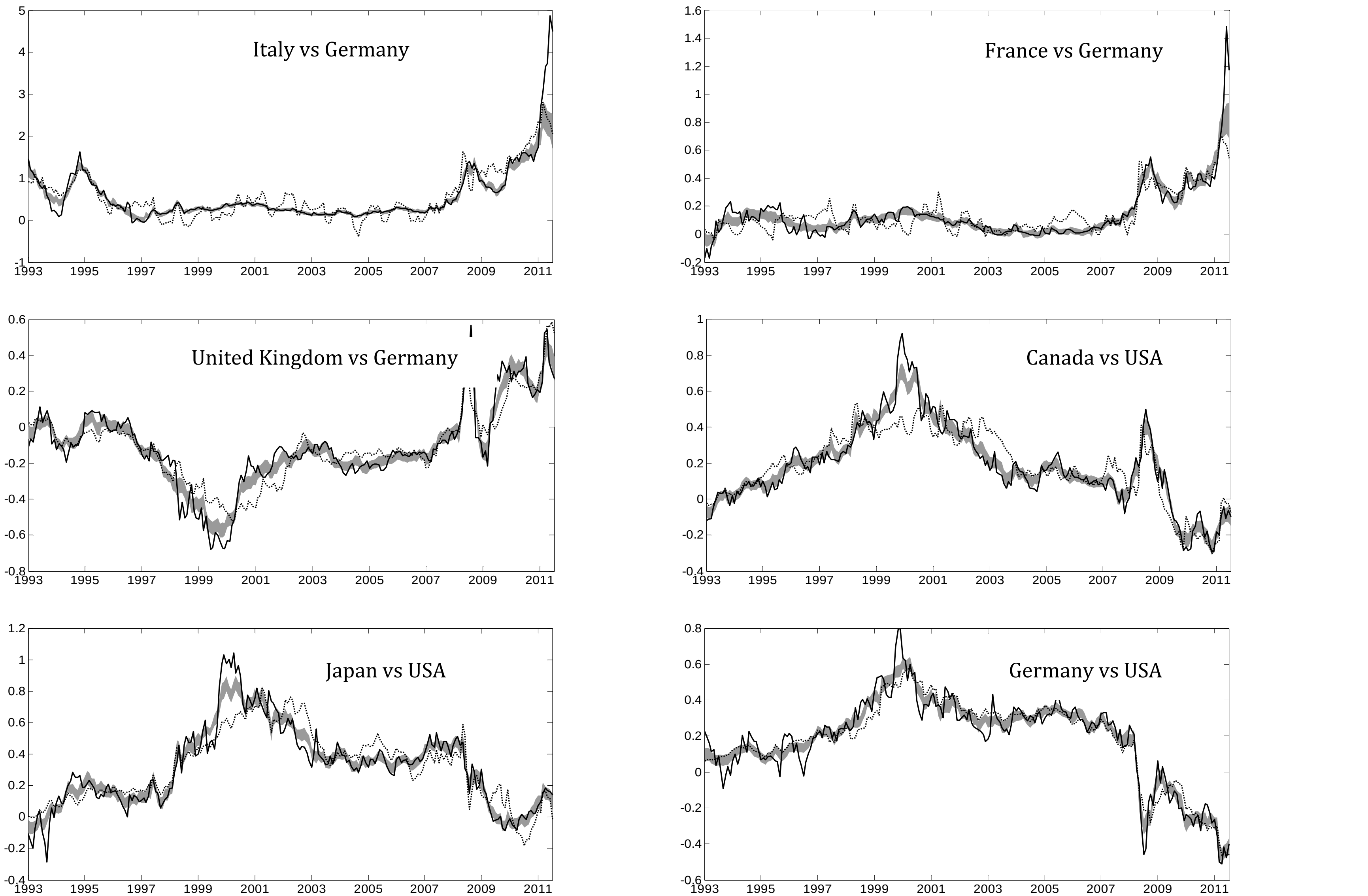
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of Japanese and U.S. bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 7: Determinants of bond yield spreads, Germany versus United States



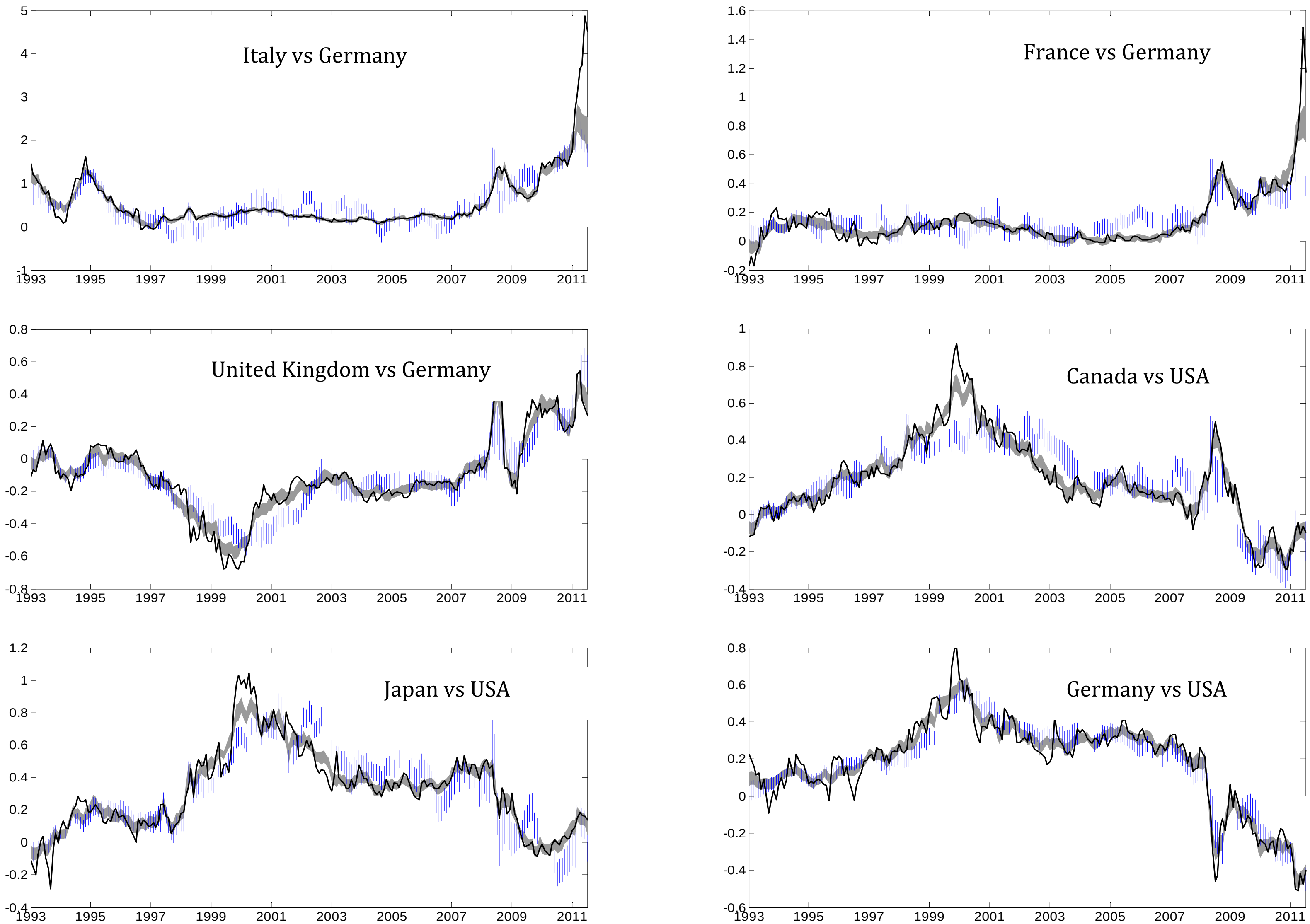
Note: The charts show posterior median values for the time-varying parameter estimates of model (1)-(3), for the spread of German and U.S. bond yields. Dotted lines are 68% posterior error bands. The dashed line shows the OLS estimate.

