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Euro Area Sovereign Bond Spreads, Risk Aversion & Monetary Policy

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Abstract

This paper provides a framework to disentangle the effects of uncertainty and risk aversion on Euro area sovereign bond yields vis-à-vis Germany, which were often confounded in previous studies. In a second step, the impact of European monetary policy on spreads through its impact on risk aversion is estimated. The background of this investigation is the reduction of spreads after Mario Draghi's OMT announcement in 2012: To what extent did the ECB lower spreads by increasing market participants' willingness to take on risks? The results indicate that risk aversion was a more important factor than uncertainty. The effect of monetary policy on spreads via risk aversion however is ambiguous at best and highly sensitive to the model specification and measurement of the policy stance.

Keywords: Monetary policy, Risk aversion, Uncertainty, Sovereign bond spreads
JEL: E43, E44, E52, E62, G12, G14, H62, H63

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Contents

List of Figures

List of Tables

1 Introduction

Since the introduction of the common European currency in 2002, the bond yields of the participating countries have experienced three markedly different periods. After introducing the Euro, all participating countries saw their sovereign yields initially converge to German yields with spreads on average below fifty basis points. This convergence ended abruptly with the outbreak of the financial crisis and the subsequent Euro crisis, when Greece applied for financial support in May 2010, followed by Ireland (November 2010) and Portugal (April 2011). However, spreads continued to increase drastically and reached a peak of around 1,100 basis points (bp) for Ireland on the 18th of July 2011, around 1,400 bp for Portugal on the 30th of January 2012 and almost 4,800 bp for Greece on the 8th of March 2012 (see Figure 1). It is perhaps worth noting that not a single country was spared from this: Even a country that is considered as relatively safe like the Netherlands saw its spreads rise on average by a factor of eight compared to the pre-crisis period. Only after Mario Draghi's now famous remarks on the 26th of July 2012, that the ECB was "ready to do whatever it takes to preserve the Euro", spreads began to decline again. However, no country has seen its yields vis-à-vis Germany return to their pre-crisis average yet.

A wide range of literature has addressed this divergence since 2009 (e.g., Attinasi, Checherita & Nickel, 2011; Barrios, Iversen, Magdalena & Setzer, 2009; Beber, Brandt & Kavajecz, 2009). Most commonly, government bond yield differentials are ascribed to three different factors: First and foremost, spreads indicate the relative risk of default. Secondly, differing degrees of liquidity influence the relative attractiveness of an asset. Finally, investors' risk aversion is a common factor influencing several countries at the same time. Risk aversion is related to the willingness of investors to bear risk, which may vary even when the amount of risk remains unchanged. Therefore, the compensation demanded to bear risk, i.e. the price of risk, also can vary independently of the underlying amount of risk. This is particularly obvious in times of heightened financial stress when investors demand a higher risk compensation because they prefer less risky assets. This common factor for risk aversion might in turn be influenced by the stance of monetary policy as, for example, Diamond and Rajan (2009) argued prominently.

Disentangling those factors is challenging in so far as they might be endogenously determined and only their aggregate effect can be observed: Deteriorating economic fundamentals could increase sovereign yields (compared to Germany), which in turn could negatively affect the fiscal outlook of the respective country again. Furthermore, data availability, especially

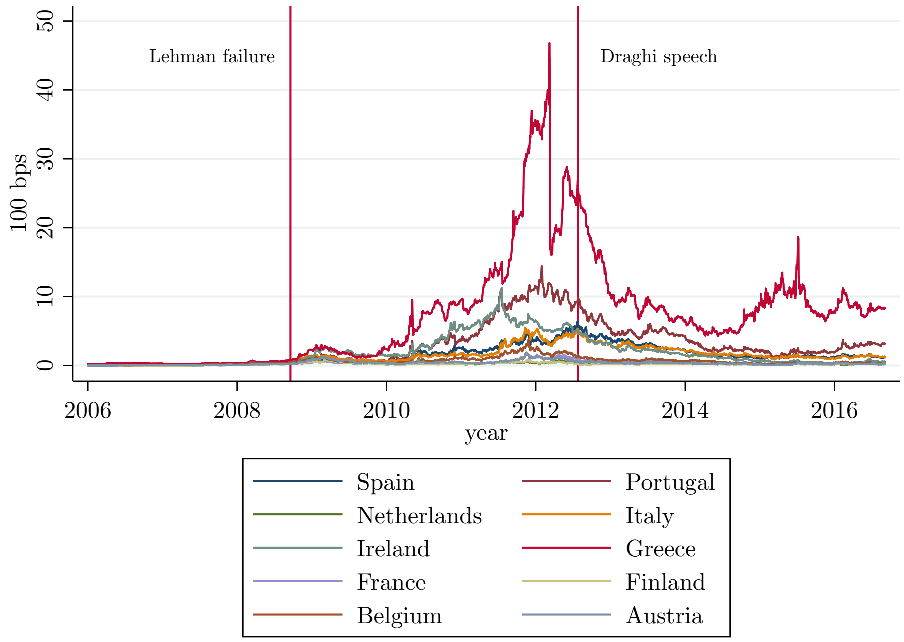

Figure 1: Sovereign yields in the Eurozone relative to Germany

for investors' risk aversion, is necessarily limited and often has to be proxied. Most commonly, risk aversion is proxied by applying a principal factor analysis (PCA) to a set of variables (e.g., Barrios et al., 2009), using the spread between AAA rated corporate bond yields and 10-year U.S. Treasury yields (e.g., Attinasi et al., 2011) or by using the VIX (e.g., Giordano, Pericoli & Tommasino, 2013). However, all approaches suffer from their inability to distinguish between changes in genuine risk aversion and changes in uncertainty (although both phenomena often go hand in hand).

The study's first main contribution is to make use of a relatively novel approach to proxy risk aversion and uncertainty separately as proposed by Bekaert, Hoerova and Lo Duca (2013). They split the VIX into a measure of expected volatility and a residual which serves as a proxy for risk aversion. As a second main contribution, this paper attempts to quantify the effect of monetary policy on yield spreads working through risk aversion, thus providing a more detailed analysis of the transmission mechanism of monetary policy during the crisis. In other words: To what extend, if at all, did monetary policy lower spreads by increasing market participants' willingness to take on risks? The third main contribution is of methodological nature as this paper combines methods of dealing with non-stationary and cointegration as well as time varying coefficients.

Disentangling the effects of uncertainty and risk aversion as well as assessing the impact of monetary policy on risk aversion is of interest in so far as it would allow for a better tailored policy response to macroeconomic problems: S. R. Baker, Bloom and Davis (2015a), for example, show that heightened economic uncertainty[1] decreases employment as well as output and also triggers jumps in stock and bond markets (S. R. Baker, Bloom & Davis, 2015b). This might require a more streamlined economic policy whereas problems induced by high risk aversion might require a more aggressive approach to handle problems in the financial markets and sovereign debt sustainability. For monetary policy authorities, high risk aversion might justify the use of more unconventional policy measures. Furthermore, many commentators, especially in German-speaking countries, argue that the decline in spreads reduces budget discipline because markets do not punish unsustainable fiscal policy as they would under normal circumstances. The key question is, does monetary policy drive down risk aversion and thereby suppresses spreads? Or is the decline in risk aversion not caused by monetary policy? Identifying the impact of monetary policy on spreads via its impact on risk aversion would help substantiating the public argument about the sensibility of the ECB's accommodative monetary policy.

The remainder of the paper is structured as follows: The next section provides an overview of the existing literature. Section 3 briefly outlines the data used for estimation. Section 4 describes the econometric model; the results are discussed in section 5. Section 6 concludes.

2 Related Literature

This section briefly reviews some of the relevant literature for the topic of this paper which builds on three different strands of research: the many proposed measures of risk aversion (Section 2.1), the extensive research on the determinants of Euro area sovereign spreads (Section 2.2) and the literature about the effect of monetary policy on the willingness and capacity of markets to take on risks (Section 2.3).

[1] Measured by newspaper coverage of policy-related uncertainty, the number of U.S. federal tax code provisions set to expire and disagreement among economic forecasters

2.1 Measuring Risk Aversion

Various asset pricing models show that risk premia are of fundamental importance to fully explain asset price movements. In theoretical models, for example the CAPM and its derivatives (Campbell, 2003), the risk premium of an asset can be decomposed into the *price of risk* (i.e. the return an investor demands to bear an amount of risk) and the *quantity of risk* (i.e. the range of possible outcomes). The price of risk is common to all assets whereas the quantity of risk is specific for each asset: A stock, for example, is riskier than an U.S. government bond, but both need to pay the same compensation for each unit of risk. The combination of the price and the quantity of risk, that is the risk premium, is the inverse of *risk appetite*, i.e. the willingness of investors to bear risk. The price of risk is often used interchangeably with risk aversion, while the quantity of risk corresponds to uncertainty.