Figure 8: Actual and fitted bond yield spreads



Note: The charts show the actual spreads (black solid line), the fitted spreads from a time-invariant OLS regression (black dotted line) and the 68% posterior error bands of the fitted spreads obtained from the Bayesian time-varying parameter models (grey shaded area).

Figure 9: Actual and fitted bond yield spreads, counterfactual simulation



Note: The charts show the actual spreads (black solid line), the 68% posterior error bands of the fitted spreads obtained from a counterfactual simulation, keeping the parameters fixed at their values in June 1994 (blue bars) and the 68% posterior error bands of the fitted spreads obtained from the Bayesian time-varying parameter models (grey shaded area).

Table 1: Country and stock market index coverage

		Obs	Mean	Std. Dev.	Min	Max
Yields (%)	Canada	224	5.193	1.627	2.021	9.348
	France	224	4.779	1.304	2.649	8.193
	Germany	224	4.599	1.282	1.870	7.576
	Italy	224	5.892	2.571	3.290	13.483
	Japan	224	1.910	0.967	0.542	4.646
	UK	224	5.320	1.538	2.116	8.801
	US	224	4.905	1.294	1.964	7.945
Spread to US (%)	Canada	224	0.289	0.638	-0.666	1.912
	France	224	-0.125	0.634	-1.420	1.441
	Germany	224	-0.305	0.507	-1.553	0.845
	Italy	224	0.987	1.891	-1.278	6.462
	Japan	224	-2.995	0.827	-4.962	-0.962
	UK	224	0.415	0.655	-1.098	2.058
Spread to Germany (%)	Canada	224	0.594	0.501	-0.789	2.254
	France	224	0.138	0.188	-0.231	1.486
	Italy	224	0.572	0.690	-0.059	4.877
	Japan	224	-2.690	0.694	-3.882	-0.867
	UK	224	0.721	0.470	-0.115	1.755
	US	224	0.305	0.507	-0.845	1.553
General risk aversion	Baa Aaa spread	224	0.960	0.458	0.548	3.385
	VIX	224	20.979	8.414	10.820	62.640
Liquidity risk: outstanding amounts of public debt (mio US$)	Canada	224	542,089	173,850	362,340	1,243,348
	France	224	775,036	369,314	270,742	1,681,307
	Germany	224	862,094	394,606	383,575	1,777,815
	Italy	224	1,130,295	407,614	578,615	2,149,013
	Japan	224	3,827,369	2,145,632	1,249,183	8,934,450
	UK	224	587,143	282,581	263,033	1,280,726
	US	224	4,214,912	1,742,323	2,799,536	9,543,056
Credit risk: expected debt to GDP ratio	Canada	224	0.437	0.122	0.273	0.619
	France	224	0.594	0.103	0.381	0.828
	Germany	224	0.317	0.080	0.188	0.450
	Italy	224	1.027	0.058	0.915	1.132
	Japan	224	1.203	0.477	0.483	2.009
	UK	224	0.474	0.111	0.381	0.855
	US	224	0.640	0.107	0.527	0.954
Credit risk: expected current account to GDP ratio	Canada	224	-0.215	1.781	-3.228	2.724
	France	224	0.281	1.560	-2.403	2.338
	Germany	224	1.870	2.540	-1.137	6.802
	Italy	224	-0.187	1.964	-3.855	3.346
	Japan	224	2.857	0.836	1.208	4.849
	UK	224	-1.706	0.805	-3.777	0.047
	US	224	-3.665	1.433	-6.407	-1.106
Credit risk: expected real GDP growth	Canada	224	2.684	0.918	-1.224	3.888
	France	224	1.866	0.960	-1.781	3.592
	Germany	224	1.555	1.130	-3.274	3.018
	Italy	224	1.424	1.049	-2.763	2.956
	Japan	224	1.102	1.273	-4.675	3.136
	UK	224	2.087	1.078	-2.446	3.388
	US	224	2.658	1.087	-2.063	4.619
Credit risk: expected unemployment	Canada	224	7.889	1.297	6.053	10.975
	France	224	10.121	1.467	7.356	12.689
	Germany	224	9.669	1.193	6.695	11.471
	Italy	224	9.489	1.808	5.905	12.136
	Japan	224	4.329	0.875	2.547	5.868
	UK	224	4.868	2.117	2.684	10.803
	US	224	5.931	1.597	4.041	9.988
Credit risk: expected consumer price inflation	Canada	224	1.882	0.429	0.435	2.773
	France	224	1.619	0.433	0.502	2.600
	Germany	224	1.721	0.604	0.532	3.701
	Italy	224	2.408	0.946	0.975	5.118
	Japan	224	0.066	0.617	-1.096	1.386
	UK	224	2.677	0.653	0.443	4.357
	US	224	2.415	0.704	-0.662	3.562

Note: The table shows summary statistics for the variables used in the econometric model.

Table 2: Preliminary spread analysis in a panel OLS setting

Panel A: macro determinants in relative terms		coeff		*std. error*	p-value
Liquidity		0.133		*0.112*	
Baa-Aaa spread		-0.029		*0.031*	
VIX		0.007	***	*0.002*	
Expected debt to GDP ratio	Relative	0.114	*	*0.063*	
Expected current account to GDP ratio	Relative	-0.009	**	*0.004*	
Expected real GDP growth	Relative	-0.032	***	*0.011*	
Expected unemployment	Relative	0.020	***	*0.006*	
Expected consumer price inflation	Relative	-0.019		*0.017*	
Observations		1344			
Adjusted R^2		0.392			

Panel B: macro determinants separately		coeff		*std. error*	p-value
Liquidity		-0.103		*0.099*	
Baa-Aaa spread		-0.015		*0.037*	
VIX		0.010	***	*0.002*	
Expected debt to GDP ratio	National	0.139	**	*0.057*	0.187
	US	0.296		*0.353*	
Expected current account to GDP ratio	National	-0.043	***	*0.005*	**0.000**
	US	-0.044	***	*0.008*	
Expected real GDP growth	National	-0.024		*0.016*	0.670
	US	0.030	**	*0.014*	
Expected unemployment	National	0.032	***	*0.005*	**0.028**
	US	-0.069	***	*0.016*	
Expected consumer price inflation	National	0.058	***	*0.019*	**0.005**
	US	-0.005		*0.021*	
Observations		1344			
Adjusted R^2		0.443			

Note: The table shows results for a time-constant model using pooled data (with spreads defined relative to the United States), allowing for country-fixed effects and estimated using simple OLS. The model is formulated as $y_{i,t} = \theta_{0,i} + \theta_1 x_{1,i,t} + \ldots + \theta_k x_{k,i,t} + \varepsilon_{i,t}$. In Panel A, the credit risk proxies are defined relative to the United States; in Panel B, they enter separately for the national economy and the United States. Figures in the column "p-value" report the p-values of a test for the equality of national and US coefficients (albeit with different signs).

Table 3: Robustness tests

Country	Variable	(1) Benchmark			(2) Alternative denominator country			(3) Adding unemployment			(4) 5-year maturity		
		Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End
Italy	Liquidity gap	-0.32	0.51	-3.16	-1.20	-1.34	-2.26	-0.85	-0.15	-3.71	-0.64	0.35	-3.45
		(-0.96 0.25)	(-0.22 1.19)	(-4.84 -1.62)	(-3.22 0.84)	(-3.43 0.92)	(-7.15 2.61)	(-1.51 -0.19)	(-0.83 0.52)	(-5.24 -2.03)	(-1.39 0.00)	(-0.47 1.23)	(-5.39 -1.61)
	$Vix*10^{-2}$	2.51	0.34	3.05	1.66	0.14	2.97	1.13	0.13	2.54	2.57	0.21	3.04
		(1.93 3.15)	(-0.07 0.81)	(1.87 4.31)	(0.94 2.52)	(-0.49 0.77)	(1.32 4.6)	(0.34 1.82)	(-0.45 0.64)	(1.06 3.99)	(1.85 3.35)	(-0.3 0.78)	(1.15 4.52)
	Debt	6.41	0.73	9.86	4.48	1.07	0.85	3.44	0.66	5.18	6.09	0.67	8.13
		(5.08 7.86)	(-0.80 2.14)	(6.96 12.74)	(2.08 7.07)	(-0.72 2.88)	(-4.12 6.4)	(1.36 5.77)	(-1.03 2.35)	(0.59 9.66)	(4.51 7.83)	(-1.02 2.40)	(4.53 11.60)
	Debt denominator country	-1.51	-1.59	-2.95	-0.34	0.35	5.86	-4.37	-0.57	-1.03	-1.42	-0.53	-5.01
		(-3.66 0.72)	(-4.55 1.00)	(-9.07 2.44)	(-1.73 1.03)	(-1.07 1.8)	(2.67 9.06)	(-5.75 -3.20)	(-2.79 1.60)	(-4.89 2.52)	(-4.34 1.29)	(-3.72 2.97)	(-11.84 1.80)
	Current account	-0.38	-0.12	-0.08	-0.28	-0.11	0.19	-0.36	-0.06	-0.04	-0.34	-0.08	-0.11
		(-0.46 -0.30)	(-0.23 -0.01)	(-0.27 0.14)	(-0.38 0.18)	(-0.19 -0.01)	(0.00 0.42)	(-0.43 -0.28)	(-0.16 0.05)	(-0.22 0.16)	(-0.44 -0.25)	(-0.20 0.04)	(-0.34 0.13)
	Current account denominator country	-0.09	0.03	0.02	0.04	0.12	-0.40	-0.04	0.06	-0.13	-0.08	0.00	-0.01
		(-0.14 -0.03)	(-0.03 0.09)	(-0.11 0.13)	(-0.06 0.13)	(0.01 0.25)	(-0.59 -0.2)	(-0.09 0.02)	(-0.01 0.12)	(-0.25 -0.01)	(-0.15 -0.02)	(-0.07 0.08)	(-0.14 0.12)
	GDP growth	-0.12	-0.08	-0.16	0.17	0.04	-0.13	-0.07	-0.06	-0.30	-0.12	-0.10	-0.26