Market based measures of risk aversion are empirically derived from various sources: In the behavioural finance literature sentiment indices are constructed using data such as closed-end fund discounts or IPO information (e.g., M. Baker & Wurgler, 2007). One approach used in empirical finance is the comparison of risk-neutral probabilities with investors subjective probabilities (e.g., Jackwerth, 2000). In practice, financial institutions use a range of indicators which Coudert and Gex (2008) survey. Among the simpler measures are indicators such as the spread between corporate bond yields and 10-year Treasury yields or the VIX, the option implied volatility of the S&P 500 index. Popular among the more elaborate indicators is the application of a principal component analysis to risk premia of different securities to identify a common factor in their variations. The first factor usually explains the largest share of the variation and is thus interpreted as representing investors' risk aversion. A prominent example, the Global risk aversion indicator, is computed by the European Systemic Risk Board using the first principal component of the Commerzbank Global Risk Perception index, the UBS FX Risk Index, Westpacs Risk Appetite Index, the BoA ML Risk Aversion Indicator and the Credit Suisse Risk Appetite Index.

Another type of indicator is the global risk aversion index (GRAI) type developed by Persaud (1996). It relies on the assumption that an increase in risk aversion should increase risk premia across all asset classes, but the increase should be higher for the riskier asset classes. Changes in the riskiness of a particular asset, however, should not affect the returns of other assets. By employing a rank correlation on price variations of different securities and their volatility, it is possible to derive a measure of risk aversion: A positive correlation indicates a decrease of risk aversion, a negative correlation an increase. By construction, the GRAI does not measure the level of risk aversion but its change.

Basically all the aforementioned measures of risk aversion however suffer from the problem that they are in fact driven by both the price of risk and the quantity or risk. This is particularly clear in case of corporate bond spreads: The spread between AAA rated corporate bond yields and 10-year Treasury yields is driven by general uncertainty about the fate of institutions considered less safe than the U.S. government and by investors' willingness to bear that additional risk. In that sense, they are closer to being a measure of *risk appetite* than *risk aversion*.

To overcome this problem Bekaert and Hoerova (2014; 2016; Bekaert et al., 2013) in a series of papers make use of the fact that the VIX measures risk-neutral volatility and not actual expected volatility and thereby follow the approach of Jackwerth (2000). The difference between risk-neutral and actual volatility lies in the adjustment for investors' risk preferences: Typically, the price of a risky payout will be below its expected value as risk averse investors demand compensation for bearing that risk. For example, it is likely that the price of a fair coin toss with payouts of either 0 or 2 will be below the expected value of 1 to compensate investors' risk aversion, where the difference between expected value and the actual price is the risk premium. The actual price can be used to infer the risk-neutral probabilities, i.e. the probability measure under which the fair price (expected-value) equals the actual price. If the actual price of the bespoke coin toss turns out to be 0.5, the risk-neutral probability of winning would be 25% instead of 50%, thereby eradicating the risk premium. The same holds true for measures of volatility: Physical expected volatility uses the actual probabilities, whereas risk-neutral volatility makes use of probabilities that are adjusted for the pricing of risk.

The risk-neutral volatility usually turns out to be higher than the physical volatility (Carr & Wu, 2009). The difference between those two measures, the variance premium, reflects the impact of the price of risk and therefore is a good proxy for risk aversion. This is emphasized by Bakshi and Madan (2006) who link the variance premium to the coefficient of relative risk aversion within a representative agent model like the CCAPM. Bekaert and Hoerova (2014) further motivate the suitability of the proxy in a theoretical model, demonstrating that the variance premium is indeed increasing in risk aversion.

Bekaert and Hoerova use this property of the VIX to decompose the index into a measure of uncertainty and a measure of risk aversion by calculating the difference between the squared VIX (i.e. the option implied variance) and the estimated expected (conditional) variance. The expected variance is used as a measure of uncertainty whereas the residual is a proxy for risk aversion. Bekaert et al. (2013) find that monetary policy affects both measures, even when controlling for business cycle movements. Of both measures, only positive uncertainty shocks exert a significant (negative) influence over industrial production. Furthermore, Bekaert and Hoerova (2014) find that risk aversion is a good predictor for stock returns while uncertainty is not. The opposite is the case for financial instability, which uncertainty predicts better than risk aversion.

2.2 Euro Area Sovereign Spreads

The pricing of sovereign bonds has been investigated in a wide range of literature beginning long before the Euro crisis. Typically, these studies explore the role of country specific credit and liquidity risks as well as risk aversion as a common factor. Bernoth, Von Hagen and Schuknecht (2012) derive the following, widely used type of regression from a standard portfolio choice model to explain spreads between bond yields:

$$y_{i,t} - y_{b,t} = \alpha_i + \beta_{1,i} CREDIT_{i,t} + \beta_{2,i} LIQUIDITY_{i,t} + \beta_{3,i} RISK_t + \epsilon_{i,t} \quad (1)$$

where $y_{i,t} - y_{g,t}$ is the spread between yields on bonds of country i and the benchmark at time t, $CREDIT_{i,t}$ the credit risk, $LIQUIDITY_{i,t}$ the liquidity risk and $RISK_t$ the common risk aversion proxy. Other technical factors (e.g., tax differences) are usually neglected because their effect is likely to be small in globally integrated financial markets like the Eurozone. The literature differs in the respective data frequency and the estimation method employed (especially more recent papers deal with possible cointegration relationships between spreads and fundamentals).[2] Many papers include further regressors in order to capture effects like the size of the respective financial sector or employ vector (error correcting) autoregressive models to estimate contagion effects.

Bayoumi, Goldstein and Woglom (1995) were among the first to study yield differences between sovereign debt issuers, thereby laying ground for the later literature on Euro area spreads. They find that the debt level of U.S. states increases their yields relative to other states. Beber et al.

[2] As Giordano et al. (2013) already noted, caution should be applied when interpreting the results of previous papers which did not consider the possible existence of a cointegration relationship

(2009) show that differences in credit quality explain a large extend of the yield spreads between April 2003 and December 2004. Liquidity on the other hand has an important impact in low credit risk countries and during times of heightened market uncertainty. This is particularly extreme during periods of large flows into or out of the bond market: In that case liquidity has a much bigger effect than otherwise, which suggests that in times of heightened market stress liquidity plays a bigger role than quality.

The global financial crisis of 2007-2009 and the subsequent Euro crisis spurred a new wave of studies. Attinasi et al. (2011) use a dynamic panel approach of Euro area sovereign spreads and find that economic fundamentals, risk aversion,[3] liquidity and announcements of bank rescue packages determine spreads. In particular, they find that international risk aversion as proxied by the corporate AAA - 10-year Treasury bond spread can explain up to 56% of the changes in daily sovereign bond yield spreads. Manganelli and Wolswijk (2009) show that a high correlation between spreads and the Euro short-term interest rate exists, even after controlling for credit ratings and international risk aversion (measured by spreads between 10-year US interest rate swaps and Treasury bonds). The correlation between spreads and the interest rate is higher for countries with low sovereign ratings. Similarly, the impact of liquidity (measured by the AAA - Bund spread) is negatively related to the interest rate. They argue further that the convergence of spreads before the crisis was not driven by the lack of credibility of the no bail-out clause because Euro area corporate bond spreads, which they argue are not subject to bailouts, experienced a similar trend.

Barrios et al. (2009) apply a PCA to weekly data of yield spreads vis-à-vis Germany to construct a sovereign risk indicator. In the same fashion, they use AAA-BBB corporate bond spreads, the VSTOXX and the volatility in the Euro-Yen exchange rate to form a general risk aversion indicator. The authors find that sovereign risk and risk aversion moved closely together until 2008Q3. After that, general risk aversion levels off while sovereign risk keeps increasing until early 2009. Barrios et al. (2009) explain this divergence with the transfer of risk from the banking to the public sector in many states, which decreased general risk aversion at the cost of increasing sovereign risk. Their econometric analysis shows that risk aversion significantly affected Belgian, French, Italian and Portuguese spreads during the crisis period (August 2007 - April 2009) but not Greek and Spanish spreads.