		(-0.32 0.07)	(-0.27 0.12)	(-0.58 0.24)	(0.08 0.26)	(-0.08 0.16)	(-0.3 0.07)	(-0.24 0.11)	(-0.24 0.13)	(-0.70 0.07)	(-0.34 0.11)	(-0.33 0.14)	(-0.73 0.22)
	GDP growth denominator country	0.20	0.01	0.18	-0.04	0.02	0.03	0.08	-0.01	0.25	0.22	0.01	0.29
		(0.06 0.34)	(-0.14 0.17)	(-0.13 0.48)	(-0.11 0.04)	(-0.05 0.09)	(-0.13 0.19)	(-0.04 0.20)	(-0.16 0.13)	(-0.01 0.55)	(0.06 0.39)	(-0.17 0.19)	(-0.08 0.65)
	CPI inflation	0.32	0.08	-0.07	0.19	-0.10	0.50	--	--	--	0.22	-0.03	-0.15
		(0.21 0.42)	(-0.10 0.25)	(-0.35 0.23)	(0.08 0.3)	(-0.25 0.05)	(0.26 0.76)				(0.09 0.34)	(-0.24 0.17)	(-0.49 0.20)
	CPI inflation denominator country	-0.01	0.01	0.30	0.08	0.08	0.04	--	--	--	0.18	0.18	0.37
		(-0.17 0.15)	(-0.14 0.18)	(-0.06 0.68)	(-0.03 0.19)	(-0.04 0.2)	(-0.21 0.3)				-(0.01 0.36)	(-0.03 0.38)	(-0.08 0.79)
	Unemployment	--	--	--	--	--	--	0.17	0.11	0.10	--	--	--
								(0.04 0.29)	(0.00 0.21)	(-0.15 0.36)			
	Unemployment denominator country	--	--	--	--	--	--	-0.22	-0.24	-0.25	--	--	--
								(-0.28 -0.15)	(-0.32 -0.15)	(-0.38 -0.11)			
France	Liquidity gap	-1.05	0.64	-1.85	0.53	-2.24	2.30	-0.65	0.44	-1.70	-0.82	0.69	-1.55
		(-1.59 -0.50)	(-0.03 1.32)	(-3.18 -0.34)	(-0.49 1.56)	(-3.70 -0.65)	(-0.19 4.86)	(-1.08 -0.16)	(-0.16 1.01)	(-2.92 -0.60)	(-1.30 -0.35)	(0.04 1.35)	(-2.80 -0.29)
	$Vix*10^{-2}$	1.04	0.24	0.95	0.44	0.05	0.39	0.56	0.13	0.76	0.41	-0.08	0.42
		(0.82 1.26)	(0.02 0.48)	(0.54 1.032)	(0.17 0.68)	(-0.24 0.35)	(-0.14 0.92)	(0.34 0.76)	(-0.09 0.37)	(0.38 1.17)	(0.23 0.61)	(-0.3 0.16)	(0.05 0.78)
	Debt	1.89	0.66	4.93	-1.11	1.16	-1.08	1.33	0.76	3.90	1.42	-0.02	3.06
		(1.30 2.51)	(-0.27 1.59)	(3.49 6.28)	(-1.67 -0.48)	(0.04 2.29)	(-2.51 0.76)	(0.85 1.82)	(-0.14 1.47)	(2.38 5.33)	(0.93 1.92)	(-0.88 0.86)	(1.74 4.34)
	Debt denominator country	-0.28	-0.63	-2.59	1.48	-1.94	2.38	0.15	-0.56	-2.49	0.14	0.22	-0.83
		(-0.74 0.15)	(-1.39 0.14)	(-3.98 -1.10)	(0.67 2.24)	(-3.18 -0.83)	(0.60 3.99)	(-0.44 0.75)	(-1.49 0.44)	(-4.38 -0.78)	(-0.23 0.54)	(-0.44 0.84)	(-2.10 0.51)
	Current account	-0.07	-0.01	0.08	0.08	-0.04	0.17	0.01	0.00	0.00	0.02	0.00	0.09
		(-0.10 -0.04)	(-0.04 0.03)	(0.01 0.16)	(0.04 0.10)	(-0.07 0.00)	(0.10 0.23)	(-0.02 0.03)	(-0.03 0.02)	(-0.06 0.06)	(-0.01 0.04)	(-0.04 0.04)	(0.02 0.16)
	Current account denominator country	-0.05	0.01	-0.01	-0.11	0.00	-0.09	-0.01	0.00	-0.05	0.01	0.01	0.01
		(-0.07 -0.03)	(-0.02 0.04)	(-0.06 0.04)	(-0.14 -0.07)	(-0.04 0.05)	(-0.16 -0.01)	(-0.03 0.01)	(-0.02 0.03)	(-0.09 -0.01)	(-0.01 0.03)	(-0.02 0.04)	(-0.03 0.05)
	GDP growth	0.06	-0.01	0.01	0.02	0.04	-0.12	0.11	0.02	0.06	0.11	-0.01	0.04
		(-0.01 0.12)	(-0.09 0.06)	(-0.15 0.16)	(0.00 0.05)	(-0.01 0.09)	(-0.20 -0.04)	(0.06 0.15)	(-0.05 0.09)	(-0.09 0.21)	(0.06 0.16)	(-0.09 0.06)	(-0.11 0.20)
	GDP growth denominator country	0.00	0.00	-0.08	-0.02	0.02	-0.01	-0.08	-0.04	-0.08	-0.06	-0.02	-0.08
		(-0.06 0.05)	(-0.07 0.06)	(-0.19 0.05)	(-0.04 0.01)	(-0.02 0.05)	(-0.08 0.05)	(-0.12 -0.04)	(-0.09 0.02)	(-0.20 0.04)	(-0.11 -0.02)	(-0.08 0.05)	(-0.19 0.04)
	CPI inflation	-0.02	-0.06	-0.09	-0.22	-0.18	-0.20	--	--	--	-0.04	-0.01	-0.07
		(-0.08 0.05)	(-0.15 0.03)	(-0.27 0.08)	(-0.28 -0.15)	(-0.28 -0.07)	(-0.34 -0.08)				(-0.09 0.01)	(-0.10 0.08)	(-0.21 0.08)
	CPI inflation denominator country	0.01	0.07	0.21	0.14	0.01	0.26	--	--	--	0.14	0.11	0.26
		(-0.05 0.06)	(-0.01 0.14)	(0.08 0.33)	(0.08 0.19)	(-0.05 0.08)	(0.16 0.34)				(0.09 0.18)	(0.03 0.18)	(0.15 0.37)
	Unemployment	--	--	--	--	--	--	0.05	0.02	-0.04	--	--	--
								(0.03 0.08)	(-0.02 0.05)	(-0.11 0.02)			
	Unemployment denominator country	--	--	--	--	--	--	-0.10	-0.08	-0.05	--	--	--
								(-0.13 -0.08)	(-0.11 -0.04)	(-0.11 -0.00)			