[3]Even though the standard measures of risk aversion often confound the price of risk and the quantity of risk and therefore are closer to risk appetite than risk aversion, as described in the previous section, in the following the term risk aversion will be used interchangeably in order to reflect the authors original terminology.

High-debt countries such as Belgium, Portugal and Italy are affected more strongly by the risk aversion indicator. Furthermore, Barrios et al. (2009) also demonstrate that the impact of risk aversion on bond spreads is higher for countries with bigger deficit forecasts and even more pronounced when the deficit forecast is accompanied by a large current account deficit.

Bernoth et al. (2012) argue that markets reacted more strongly to deficit numbers than debt levels before the introduction of the Euro. However, markets paid more attention to debt levels than to deficits between the introduction of the Euro and the crisis. During the crisis, all measures of fiscal performance become significant. Risk aversion (measured by the corporate Baa - 10-year Treasury bond spread) affected spreads before the crisis only on US Dollar denominated bonds issued by European sovereigns, but not on DM or Euro denominated issues. This is taken as a sign that European bonds did not enjoy a safe haven status during that time. With the onset of the crisis German bonds gained a safe-haven status as increases in risk aversion affected yield spreads in both currencies equally. They estimate risk aversion to explain 120 bp of the spreads of all countries.

Bernoth and Erdogan (2012) apply a time-varying coefficient panel model to European sovereign yield spreads between 1999-2010. Similarly to Bernoth et al. (2012), they argue that spreads before the onset of the crisis were not affected by deficit spreads relative to Germany, only by debt to GDP ratios. Furthermore, they find that risk aversion (proxied by the corporate BBB - 10-year Treasury bond spread) played an important role during the introduction of the common currency but between Q1/2001 and Q3/2006 its impact decreased significantly and even turned insignificant for a while, indicating that Germany lost its status as a safe haven during that period. However, already two years before the bankruptcy of Lehmann, risk aversion became a significant factor again, with its coefficient reaching an unprecedented level in early 2010. Liquidity risk on the other side was insignificant during the whole time.

Kilponen, Laakkonen and Vilmunen (2012) study the impact of decisions by the ECB and other European crisis resolution mechanisms (the European Economic Recovery Plan, the EFSF and its successor, the ESM) during the financial crisis and contagion effects. They find only mixed evidence of systemic contagion. Of all crisis resolution mechanisms, the ECB's Securities Market Program had the biggest impact on yield spreads. Decisions to grant rescue packages by the ESFS or the ESM decreased yields in the receiving countries, but increased yields in Spain and Italy. Beetsma, Giuliodori, de Jong and Widijanto (2013) on the other hand use the Eurointelligence newsflash to construct news variables and find evidence for significant spillovers between GIIPS countries after bad news, especially when the affected country has large cross-border bank holdings. Spillovers from GIIPS to non-GIIPS countries following bad news are also observed but are much less pronounced. Similarly, Gerlach, Schulz and Wolff (2010) find that the size of the banking sector positively affects the spread of its sovereign in times of heightened risk aversion. The effect is even more pronounced when the banking sector is underfunded, i.e. when the aggregate equity buffer is small. Both effects, however, are not present when risk aversion is on a normal level.

More recent papers like Santis (2012) use multiequation econometric techniques. Arezeki et al. (2011) estimate a vector auto regression (VAR) model and find that sovereign rating news have significant contagion effects on both other countries and financial markets. Santis (2012) allows for a long-run cointegrating relationship between spreads and other variables, emphasizing the results of Arezeki et al. (2011) regarding contagion across countries. However, Santis (2012) also argues that international risk aversion, as measured by the corporate BBB - 10-year Treasury bond spread, is insufficient to explain the surge in spreads during the crisis and additionally uses the KfW-Bund spread as a measure of regional risk aversion. Giordano et al. (2013) on the other hand find no direct contagion effects of Greek rating changes, only indirect effects when interacted with fiscal or macroeconomic variables.

Overall, the existing literature suggests that fiscal variables significantly affect sovereign spreads. The magnitude of that effect, however, is much larger during the crisis than before. This is consistent with some degree of mispricing of sovereign debt before the crisis. The impact attributed to liquidity in most papers is rather small, in some papers even insignificant. Contagion effects seem to be present at least for some countries to some extent and most papers suggest that a large and underfunded banking sector has an adverse effect on sovereign yields. Furthermore, global risk aversion and public risk seem to decouple in 2009 consistent with the transfer of risk from private to public balance sheets. However, much of the previous literature has used ambiguous indicators of risk aversion that confound the price of risk and the quantity of risk. Disentangling those two phenomena will be this paper's first main contribution.

2.3 Risk Aversion and Monetary Policy

A growing body of literature investigates the micro-effects of monetary policy on risk taking on financial markets. The risk-taking channel of monetary policy, as first described by Borio and Zhu (2012), explains how low interest rates can promote risky behaviour on financial markets by two channels: Low interest rates might alter asset managers' incentives because they are bound to sticky targets of return by their investors who are subject to money illusion. Secondly, the ability of financial institutions to take on risk increases because low interest rates increase the value of their collateral and their interest rate margins (e.g., Adrian & Shin, 2010; Adrian, Moench & Shin, 2010) and reduces the attractiveness of safe assets relative to riskier assets with higher expected returns (Rajan, 2006). Furthermore, monetary policy mainly affects short-term rates, which is especially important for banks mainly depending on short-term wholesale funding. Increasing liquidity of short-term debt combined with agency problems can lead to a decrease in lending standards (Diamond & Rajan, 2009).

This kind of argument has been corroborated empirically by, for example, Jiménez, Ongena, Peydró and Saurina (2014), who use 22 years of data from the Spanish credit registry to find that banks soften their lending standards and provide more risky new loans in times of low short-term interest rates. The credit risk of outstanding loans on the other hand is reduced through lower rates. Ioannidou, Ongena and Peydró (2015) use the full dollarisation in Bolivia to treat US monetary policy as exogenous. Their findings are similar to Jimenez et al. (2014), adding that loan spreads do not increase with riskier new loans which implies that banks do not price risk appropriately. Delis and Kouretas (2011) find those effects to be more pronounced for European banks that engage to a larger extend in non-traditional activities, as measured by a higher volume of off-balance sheet items. This holds across all levels of capitalisation. Adrian and Shin (2010) find that lower nominal rates affect interest margins of financial intermediaries and thereby the size of their balance sheet, which in turn has consequences on credit supply and output growth. This behaviour can be amplified through the financial accelerator (Bernanke, Gertler & Gilchrist, 1999) with the only difference that the effect starts out from lenders instead of borrowers.

While a range of empirical studies certainly make a case for increased bank risk-taking under low interest rate regimes, there is much less empirical literature on the link between monetary policy, risk aversion and asset prices. Among those Baekert and Hoerova (2014) show that their risk aversion estimate (as presented in Section 2.1) is a significant predictor of stock returns but not of output growth. This paper's second main contribution therefore will be to extend the literature about monetary policy on risk aversion to sovereign yield spreads.

3 The Data

The dataset covers eleven Euro area countries (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal, Spain and the Netherlands) from January 2003 to August 2016. Luxembourg was excluded because it lacks any noteworthy public debt, both in relative and absolute terms.[4] The remaining countries were excluded because they all joined the Euro relatively late (Slovenia in 2007, Malta & Cyprus in 2008, Slovakia in 2009, Estonia in 2011, Latvia in 2014 and Lithuania in 2015) and thus provide only a limited number of observations. The variable to be explained is the difference between the yield on the respective country's 10-year bonds and the corresponding yield of German 10-year bonds.

3.1 Risk Aversion and Uncertainty

In order to create a common proxy for uncertainty and risk aversion in the Eurozone, this paper follows the approach of Bekaert et al. (2013): The option implied volatility index of the EURO STOXX 50, the VSTOXX, is decomposed into an uncertainty and a risk aversion series by subtracting the expected future realized variance of the underlying index from the squared VSTOXX. Given that the VSTOXX reflects options on the fifty largest and most liquid stocks in the Eurozone, there is little reason to doubt the functioning of the price discovery process of neither the underlying market nor the derived option market. Since the VSTOXX is derived from that option market there are no liquidity issues either.