Table 3 (continued): Robustness tests

Country	Variable	(1) Benchmark			(2) Alternative denominator country			(3) Adding unemployment			(4) 5-year maturity		
		Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End
United Kingdom	Liquidity gap	-0.36	-0.98	-0.01	-1.11	-1.85	0.82	-0.15	-0.92	0.54	-3.30	-3.15	-3.43
		(-0.57 -0.13)	(-1.31 -0.59)	(-0.46 0.46)	(-1.97 -0.28)	(-3.31 -0.44)	(-1.47 3.04)	(-0.42 0.10)	(-1.30 -0.53)	(-1.04 -0.11)	(-4.14 -2.31)	(-4.66 -1.66)	(-5.61 -1.07)
	Vix*10^{-2}	0.63	-0.01	0.73	-0.12	-0.21	0.43	0.42	-0.13	1.02	0.68	0.17	0.76
		(0.39 0.88)	(-0.28 0.25)	(0.31 1.12)	(-0.31 0.07)	(-0.45 0.07)	(0.04 0.78)	(0.19 0.66)	(-0.41 0.19)	(0.62 1.42)	(0.49 0.87)	(-0.07 0.42)	(0.35 1.17)
	Debt	1.47	1.94	1.24	-0.33	0.11	-0.65	1.37	1.73	0.22	-0.93	0.64	-0.49
		(1.07 1.85)	(1.30 2.54)	(0.66 1.82)	(-0.73 0.11)	(-0.66 0.83)	(-1.62 0.24)	(0.95 1.72)	(1.02 2.53)	(-0.37 0.93)	(-1.36 -0.43)	(-0.07 1.40)	(-1.61 0.61)
	Debt denominator country	-1.41	-0.85	-2.66	0.10	-0.38	0.42	-2.37	-1.58	0.85	0.89	-0.92	-0.45
		(-2.00 -0.90)	(-1.79 0.00)	(-4.06 -1.40)	(-0.41 0.60)	(-1.26 0.48)	(-0.76 1.68)	(-3.00 -1.73)	(-2.65 -0.42)	(-0.63 2.35)	(0.29 1.42)	(-1.72 -0.06)	(-1.83 0.91)
	Current account	0.08	-0.03	0.00	-0.01	-0.04	-0.03	0.08	0.01	0.08	0.04	-0.04	-0.02
		(0.03 0.13)	(-0.09 0.03)	(-0.09 0.11)	(-0.03 0.01)	(-0.08 0.01)	(-0.09 0.03)	(0.04 0.11)	(-0.06 0.07)	(0.00 0.18)	(0.02 0.06)	(-0.09 -0.00)	(-0.08 0.05)
	Current account denominator country	0.08	-0.01	0.09	-0.03	-0.01	0.00	0.08	0.02	0.09	-0.05	0.00	-0.02
		(0.06 0.09)	(-0.03 0.01)	(0.06 0.13)	(-0.04 -0.02)	(-0.03 0.01)	(-0.03 0.04)	(0.07 0.10)	(-0.01 0.04)	(0.05 0.12)	(-0.07 -0.04)	(-0.03 0.02)	(-0.06 0.02)
	GDP growth	0.14	-0.01	0.06	0.04	0.02	0.04	0.17	0.04	0.05	0.07	0.01	0.00
		(0.10 0.17)	(-0.07 0.04)	(-0.01 0.14)	(0.02 0.06)	(-0.03 0.06)	(-0.02 0.10)	(0.13 0.20)	(-0.01 0.10)	(-0.03 0.11)	(0.04 0.09)	(-0.04 0.05)	(-0.06 0.06)
	GDP growth denominator country	-0.12	-0.02	0.01	-0.04	-0.01	0.04	-0.16	-0.05	0.06	-0.04	-0.01	0.02
		(-0.15 -0.10)	(-0.06 0.03)	(-0.06 0.07)	(-0.06 -0.03)	(-0.04 0.01)	(-0.01 0.08)	(-0.20 -0.13)	(-0.10 0.00)	(-0.01 0.13)	(-0.06 -0.02)	(-0.03 0.02)	(-0.04 0.06)
	CPI inflation	0.02	0.05	-0.04	0.03	0.03	-0.02	--	--	--	-0.03	-0.04	0.06
		(-0.03 0.07)	(-0.03 0.13)	(-0.15 0.06)	(-0.00 0.06)	(-0.03 0.08)	(-0.10 0.06)				(-0.07 0.00)	(-0.09 0.01)	(-0.03 0.14)
	CPI inflation denominator country	0.05	0.08	-0.07	0.00	-0.02	-0.02	--	--	--	0.05	0.05	-0.04
		(-0.00 0.09)	(0.01 0.14)	(-0.19 0.04)	(-0.03 0.02)	(-0.07 0.02)	(-0.07 0.04)				(0.02 0.08)	(0.01 0.10)	(-0.10 0.02)
	Unemployment	--	--	--	--	--	--	-0.02	-0.01	0.10	--	--	--
								(-0.04 -0.00)	(-0.05 0.03)	(0.05 0.16)			
	Unemployment denominator country	--	--	--	--	--	--	-0.05	-0.07	0.01	--	--	--
								(-0.07 -0.03)	(-0.11 -0.03)	(-0.04 0.06)			
Canada	Liquidity gap	6.81	0.70	3.50	0.20	0.45	1.00	0.69	-5.28	0.08	7.10	1.03	5.54
		(3.94 9.76)	(-3.73 5.00)	(-0.05 6.94)	(-0.08 0.48)	(0.05 0.86)	(0.27 1.64)	(-2.11 3.78)	(-8.96 -1.26)	(-3.42 4.09)	(2.28 12.27)	(-6.34 7.81)	(-0.81 12.23)
	Vix*10^{-2}	1.19	0.55	0.12	1.05	0.61	1.25	0.66	-0.05	-0.01	1.49	0.94	-0.04
		(0.96 1.43)	(0.22 0.87)	(-0.26 0.51)	(0.85 1.23)	(0.39 0.85)	(0.86 1.69)	(0.38 0.93)	(-0.42 0.03)	(-0.48 0.44)	(1.11 1.86)	(0.36 1.51)	(-0.74 0.69)
	Debt	1.08	0.38	0.43	0.70	0.19	-0.18	1.79	2.21	2.88	0.89	-0.18	-0.53
		(0.70 1.45)	(-0.35 1.12)	(-0.45 1.38)	(0.06 1.36)	(-0.89 1.23)	(-2.02 1.50)	(1.41 2.24)	(1.34 3.03)	(1.78 4.05)	(0.22 1.53)	(-1.37 1.06)	(-2.12 1.11)
	Debt denominator country	-0.49	-1.90	-1.34	1.23	1.15	0.13	-1.75	-2.30	-1.23	0.14	-1.21	-1.05
		(-0.80 -0.19)	(-2.62 -1.16)	(-2.03 -0.57)	(0.44 1.96)	(0.01 2.30)	(-1.79 2.02)	(-2.26 -1.24)	(-3.09 -1.51)	(-2.09 -0.33)	(-0.37 0.65)	(-2.41 0.12)	(-2.39 0.16)
	Current account	-0.02	-0.02	-0.01	-0.02	-0.02	-0.02	-0.02	-0.02	0.04	-0.07	-0.04	-0.06
		(-0.04 0.00)	(-0.05 0.01)	(-0.06 0.04)	(-0.03 -0.01)	(-0.04 -0.00)	(-0.05 0.01)	(-0.04 -0.00)	(-0.04 0.01)	(-0.00 0.08)	(-0.10 -0.03)	(-0.09 0.01)	(-0.13 0.02)
	Current account denominator country	-0.10	0.02	0.06	0.00	-0.01	0.01	-0.02	0.09	0.01	-0.18	-0.03	0.03
		(-0.13 -0.07)	(-0.05 0.07)	(-0.02 0.14)	(-0.02 0.02)	(-0.04 0.02)	(-0.04 0.06)	(-0.05 0.02)	(0.03 0.15)	(-0.07 0.10)	(-0.22 -0.12)	(-0.13 0.06)	(-0.10 0.16)
	GDP growth	0.01	-0.05	-0.03	-0.06	0.00	-0.13	-0.01	-0.10	-0.05	-0.04	-0.02	-0.02
		(-0.04 0.05)	(-0.11 0.01)	(-0.13 0.07)	(-0.09 -0.04)	(-0.04 0.03)	(-0.19 -0.05)	(-0.05 0.02)	(-0.15 -0.06)	(-0.14 0.03)	(-0.11 0.02)	(-0.12 0.09)	(-0.17 0.13)
	GDP growth denominator country	-0.02	0.02	-0.06	0.02	-0.02	0.02	-0.02	0.05	-0.07	-0.04	-0.05	-0.13
		(-0.05 0.02)	(-0.02 0.07)	(-0.15 0.02)	(0.00 0.04)	(-0.05 0.01)	(-0.03 0.07)	(-0.04 0.02)	(0.01 0.09)	(-0.15 0.00)	(-0.10 0.01)	(-0.13 0.03)	(-0.26 0.01)
	CPI inflation	0.04	-0.04	0.11	0.01	0.03	0.04	--	--	--	-0.02	-0.11	0.11
		(-0.01 0.08)	(-0.10 0.03)	(-0.00 0.22)	(-0.03 0.04)	(-0.01 0.07)	(-0.07 0.15)				(-0.09 0.06)	(-0.21 -0.00)	(-0.07 0.30)
	CPI inflation denominator country	-0.06	0.05	-0.07	0.00	0.02	0.01	--	--	--	0.00	0.10	-0.06
		(-0.09 -0.02)	(-0.00 0.10)	(-0.14 -0.01)	(-0.04 0.03)	(-0.04 0.07)	(-0.07 0.10)				(-0.06 0.06)	(0.02 0.19)	(-0.18 0.05)
	Unemployment	--	--	--	--	--	--	-0.17	-0.28	-0.19	--	--	--
								(-0.21 -0.12)	(-0.35 -0.20)	(-0.29 -0.09)			
	Unemployment denominator country	--	--	--	--	--	--	0.09	0.10	0.14	--	--	--
								(0.05 0.13)	(0.04 0.15)	(0.06 0.22)			