The expected future realized variance of the EURO STOXX 50 is estimated by regressing the future realized monthly variance onto a set of variables. While Bekaert et al. (2013) use the lagged realized monthly variance and the squared VIX to estimate the expected variance for the S&P500, a horse race was conducted to find the model with the best fit for the EURO STOXX 50 in terms of RMSE, MAE and MAPE (see Section A1 for the detailed results). The best model in each statistic contains the lagged realized monthly, weekly and daily variance as well as the squared VSTOXX. [5] This specification optimally balances the most recent information contained in the lagged daily variance with longer term trends contained in the lagged realized weekly and monthly variances, thereby also smoothing out

[4] In 2007 the debt to GDP ratio was 7.2%, at its peak in 2013 it was at 23.3%

[5] The model is estimated with all variables even though some of them are not statistically significant in order to avoid omitted variable bias in the other coefficients.

any spikes in the daily variance.

$$\begin{aligned} RVAR^{(22)}_{EU,t} = \underset{(.0001)}{0.0005} + \underset{(.0574)}{0.4017} VSTOXX^2_{t-22} - \underset{(0.0648)}{0.0169} RVAR^{(22)}_{EU,t-22} \\ + \underset{(0.3002)}{0.5070} RVAR^{(5)}_{EU,t-22} + \underset{(0.8920)}{2.8115} RVAR^{(1)}_{EU,t-22} + \epsilon_{EU,t} \end{aligned} \tag{2}$$

The realized monthly variance at day t is defined as the sum of daily realised variances over 22 (business) days $RVAR^{(22)}_t = \sum_{j=1}^{22} RV_{t-j+1}$. RV_t is the daily realized variance computed by the Oxford Man Institute by summing up squared 5 minute returns (Heber, Lunde, Shephard & Sheppard, 2013). The regression standard errors are corrected for heteroskedasticity and serial correlation using 12 Newey-West (1987) lags.[6]

Table 1: Summary statistics - Risk aversion & uncertainty

Variable	Mean	Std. Dev.	Min.	Max.	N
VSTOXX	23.936	7.729	11.6	87.510	3569
- pre-crisis	20.820	18.561	11.6	53.14	1488
- crisis	31.391	10.661	18.360	87.510	1009
- post-crisis	21.245	4.707	12.710	40.800	1072
Uncertainty (Eurozone)	34.188	28.917	9.954	323.49	3569
- pre-crisis	27.253	21.798	9.954	156.773	1488
- crisis	53.088	40.109	17.231	323.490	1009
- post-crisis	26.027	11.912	11.191	111.779	1072
Risk aversion (Eurozone)	21.548	24.261	-8.628	314.677	3569
- pre-crisis	15.685	18.561	-8.628	117.453	1488
- crisis	38.399	33.165	9.391	314.677	1009
- post-crisis	13.826	8.506	-8.616	65.841	1072

The fitted values of the regression are the expected realised variance or uncertainty series (UC_t). The difference between the squared VSTOXX (i.e. the implied variance for the next 30 days) and the expected realised variance over the next 22 business days (approximately corresponding to one calendar month, thereby having the same horizon as the VSTOXX) at time t, i.e. the fitted value of the regression, is the proxy for risk aversion: $RA_t = VSTOXX^2_t - E_t[RVAR_{t+22}]$. Table 1 provides the summary statistics of the series and Figure 2 shows that both series behave very similar to the ones constructed by Bekaert et al. (2013).[7] The regression was estimated using a dataset starting in January 2000, the earliest date at which data on realised

[6] Number of lags was chosen according to the rule of thumb $L = 0.75N^{1/3}$

[7] Both series are expressed in percentages squared for convenience.

variance is available from the Oxford Man institute.[8] The fitted values before 2003 were dropped as they lie outside the period under observation for the other variables.

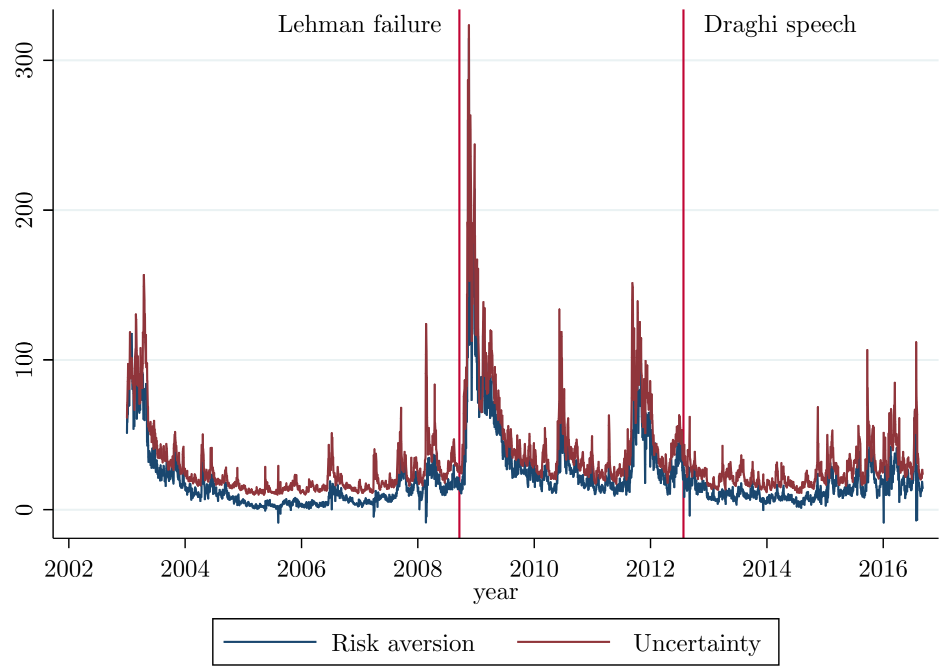

Figure 2: Risk aversion and uncertainty estimates in the Eurozone

During the crisis (the period between the 15th September 2008, the day Lehman Brothers went bankrupt, and the 16th July 2012, the day of Mario Draghi's "whatever it takes" speech) both uncertainty and risk aversion rose to levels of more than twice their pre-crisis (1st January 2003 until 14th September 2008) levels, as Figure 2 and Table 1 show. After Mario Draghi's speech both series fell drastically again and have ever since remained depressed with values even below their pre-crisis means. The same pattern is observable for the variance of both series. Only the VSTOXX has remained slightly elevated compared to pre-crisis levels and did not see a comparable

[8]The VSTOXX, obtained from Thomson Reuters Datastream, is calculated as a rolling index with a horizon of 30 days and expressed in annualised terms. In order to have the VSTOXX on the same scale as the dependent variable in Equation 2, the VSTOXX is brought down to a monthly level and expressed as percentage term (i.e. by dividing it through the square root of 12 and multiplying by 100).

spike in variance.

3.2 Spreads

This paper uses yields of generic 10-year government bonds obtained from Thomson Reuters Datastream. Spreads are computed as the difference between the sovereign yield of the respective country and the yield on German bonds. The data are collected at the end of each business day.

Table 2: Summary statistics - Sovereign bond yields vis-à-vis Germany

Variable	Mean	Std. Dev.	Min.	Max.	N
Spain	1.199	1.366	-0.059	6.354	3569
Portugal	2.328	2.953	-0.038	14.41	3569
Netherlands	0.19	0.165	-0.134	0.842	3569
Italy	1.224	1.201	0.061	5.498	3569
Ireland	1.498	2.089	-0.237	11.249	3569
Greece	5.786	7.56	0.069	46.802	3569
France	0.307	0.31	-0.113	1.887	3569
Finland	0.161	0.176	-0.2	0.827	3569
Belgium	0.499	0.546	-0.106	3.66	3569
Austria	0.278	0.295	-0.136	1.832	3569

Gaps in the data for the VSTOXX as well as uncertainty and risk aversion were filled with the last available data point to match the data availability of government spreads. Gaps in the latter remain unchanged.

3.3 Country Specific Risk

The most commonly used indicators of a country's credit risk are credit default swap (CDS) prices, (expected) fiscal fundamentals (e.g., the debt to GDP ratio) or sovereign ratings. Of all three indicators, CDS prices are available at the highest frequency and in theory should quickly incorporate all news relevant to assess sovereign credit quality: As investors use all available information at each point in time to form expectations about a country's fiscal position, there should be no anticipated changes in CDS prices. Therefore, every change in prices should be a response to new information. This seems to be especially important during the height of the Euro crisis when a country's fiscal position was likely dependent on the stance of the negotiations about sovereign and financial bailouts. However, CDS markets are still relatively young and exist only for a few European countries. Furthermore, CDS prices are not only driven by fundamentals but also by counterparty risk (however, this effect in practice is small as per Duffie and Huang (1996)), liquidity and by policy interventions like the ban of naked short selling (e.g., Hendershott, Namvar & Phillips, 2013). This casts doubt upon the efficiency of price discovery in CDS markets.

Ratings have the advantage of being released much closer to relevant news than official budget numbers and significantly affect investment decisions of institutional investors. However, it seems likely that market participants often anticipate rating changes so that they do not contain new information for market yields. Furthermore, the combination of rapidly decreasing spreads and constant ratings since 2012 suggests that rating agencies changed their assessment of sovereign credit risk and became more cautious, as argued by de Vries and de Haan (2015). They also show that the impact of sovereign ratings on GIIPS spreads is likely to have changed. Simulating this regime change would add unnecessary complexity to the model.

Therefore, this paper uses fiscal and macroeconomics fundamentals as the preferred measure of credit risk. Fiscal fundamentals include the debt to GDP ratio and the budget surplus to GDP ratio. Macroeconomic fundamentals include real GDP growth and the current account balance as share of GDP. The first three variables are clearly indicators of a country's ability to repay its debt. The current account balance has played only a minor role in the existing literature on Euro spreads. It is however commonly used in research about emerging markets and imbalances between Euro area countries have been publicly identified as one of the main drivers of the crisis. The relevance of those indicators is emphasized by the fact that all four are part of the European Commission's Macroeconomic Imbalance Procedure. The data is obtained on a quarterly basis from Eurostat. For the lack of detailed release dates it is worth noting that budget data is released with a 3-4 month delay whereas GDP numbers are released with a 1-2 month delay ("preliminary flash" and "flash" estimates) and current account numbers ("1st release") with a 2 month delay. This paper therefore assumes that yields react immediately to liquidity, uncertainty and risk aversion, with a two-month lag to macroeconomic variables and with a four-month lag to budget numbers. This conveniently also addresses the risk of possible reverse causality between borrowing costs and fundamentals.

Table 3: Summary statistics - Fiscal & macro fundamentals

Variable	Mean	Std. Dev.	Min.	Max.	N
		GDP growth			
Spain	0.305	0.702	-1.561	1.131	55
Portugal	0.007	0.763	-2.302	1.109	55
Netherlands	0.284	0.721	-3.238	1.482	55
Italy	-0.031	0.716	-2.86	1.155	55
Ireland	1.001	3.322	-4.167	20.342	55
Greece	-0.191	1.65	-4.762	3.11	55
France	0.273	0.493	-1.705	1.091	55
Finland	0.244	1.332	-6.814	2.764	55
Belgium	0.346	0.570	-2.102	1.575	55
Austria	0.333	0.727	-1.766	1.835	55
		Current account surplus			
Spain	-3.976	4.074	-11.6	3	55
Portugal	-6.42	4.964	-13.6	3.9	55
Netherlands	7.582	2.308	2.2	12.7	55
Italy	-0.858	2.091	-5.8	3.5	55
Ireland	-0.965	5.196	-10	14.3	55
Greece	-8.353	6.923	-21	11.9	55
France	-0.407	1.115	-2.9	1.7	55
Finland	1.805	3.641	-4.2	10.4	55
Belgium	0.882	3.684	-7.7	9.4	55
Austria	2.582	2.346	-1.1	8.800	55
		Budget surplus ratio			
Spain	-4.154	5.953	-18.2	8	56
Portugal	-5.843	3.594	-18.5	0.1	56
Netherlands	-2.243	2.686	-8	4	56
Italy	-3.327	2.494	-9.4	2.2	56
Ireland	-5.648	9.914	-42	8.5	56
Greece	-8.652	5.207	-30.1	-0.8	56
France	-4.125	2.883	-10.3	1	56
Finland	0.475	4.773	-9.6	10.8	56
Belgium	-2.509	5.917	-17.9	6.7	56
Austria	-2.571	2.673	-9.5	2.8	56
		Gross debt ratio			
Spain	61.189	22.775	34.7	100.5	56
Portugal	89.539	29.035	53.2	132.8	56
Netherlands	56.098	8.495	42.7	69.100	56
Italy	113.736	12.038	99.8	136	56
Ireland	63.788	37.5	23.6	124	56
Greece	131.866	30.385	101.5	181.8	56
France	77.107	12.988	59.5	97.7	56
Finland	44.786	9.227	28.8	63.6	56
Belgium	101.754	6.369	87	110.9	56
Austria	76.009	7.029	64.8	86.900	56

As a robustness check this paper also considers the European Economic Forecast by the DG for Economic and Financial Affairs which is published in the spring, autumn and, since 2013, also in the winter of every year. From each release this paper uses the forecasts for the upcoming year (i.e. the 2016 forecast comes from the winter, spring and autumn 2015 release) for GDP growth, the primary surplus to GDP ratio and the gross debt to GDP ratio. Each forecast is considered the best available forecast until the day of the next release. The advantage of forecasts is that markets in theory should be forward looking and therefore should pay more attention to forecasts than to historical values. Furthermore, contrary to the official data, the exact release dates of the forecasts are known. The main disadvantage is that the number of observations available is much lower.

Table 4: Summary statistics - Forecasts

Variable	**Mean**	**Std. Dev.**	**Min.**	**Max.**	**N**
		GDP growth forecast			
Spain	1.741	1.389	-1.4	3.4	32
Portugal	0.897	1.24	-3	2.2	32
Netherlands	1.553	1.033	-0.4	5	32
Italy	1.128	0.612	-0.5	2.1	32
Ireland	2.888	1.931	-2.6	5.3	32
Greece	1.803	2.286	-4.2	4.2	32
France	1.538	0.647	-0.2	2.4	32
Finland	1.878	0.953	0.2	3.5	32
Belgium	1.572	0.661	-0.2	2.5	32
Austria	1.716	0.585	-0.1	2.7	32
		Primary budget surplus forecast			
Spain	-0.941	3.063	-7.8	2.9	32
Portugal	0.037	1.823	-4.9	2	32
Netherlands	-0.097	1.792	-3.7	3.7	32
Italy	2.022	1.198	-0.6	4.5	32
Ireland	-1.197	3.835	-12.5	2.1	32
Greece	1.369	2.537	-6.600	5.4	32
France	-1.391	1.391	-5.4	0.5	32
Finland	1.313	2.826	-3.1	5.9	32
Belgium	1.369	2.199	-2.1	5.7	32
Austria	0.716	1.223	-2.5	3	32
		Gross debt ratio forecast			
Spain	70.553	26.686	34.6	104.8	32
Portugal	91.159	30.021	23.5	127.2	32
Netherlands	60.638	11.873	29.3	76.400	32
Italy	118.172	11.84	102.6	134	32
Ireland	75.575	39.959	21.7	122.5	32
Greece	139.331	37.846	90.2	199.7	32
France	81.669	14.808	59.3	99.8	32
Finland	52.084	15.679	29.1	95.900	32
Belgium	94.478	18.796	0.8	107.8	32
Austria	71.309	8.925	56.8	85.8	32

Liquidity risk is measured by the spread between closing bid and ask prices because the literature suggests that deeper bond markets decrease bid-ask spreads (e.g., BIS, 1999). Bid and ask prices are obtained as daily closing prices from Bloomberg. Unfortunately, the data provided by Bloomberg has some gaps, so the periods affected will not be estimated. Apart from Greece, bid-ask spreads on average are quite low, with Portugal having the highest average bid-ask spread of 5.4 basis points. This indicates a well-integrated financial market.

Table 5: Summary statistics - Liquidity

Variable	Mean	Std. Dev.	Min.	Max.	N
Spain	0.008	0.01	-0.199	0.066	3572
Portugal	0.054	0.105	-0.229	0.789	3392
Netherlands	0.005	0.003	0	0.042	3548
Italy	0.01	0.006	0	0.079	3574
Ireland	0.022	0.049	0	0.294	3044
Greece	0.292	0.833	0.002	7.082	3560
France	0.006	0.004	0	0.037	3574
Finland	0.004	0.003	0.001	0.044	3575
Belgium	0.007	0.006	0	0.052	3574
Austria	0.011	0.012	-0.016	0.062	3533

The country specific variables presented here are used as absolutes and not relative to Germany. It seems likely that Germany enjoys such a high level of credibility that investors make their decisions based on a country's budget deficit and not relatively to Germany. Indeed, Santis (2012) finds no significant difference between both approaches.