Table 3 (continued): Robustness tests

Country	Variable	(1) Benchmark			(2) Alternative denominator country			(3) Adding unemployment			(4) 5-year maturity		
		Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End	Beginning	Middle	End
Japan	Liquidity gap	-0.07	-0.35	-0.52	0.02	0.08	-0.02	0.35	0.02	-0.29	-0.10	-0.47	-0.28
		(-0.25 0.13)	(-0.67 -0.03)	(-0.99 -0.01)	(-0.02 0.06)	(0.03 0.14)	(-0.10 0.05)	(0.15 0.55)	(-0.29 0.38)	(-0.81 0.19)	(-0.31 0.12)	(-0.83 -0.13)	(-0.82 0.27)
	$Vix*10^{-2}$	1.21	0.30	0.50	1.07	0.47	1.36	0.50	0.37	0.54	1.89	0.48	0.40
		(0.94 1.49)	(-0.06 0.64)	(0.1 0.94)	(0.86 1.26)	(0.24 0.72)	(1.06 1.65)	(0.19 0.79)	(0.04 0.7)	(0.09 0.99)	(1.61 2.2)	(0.13 0.83)	(-0.05 0.87)
	Debt	0.28	0.53	0.27	0.09	0.00	0.55	0.13	0.51	0.33	0.06	0.52	0.40
		(0.03 0.53)	(0.17 0.84)	(-0.16 0.67)	(-0.10 0.30)	(-0.31 0.31)	(0.15 0.92)	(-0.13 0.40)	(0.17 0.88)	(-0.06 0.70)	(-0.22 0.34)	(0.16 0.91)	(-0.07 0.90)
	Debt denominator country	-1.96	-2.96	-1.40	-0.07	-0.03	-2.09	-0.60	-1.67	-0.63	-0.94	-2.72	-1.75
		(-2.65 -1.25)	(-3.81 -1.99)	(-2.56 -0.14)	(-0.93 0.79)	(-1.25 1.23)	(-3.94 -0.11)	(-1.37 0.14)	(-2.81 -0.54)	(-1.89 0.54)	(-1.68 -0.11)	(-3.83 -1.73)	(-3.18 -0.35)
	Current account	-0.08	-0.03	-0.05	-0.10	-0.01	-0.06	-0.05	-0.06	-0.03	-0.03	0.03	-0.02
		(-0.11 -0.05)	(-0.08 0.02)	(-0.11 0.01)	(-0.13 -0.07)	(-0.06 0.03)	(-0.12 0.00)	(-0.08 -0.02)	(-0.12 -0.00)	(-0.08 0.03)	(-0.06 0.00)	(-0.03 0.08)	(-0.08 0.03)
	Current account denominator country	-0.03	0.08	-0.04	0.04	0.01	0.01	0.03	0.10	0.01	-0.12	0.07	0.03
		(-0.08 0.02)	(0.01 0.14)	(-0.14 0.05)	(0.01 0.06)	(-0.03 0.04)	(-0.04 0.06)	(-0.02 0.10)	(0.02 0.17)	(-0.09 0.10)	(-0.18 -0.06)	(-0.00 0.15)	(-0.07 0.13)
	GDP growth	0.00	0.01	0.00	-0.01	0.01	0.01	0.02	0.00	0.00	-0.02	0.03	0.01
		(-0.02 0.03)	(-0.02 0.05)	(-0.04 0.04)	(-0.02 0.01)	(-0.02 0.03)	(-0.02 0.04)	(-0.01 0.04)	(-0.03 0.04)	(-0.05 0.05)	(-0.04 0.00)	(-0.01 0.07)	(-0.04 0.05)
	GDP growth denominator country	0.01	0.03	-0.02	0.06	0.01	0.00	-0.01	0.00	-0.03	-0.02	-0.01	-0.06
		(-0.02 0.03)	(-0.01 0.06)	(-0.07 0.03)	(0.05 0.07)	(-0.02 0.04)	(-0.03 0.03)	(-0.04 0.01)	(-0.04 0.04)	(-0.08 0.03)	(-0.05 0.01)	(-0.06 0.03)	(-0.11 -0.00)
	CPI inflation	-0.16	-0.10	0.03	-0.14	-0.10	-0.08	--	--	--	-0.15	-0.05	0.05
		(-0.20 -0.12)	(-0.17 -0.02)	(-0.05 0.11)	(-0.20 -0.09)	(-0.17 -0.02)	(-0.18 0.04)				(-0.19 -0.10)	(-0.13 0.03)	(-0.04 0.13)
	CPI inflation denominator country	0.13	0.04	0.00	0.11	0.09	0.15	--	--	--	0.16	0.05	0.02
		(0.09 0.17)	(-0.02 0.11)	(-0.07 0.09)	(0.08 0.15)	(0.03 0.14)	(0.05 0.24)				(0.12 0.21)	(-0.02 0.12)	(-0.07 0.11)
	Unemployment	--	--	--	--	--	--	0.10	0.08	0.04	--	--	--
								(0.06 0.14)	(0.01 0.14)	(-0.05 0.14)			
	Unemployment denominator country	--	--	--	--	--	--	-0.12	-0.13	-0.10	--	--	--
								(-0.14 -0.09)	(-0.18 -0.08)	(-0.16 -0.04)			
Germany	Liquidity gap	-1.26	-1.72	-0.12	--	--	--	-0.66	-1.64	-0.33	-0.41	-0.43	-0.72
		(-2.00 -0.51)	(-2.70 -0.58)	(-1.65 1.53)				(-1.27 -0.06)	(-2.62 -0.72)	(-1.94 1.22)	(-1.12 0.35)	(-1.54 0.72)	(-2.36 0.78)
	$Vix*10^{-2}$	-0.46	-0.46	-0.81	--	--	--	-0.30	-0.24	-0.84	0.17	0.05	-0.51
		(-0.68 -0.26)	(-0.69 -0.21)	(-1.14 -0.49)				(-0.49 -0.12)	(-0.48 0.03)	(-1.13 -0.53)	(-0.06 0.38)	(-0.21 0.31)	(-0.87 -0.19)
	Debt	1.80	1.64	1.47	--	--	--	2.29	1.46	2.12	1.43	2.17	2.53
		(1.28 2.31)	(0.84 2.49)	(0.16 2.89)				(1.77 2.74)	(0.75 2.32)	(0.84 3.43)	(0.89 2.05)	(1.27 3.07)	(0.99 4.16)
	Debt denominator country	-2.61	-2.54	-1.96	--	--	--	-1.54	-1.54	-2.27	-1.92	-2.03	-2.81
		(-2.93 -2.28)	(-3.11 -1.90)	(-2.55 -1.31)				(-1.92 -1.13)	(-2.27 -0.81)	(-2.94 -1.60)	(-2.27 -1.56)	(-2.64 -1.38)	v-3.57 -2.08)
	Current account	0.00	0.01	-0.03	--	--	--	-0.02	0.00	-0.03	0.00	0.01	0.02
		(-0.02 0.02)	(-0.02 0.04)	(-0.06 -0.00)				(-0.03 -0.00)	(-0.02 0.03)	(-0.06 0.00)	(-0.02 0.01)	v-0.02 0.03)	(-0.01 0.06)
	Current account denominator country	0.04	0.04	-0.01	--	--	--	0.05	0.02	0.03	0.05	0.10	0.11
		(0.01 0.07)	(-0.01 0.09)	(-0.07 0.06)				(0.02 0.07)	(-0.03 0.06)	(-0.04 0.10)	(0.02 0.09)	(0.05 0.15)	(0.04 0.19)
	GDP growth	0.01	0.02	0.00	--	--	--	-0.01	-0.03	-0.01	0.01	0.03	0.00
		(-0.00 0.03)	(-0.01 0.05)	(-0.03 0.04)				(-0.03 0.01)	(-0.07 0.00)	(-0.05 0.03)	(-0.00 0.03)	(0.00 0.06)	(-0.03 0.03)
	GDP growth denominator country	0.00	0.00	-0.01	--	--	--	0.01	0.02	0.01	-0.05	-0.02	-0.05
		(-0.02 0.01)	(-0.02 0.03)	(-0.06 0.03)				(-0.01 0.03)	(-0.00 0.05)	(-0.04 0.05)	(-0.07 -0.03)	(-0.05 0.00)	(-0.10 -0.01)
	CPI inflation	-0.07	-0.08	-0.03	--	--	--	--	--	--	-0.13	-0.11	-0.04
		(-0.10 -0.04)	(-0.13 -0.03)	(-0.11 0.04)							(-0.16 -0.10)	(-0.16 -0.06)	v-0.13 0.04)
	CPI inflation denominator country	0.03	0.04	0.01	--	--	--	--	--	--	0.07	0.07	0.01
		(0.00 0.06)	(-0.01 0.08)	(-0.05 0.07)							(0.04 0.10)	(0.03 0.12)	(-0.05 0.08)
	Unemployment	--	--	--	--	--	--	0.00	0.00	-0.01	--	--	--
								(-0.01 0.02)	(-0.03 0.03)	(-0.05 0.03)			
	Unemployment denominator country	--	--	--	--	--	--	-0.07	-0.07	-0.01	--	--	--
								(-0.10 -0.05)	(-0.11 -0.04)	(-0.05 0.03)			

Note: The table shows results for robustness tests of model (1)-(3). For each model, the table reports the posterior median values of the parameter estimates at the beginning, in the middle and at the end of the sample period, along with the 68% posterior error bands (in brackets). Panel 1 shows the results for the benchmark model, panel 2 for models where the denominator country is changed (to the United States for France and Italy, to Germany for Canada, Japan and the United Kingdom). Panel 3 reports results for a model with unemployment expectations, but without inflation expectations, panel 4 for models of the 5-year government bond spreads.

Appendix

Estimation and priors specification

The time varying model is estimated using Bayesian methods.

It is worth emphasizing that the algorithm used in this paper allow us to compute error bands around the median estimates of the coefficients, thereby providing a very natural way to assess their statistical significance.

As for the specification of the priors, we assume that the priors for the initial states θ_0 of the time varying coefficients and log standard errors $\log \sigma$ are normally distributed. The prior for the hyperparameters, Ω, is assumed to be distributed as an inverse-Wishart, while the distribution of ζ is assumed to be an Inverse Gamma (IG). More specifically, we have the following priors.

- Time varying coefficients: $P(\theta_0) \sim N(\hat{\theta}, \hat{V}_\theta)$ and $P(\Omega) \sim IW(\Omega_0^{-1}, \rho)$.
- Stochastic volatilities: $P(\log \sigma_0) \sim N(\log \hat{\sigma}, I_n)$.
- $P(\zeta) \sim IG(\zeta_0^{-1}, v_0)$.

The hyper-parameters are calibrated using a time invariant OLS regression estimated over all the the sample of size T_0. For the initial states θ_0 we set the mean $\hat{\theta}$ and the variances, $\hat{V}_\theta$ as the maximum likelihood point estimates (the variance il multiplied by four). For the initial states of the log volatilitie, $\log \sigma_0$, the mean of the distribution is chosen to be the logarithm of the point estimates of the standard errors of the residuals of the estimated time invariant OLS regression.

The degrees of freedom for the covariance matrix of the innovations to the drifting coefficients, ρ, are set equal to T_0, the size of the pre-sample. The degrees of freedom v_0, for the prior on the stochastic volatilitie variance ζ, are set equal to 0.001, while the prior d_0, in the scale matrix ζ_0^{-1}, is set equal to 1. The matrix $\Omega_0^{-1} = \lambda \hat{V}_\theta$; λ is the parameter governing the amount of time-variation in the unobserved states, it is fixed to track the optimal percentage of residuals outside the confidence band for a given percentile. Very loose values of λ would imply a large variance of the distribution of the coefficients, and hence a large variance of the distribution of the fitted values. In this case the model would tend to overfit the data, and an overly large percentage of observed data would lie within the confidence bands around the fitted values. The opposite would happen if $\lambda$is very tight. Ideally, we would like to observe 1% of the observed data to lie outside the 1% confidence bands, 2% to lie outside the 2% confidence bands and so on; in other words, plotting the percentages of observations outside the respective confidence bands against the theoretical percentages, we should expect a line that is close to the 45 degree line. We fix the parameter lambda as the value that minimizes the distance from the theoretical 45 degree line.