3.4 Monetary Policy

This paper uses the real interest rate as the main measures of the monetary policy stance. The real rate is defined as the nominal rate (i.e. the ECB's main refinancing operation (MRO), irrespective of the nature of the tender) minus the harmonised consumer price index (HCPI) 12-month average rate of change. The nominal rate is obtained from the ECB whereas the HCPI is obtained from Eurostat. The real rate is a better measure of the monetary policy stance than the nominal rate because it accounts for inflation: A nominal rate of, for example, 2 percent can be considered either tight or loose, depending on the rate of inflation. Figure 3 illustrates both the nominal and the real interest rate series.

In Section 4.5 alternative measures are used as robustness check: the nominal rate and unexpected moves in the ECB's monetary policy, including the announcements of unconventional policy measures. Following Gürkaynak, Sack and Swanson (2005), a monetary policy surprise is defined

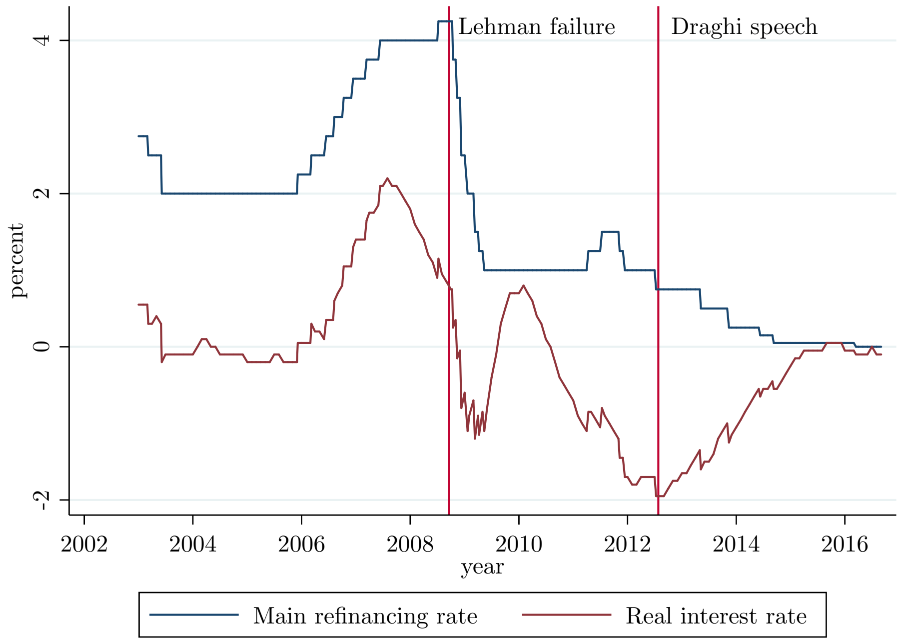

Figure 3: Monetary policy stance in the Eurozone

as the movement of interest rate futures around the time of a monetary policy decision. Intuitively, prices of interest rate futures should reflect the best available prediction of the future monetary policy stance. Any change of that price in response to a monetary policy announcement implies that the announcement caught markets by surprise, thereby solving the orthogonalization problem frequently arising in identifying the stance of monetary policy. Following Bernoth and von Hagen (2004) the appropriate future in this case is the continuous three-month Euribor obtained from Thomson Reuters Datastream. An unexpected monetary policy decision therefore is defined as

$$\Delta r_t^u = f_t - f_{t-1} \tag{3}$$

where Δr_t^u is the unexpected change in interest rates at time t and $f_t - f_{t-1}$ is the difference between the rate implied by the spot prices of the continuous three-month Euribor future at time t and time $t-1$.[9] As in Bredin, Hyde, Nitzsche and O'Reilly (2009) a daily window is used. The expected rate

[9] The implied rate is defined as 100 minus the futures spot price, multiplied by 100.

change conversely is defined reversely as

$$\Delta r_t^e = \Delta r_t - \Delta r_t^u \tag{4}$$

Policy surprises were calculated on each day of a regular ECB governing council meeting. This was also done when the MRO was left unchanged, because even this could come as a surprise to market expectations. The meetings of the governing council currently take place eight times a year, before 2015 twelve times a year. Besides the normal meetings the dates of unconventional monetary policy announcements were also considered. Unlike interest rates this approach also captures the market reaction to the announcement of, for example, asset purchase programs: In many theoretical models they should affect real activity but are not reflected in the policy rate. The relevant dates are aggregated from Kilponen et al. (2012), Szczerbowicz (2015) as well as Dewachter et al. (2016) and presented in Table A9. When an announcement was made over the weekend or after markets closed, the following business day was used. Figure 4 presents the estimated 180 monetary policy shocks (of which 153 are different from zero). It is notable that the frequency and magnitude of policy shocks increased during the crisis period. Since 2012 however, shortly after Mario Draghi took office, the frequency and magnitude of shocks significantly decreased again, even below their pre-crisis levels. This could imply that the ECB became better at communicating its strategy to the markets.

Both the real and nominal interest rate series follow a unit root process ($Z_{\rho,rate} = -4.609; Z_{\rho,realrate} = -4.667$) and therefore enter in first differences. The series of monetary policy surprises is stationary ($Z_{\rho,policyshock} = -3,563.601$) and therefore enters in levels.

4 The Empirical Model

Visual inspection of Figure 1 suggests that the spreads between German and other European sovereign bond yields are highly persistent during the crisis years. As previous literature suggests, this requires to test for the presence of unit roots and cointegration relationships. Furthermore, even stationary but highly autocorrelated variables can produce spurious relationships, as Granger, Hyung and Jeon (2001) point out. Following the approach of Giordano et al. (2013) this paper first estimates a stationary model as this is assumed by most of the early literature on that topic (Section 4.1). However, after showing that this assumption is questionable (Section 4.2), this paper drops the assumption of stationarity and corrects the model accordingly

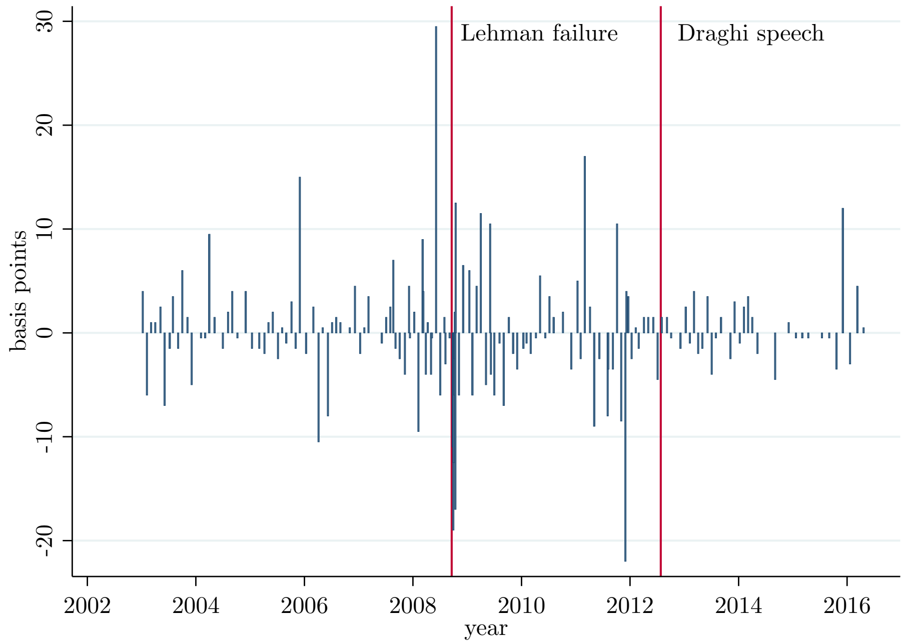

Figure 4: Monetary policy shocks in the Eurozone

(Section 4.3). This model is also estimated on a rolling basis to check for any variation in the parameters over time (Section 4.4). Finally, a range of alternative model specifications is estimated to test the robustness of the model (Section 4.5).