Gibbs sampling algorithm

Estimation is performed using Bayesian methods. To draw from the joint posterior distribution of model parameters we use a Gibbs sampling algorithm similar to the one described by Primiceri (2005). The idea behind the algorithm is to draw sets of coefficients from known conditional posterior distributions. The algorithm is initialized at some values and, under some regularity conditions, the draws converge to a draw from the joint posterior after a burn in period. Let z be a $(q \times 1)$ vector, and z^T denote the sequence $[z'_1, ..., z'_T]'$. Each repetition is then composed of the following steps, with s^T to be defined below:

1. $p(s^T|x^T, \theta^T, \sigma^T, \Omega, \zeta)$
2. $p(\sigma^T|x^T, \theta^T, \Omega, \zeta, s^T)$
3. $p(\theta^T|x^T, \sigma^T, \Omega, \zeta, s^T)$
4. $p(\Omega|x^T, \theta^T, \sigma^T, \zeta, s^T)$
5. $p(\zeta|x^T, \theta^T, \sigma^T, \Omega, s^T)$

- Step 1: sample from $p(s^T|y^T, \theta^T, \sigma^T, \Omega, \zeta)$

Conditional on $y^{**}_{i,t}$ and r^T, we independently sample each $s_{i,t}$ from the discrete density defined by $Pr(s_{i,t} = j|y^{**}_{i,t}, r_{i,t}) \propto f_N(y^{**}_{i,t}|2r_{i,t} + m_j - 1.2704, v_j^2)$, where $f_N(y|\mu, \sigma^2)$ denotes a normal density with mean μ and variance σ^2.

- Step 2: sample from $p(\sigma^T|y^T, \theta^T, \phi^T, \Omega, \zeta, \Psi, s^T)$

To draw σ^T we use the algorithm of Kim, Shephard and Chibb (KSC) (1998). Consider the system of equations $y^*_t \equiv (y_t - X'_t\theta_t) = \varepsilon_t^{1/2} u_t$, where $u_t \sim N(0, I)$, $X_t = (I_n \otimes x'_t)$, and $x_t = [1_n, x_{1,t}...x_{k,t}]$. Conditional on y^T,and θ^T y^*_t is observable. Squaring and taking the logarithm, we obtain

$$y^{**}_t = 2r_t + \upsilon_t \tag{1}$$
$$r_t = r_{t-1} + \xi_t \tag{2}$$

where $y^{**}_{i,t} = \log((y^*_{i,t})^2 + 0.001)$ –the constant (0.001) is added to make estimation more robust– $\upsilon_{i,t} = \log(u^2_{i,t})$ and $r_t = \log \sigma_t$. Since, the innovation in (1) is distributed as $\log \chi^2(1)$, we use, following KSC, a mixture of 7 normal densities with component probabilities q_j, means $m_j - 1.2704$, and variances v_j^2 (j=1,...,7) to transform the system in a Gaussian one, where $\{q_j, m_j, v_j^2\}$ are chosen to match the moments of the $\log \chi^2(1)$ distribution. The values of the parameters are reported in table 1.

Let $s^T = [s_1, ..., s_T]'$ be a matrix of indicators selecting the member of the mixture to be used for each element of υ_t at each point in time. Conditional on s^T, $(\upsilon_{i,t}|s_{i,t} = j) \sim N(m_j - 1.2704, v_j^2)$, we can use the algorithm of Primiceri (2005) to draw r_t (t=1,...,T) from $N(r_{t|t+1}, R_{t|t+1})$, where the mean $r_{t|t+1} = E(r_t|r_{t+1}, y^t, \theta^T, \Omega, \zeta, s^T,)$ and the variance $R_{t|t+1} = Var(r_t|r_{t+1}, y^t, \theta^T, \Omega, \zeta, s^T)$.

- Step 3: sample from $p(\theta^T|y^T, \sigma^T, \Omega, \zeta, s^T)$

Conditional on all other parameters and the observables we have

$$y_t = X'_t\theta_t + \varepsilon_t \tag{3}$$
$$\theta_t = \theta_{t-1} + \omega_t \tag{4}$$

Table 1: ***Parameters Specification***

j	q_j	m_j	v_j^2
1.0000	0.0073	-10.1300	5.7960
2.0000	0.1056	-3.9728	2.6137
3.0000	0.0000	-8.5669	5.1795
4.0000	0.0440	2.7779	0.1674
5.0000	0.3400	0.6194	0.6401
6.0000	0.2457	1.7952	0.3402
7.0000	0.2575	-1.0882	1.2626

Draws for θ_t can be obtained from a $N(\theta_{t|t+1}, P_{t|t+1})$, where $\theta_{t|t+1} = E(\theta_t|\theta_{t+1}, y^T, \sigma^T, \phi^T, \Omega, \zeta, \Psi)$ and $P_{t|t+1} = Var(\theta_t|\theta_{t+1}, y^T, \sigma^T, \Omega, \zeta)$ are obtained with the algorithm of Primiceri (2005).

- Step 4: sample from $p(\Omega|y^T, \theta^T, \sigma^T, \zeta, s^T)$

Conditional on the other coefficients and the data, Ω has an Inverse-Wishart posterior density with scale matrix $\Omega_1^{-1} = (\Omega_0 + \sum_{t=1}^{T} \Delta\theta_t(\Delta\theta_t)')^{-1}$ and degrees of freedom $df_{\Omega_1} = df_{\Omega_0} + T$, where Ω_0^{-1} is the prior scale matrix, df_{Ω_0} are the prior degrees of freedom and T is length of the sample use for estimation. To draw a realization for Ω, we make df_{Ω_1} independent draws z_i (i=1,...,df_{Ω_1}) from $N(0, \Omega_1^{-1})$ and compute $\Omega = (\sum_{i=1}^{df_{\Omega_1}} z_i z_i')^{-1}$ (see Gelman et. al., 1995).

- Step 5: sample from $p(\zeta|y^T, \theta^T, \sigma^T, \Omega, s^T)$

Conditional to the other coefficients and the data, ζ has an Inverse-Gamma posterior density with scale matrix $\zeta_1^{-1} = (\zeta_0 + \sum_{t=1}^{T} \Delta \log \sigma_t(\Delta \log \sigma_t)')^{-1}$ and degrees of freedom $df_{\zeta_1} = df_{\zeta_0} + T$ where ζ_0^{-1} is the prior scale matrix and df_{ζ_0} the prior degrees of freedom.