4.1 Stationary Model

The model is estimated assuming stationary at first. It allows for heteroskedasticity and autocorrelation in the residuals but corrects for heteroskedasticity and autocorrelation in the standard errors using Newey-West (1987) standard errors with eight lags. In this equation, i and t denote country and day, respectively.

$$\begin{aligned} y_{i,t} - y_{G,t} = \alpha_i &+ \beta_{1,i} GROWTH_{i,t} + \beta_{2,i} ACCOUNT_{i,t} \\ &+ \beta_{3,i} BUDGET_{i,t} + \beta_{4,i} DEBT_{i,t} + \beta_{5,i} LIQUIDITY i,t \\ &+ \beta_{6,i} UNCERTAINTY_t + \beta_{7,i} RISKAVERSION_t + \epsilon_{i,t} \end{aligned} \tag{5}$$

where $y_{i,t} - y_{G,t}$ is the spread between bond yields of country i and German yields, $GROWTH_{i,t}$ the is the lagged real GDP growth, $ACCOUNT_{i,t}$ the lagged current account balance as a share of GDP, $BUDGET_{i,t}$ the lagged budget surplus as share of GDP, $DEBT_{i,t}$ the lagged gross debt as share of GDP, $LIQUIDITY_{i,t}$ the daily bid ask spread and $UNCERTAINTY_t$ and $RISKAVERSION_t$ are the respective series constructed in Section 3.1. This specification is relatively close to that of Giordano et al. (2013), only that this paper replaces the private debt ratio with the budget surplus ratio and decomposes the VIX into an uncertainty and risk aversion component. Furthermore, unlike in Giordano et al. (2013), the model is estimated for each country separately and not as a panel.

The first three coefficients of this model are expected to have a negative sign as high growth rates, current account and budget surpluses are generally considered to be a sign of a healthy economy which makes it easier to repay public debt and thereby should lower spreads. For the same reason, the coefficient on $DEBT_{i,t}$ is expected to be positive as high debt levels impede a country's ability to repay its debts and are associated with lower growth (e.g., Woo & Kumar, 2015). The last three coefficients are expected to have a positive sign, as higher bid-ask spreads (i.e. lower liquidity), uncertainty and risk aversion all decrease the attractiveness of an asset and, therefore, should increase spreads.

In order to deal with the presence of mixed sample frequencies, the lower-frequency data of economic and fiscal fundamentals is brought to the highest available frequency (i.e. that of the spread series) by simply keeping the values constant until new data is released. The intuition behind this is that the available information set does not significantly change without the release of new official data. This approach is preferred over aggregating up the daily data to the lowest available frequency in order to avoid losing much of the variation in the uncertainty and risk aversion series. Furthermore, most of the changes in monetary policy take place during the month, so that by aggregating the data to the lowest frequency available it would not be possible to detect the immediate effect of monetary policy on risk aversion. Other techniques that attempt to convert lower frequency variables into higher frequency variables by modelling them as latent variables (e.g., Mariano & Murasawa, 2003) were not used because they would impute that markets have the same information set. While the use of constant variables also imputes a certain information set, i.e. that markets do not incorporate other information about the fiscal position of a country until the next data release, it seems more likely that markets follow the official data rather than use an information set imputed by a complex disaggregation of low frequency data.

Part I of Table 6 presents the results of this regression. In a second step,

Table 6: Regression results - Stationary model

	ES	PT	NL	IT	IR	GR	FR	FN	BE	OE
					Part I: Baseline without instrumenting for risk aversion					
Constant	-1.281***	-3.928***	-.512***	-10.792***	-.587***	-11.041***	-1.275***	.067	-2.118***	-1.823***
GDP Growth	-.944***	-.390***	-.047***	-.433***	-.024***	-.851***	-.090***	-.019***	-.133***	.006
Account balance	-.062**	-.084***	.017***	-.244***	-.046***	-.077***	-.051***	-.015***	-.027***	.012***
Budget surplus	-.026***	-.026*	.004	.030**	-.074***	-.020	.002	-.009***	-.009**	-.003
Debt ratio	.035***	.051***	.008***	.100***	.013***	.107***	.018***	-.0004	.022***	.024***
Liquidity	41.933**	20.386***	15.917***	62.800***	19.215***	5.551***	21.590***	16.290***	54.545***	13.980***
Euro Uncertainty	.002	.001	.001**	.003	.002	-.007	-.0002	.001**	.003	.001
Euro Risk aversion	-.006**	-.001	.002**	-.001	.004	.017**	.003***	.002***	-.001	.002***
					Part II: Instrumenting for risk aversion					
Constant	-2.134***	-3.836***	-.510***	-11.182***	-.209	-9.795***	-1.306***	.077	-2.049***	-1.859***
GDP Growth	-.6463***	-.193	-.018	-.246	-.017	-.681***	-.053*	-.013**	-.056	.016
Account balance	-.1301**	-.094***	.027***	-.231***	-.053***	-.098**	-.057***	-.015***	-.027***	.013***
Budget surplus	-.0129	.004	.006**	.033*	-.054	.084	.003	-.010***	-.011***	-.004
Gross Debt to GDP	.054***	.057***	.008***	.105***	.017***	.127***	.0192***	.000	.022***	.026***
Liquidity	54.362***	20.669***	11.601***	67.929***	21.764***	5.457***	22.766***	15.930***	51.703***	11.555***
Euro Uncertainty	-.115*	-.099*	-.013*	-.067	-.081	-.469	-.011	-.006	-.021	-.012
$Euro\widehat{Risk}aversion$	.147*	.128*	.020**	.091*	.111	.607	.018*	.011	.030	.020**

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

the risk aversion term is instrumented by the real interest rate, the monetary policy variable. It could be argued that this approach suffers from the problem that risk aversion and the stance of monetary policy are endogenously determined: Monetary policy might affect asset prices via risk aversion but monetary policy in turn also might react to asset prices. That monetary policy indeed reacts to asset prices was shown, for example, by Rigobon and Sack (2004) for the Federal Reserve. For this reason, Bekaert et al. (2013) opt to use a vector autoregressive system. Unlike their model however, this paper uses daily instead of monthly frequency data. It seems unlikely that monetary policy reacts to such short-term changes in risk aversion or uncertainty. Furthermore, Bekaert et al. (2013) find that monetary policy is largely unresponsive to risk aversion even with monthly data when controlling for business cycle movements. Table 6 Part II presents the results of the instrumental variable regression. The coefficients of the stationary model without instrumenting risk aversion with monetary policy by and large have the expected signs, with some exceptions for the current account surplus (the Netherlands, Austria) and the budget surplus (Italy). GDP growth, the gross debt ratio and liquidity are almost always significant and have the expected signs. Uncertainty is only (slightly) significant for the Netherlands and Finland, whereas Risk aversion is significant for Spain, the Netherlands, Greece, France, Finland and Austria, all except Spain with the expected positive sign.

Several coefficients on GDP growth and the budget balance become insignificant after instrumenting risk aversion. This suggests possible issues of multicollinearity between the instrument and those variables. The signs of the coefficients however stay the same. In terms of uncertainty, the coefficient on Austria becomes insignificant while the coefficients on Spain, Portugal and the Netherlands all are slightly significant with an unexpected negative sign. The instrumented risk aversion variable has a (slightly) significant impact in Spain, Portugal, the Netherlands, Italy and Austria, all with the expected positive sign.

4.2 Testing for Unit Roots and Cointegration

Some parts of the previous literature raise the possibility of non-stationarity among the variables. This paper therefore uses the augmented-Dickey Fuller (ADF) test and the Phillips-Perron test. The ADF test uses additional lags of the first-differenced variable whereas the Philips-Perron test uses Newey-West (1987) standard errors to account for serial correlation as well as heteroskedasticity. The appropriate lag length for the ADF was chosen using

Table 7: Unit root tests

	ES	PT	NL	IT	IR	GR	FR	FN	BE	OE
Spreads										
- Z_ρ	-4.341	-4.116	-16.499	-5.167	-3.894	-9.334	-11.644	-14.947	-8.713	-13.187
- ADF	-1.611	-1.561	-3.021	-1.756	-1.514	-2.016	-2.595	-2.927	-2.235	-2.865
GDP growth										
- Z_ρ	-7.085	-31.868	-32.528	-19.166	-66.409	-43.177	-25.574	-42.054	-21.999	-40.509
- ADF	-1.557	-4.619	-4.668	-3.116	-7.596	-1.914	-3.817	-5.458	-3.472	-5.263
Current account balance										
- Z_ρ	-2.896	-3.428	-15.616	-15.675	-2.418	-32.128	-45.801	-22.366	-56.553	-37.412
- ADF	0.002	0.285	-3.526	0.456	1.914	0.676	-1.875	-1.720	-2.077	-1.408
Budget surplus										
- Z_ρ	-49.132	-57.261	-25.925	-65.217	-20.763	-51.619	-95.895	-26.291	-107.268	-60.960
- ADF	-0.930	-6.698	-1.162	-1.878	-1.267	-6.351	-1.033	-0.862	-1.802	-7.912
Gross debt ratio										
- Z_ρ	0.992	-0.152	-1.361	0.964	-1.175	-0.169	0.025	0.975	-3.320	-2.189
- ADF	0.205	-0.087	-0.598	0.469	-1.947	-0.255	-0.191	0.919	-1.114	-1.326
Liquidity										
- Z_ρ	-443.596	-56.434	-145.483	-70.446	-8.443	-65.674	-88.302	-173.908	-66.689	-142.213
- ADF	-14.485	-7.883	-9.129	-7.655	-2.123	-7.042	-6.522	-8.634	-6.054	-10.375
Uncertainty										
- Z_ρ					-94.934					
- ADF					-6.242					
Risk aversion										
- Z_ρ					-112.395					
- ADF					-6.140					

Critical values for the Philips-Perron test for daily/quarterly data are -11.300/-10.730 (10%), -14.100/-13.340 (5%) and -20.700/-18.990 (1%) respectively. Critical values for the ADF test for daily/quarterly data are -2.570/-2.598 (10%), -2.860/-2.926 (5%) and -3.430/-3.573 (1%) respectively.

Table 8: Engle-Granger tests for cointegration

Variable	ES	PT	NL	IT	IR	GR	FR	FN	BE	OE
Section I: Spreads as dependent variable										
GDP growth	-2.324	-2.484	-3.941	-2.080	-1.733	-3.712	-2.631	-3.765	-2.310	-3.219
Current account balance	-2.078	-1.754	-2.911	-1.652	-1.452	-2.413	-3.201	-3.588	-2.800	-2.622
Budget surplus	-4.124	-2.169	-3.011	-1.666	-4.446	-2.250	-2.797	-3.061	-2.326	-2.594
Gross debt ratio	-1.578	-1.675	-3.078	-1.752	-1.565	-2.809	-2.850	-2.859	-2.134	-2.750
Section II: Spreads as explanatory variable										
GDP growth	-2.346	-4.547	-4.879	-3.243	-5.640	-5.399	-3.711	-5.202	-3.387	-4.877
Current account balance	-2.132	-2.033	-3.278	-3.132	-1.428	-4.426	-5.108	-4.548	-5.804	-5.086
Budget surplus	-6.046	-5.451	-4.002	-6.264	-5.473	-5.016	-6.516	-4.350	-7.315	-5.636
Gross debt ratio	0.225	-0.825	-1.143	-0.282	-0.941	-1.725	-1.419	0.171	-1.610	-1.618

The critical values for the Z(t) statistic with an intercept in the estimated cointegrating relationship but no intercept in the error term regression with 1 explanatory variable in the first stage regression (excl. intercept) correspond to case 2 in Table B.9, pg. 766 Hamilton (1994) and were adjusted using the Bonferroni correction: -3.37 (10 percent), -3.64 (5 percent).

the Schwartz Bayesian information criterion due to its superior small sample properties. The number of Newey-West lags for the spreads, uncertainty and risk aversion series was set to 9, for the liquidity series to 8 and for the quarterly series to 3. Table 7 presents the augmented Dickey-Fuller Z_t statistic and the Phillips-Perron Z_ρ statistic.[10] The null hypothesis of a unit root cannot be rejected for seven out of ten series of spreads (except for the Netherlands, Finland and partially Austria) and all series of gross debt. It can be rejected for risk aversion, uncertainty, and in all but one case for liquidity (Ireland) and GDP growth (Spain). The test results for the current account balance and the budget surplus are more ambiguous: Using the augmented Dickey-Fuller test, the null of a unit root is only rejected for the current account balance of the Netherlands and only for the budget surplus ratios of Portugal, Greece and Austria. The Philips-Perron test on the other hand rejects the null in seven out of ten cases for the current account balance and in all series of budget surplus ratios. However, both tests are known for their poor finite sample performance. Especially the quarterly series of fiscal and macroeconomic fundamentals with a sample size of $N = 55$ are likely to be subject to those deficiencies.

For this reason, the modified Dickey-Fuller test is also used. The modified Dickey-Fuller test, as proposed by Elliott, Rothenberg and Stock (1996, from here on called ERS) performs the ADF test on a series that has been transformed by a generalized least squares regression. The results of the ERS test are not reported here due to space reasons but are available in the Stata documentation of this paper. The results of the ERS test are much more pronounced than the previous results: For all spread and credit risk variables the null hypothesis cannot be rejected at some lag length (using a 5% significance threshold). It is striking that unit roots appear to be present not only in the full sample but also when considering only the pre-crisis sample. In case of spreads the results of the ERS test on the pre-crisis sample are even more suggestive of unit root processes than before. For the liquidity, risk aversion and uncertainty series the ERS test rejects the null hypothesis of a unit root process just like the previous tests.

Following the results of the unit root tests it is necessary to test for cointegration relationships between the spreads and their determinants. This paper uses the residual-based Engle-Granger test as laid out in Hamilton (1994, p. 599): In the first step an OLS regression involving the potentially cointegrated variables is estimated and in the second step the saved

[10] In case of the quarterly variables both tests were conducted before interpolating the data.

residuals are regressed onto their lagged values. The Engle-Granger test unfortunately does not reveal the number of cointegration relationships, instead it only indicates whether a cointegration relationship exists. Some authors try to circumvent that problem by adding one variable at a time to the first stage regression, each time testing the residual series for stationarity. However, adding an independent variable to a cointegrating pair of variables will not change the outcome of the test. If the third variable is not part of the cointegration relationship its coefficient will be put to zero with the residual series being unchanged. Adding a third variable to a cointegrating relationship of two variables therefore runs at risk of indicating a cointegration relationship with the third variable even if in reality there is none (Sjö, 2008). Therefore, only bivariate cointegrating regressions are tested.

The critical values for the null hypothesis of no cointegration (i.e. a unit root in the regression residuals) differ from the critical values of normal unit root tests as OLS is biased towards creating stationary residuals. The appropriate critical values are higher (in absolute terms) than those of normal unit root tests and depend on the number of explanatory variables in the first regression as well as the existence of deterministic terms (i.e. drift and trend) in the first and second stage regressions. Furthermore, since both possible orders of the first stage regressions are tested, the Bonferroni correction is applied to the critical values to counteract the problem of multiple comparisons (Sandberg, 2016): Each individual hypothesis is tested at the significance level $p = \alpha/m$ where α is the desired alpha and m is the number of hypotheses tested ($m = 2$ in this case). The resulting critical values for two hypotheses at the 10 and 5 percent threshold conveniently correspond to the critical values for 5 and 2.5 percent, respectively. Table 8 Part I displays the Philips-Perron Z_t statistic for the presence of a unit root with six Newey-West (1987) lags.

Most residual series follow a unit root and therefore fail to reject the null hypothesis of no cointegration. However, there are a few notable exceptions: The Engle-Granger test indicates a cointegration relationship between Spanish and Irish sovereign spreads and their respective budget surplus as well as between Dutch, Greek and Finnish spreads and GDP growth when using the conventional 5 percent significance threshold. When using the loser 10 percent threshold Finnish spreads and the current account balance also appear to be cointegrated.

The results change when the ordering of the first stage regression is reversed: The number of cointegration relationships increases drastically (see Table 8 Part II). In all countries but Spain, Italy and Belgium the test indicates a cointegration relationship between spreads and GDP growth, in all but Spain, Portugal, the Netherlands, Italy and Ireland between spreads and the current account surplus and in all countries between spreads and the budget surplus. Only in case of the gross debt ratio the number of cointegration relationships is zero regardless of the normalisation.

This is a well-known deficiency of the Engle-Granger test, which can occur if there are large differences in the variances of the variables or if the variables are near-integrated (i.e. have a large negative MA(1) component) (Sjö, 2008). Unfortunately, there is no obvious solution to this problem. However, based on economic theory, it seems likely that a long-run relation between a country's yields vis-à-vis Germany and its fundamentals exists, i.e. that spreads return to a level that is justified by the respective country's economic fundamentals. This paper therefore proceeds by treating the spread and the non-stationary variables as cointegrated.

4.3 Non-Stationary Model

Given the outcome of the Engle-Granger test displayed in Table 8, some of the estimated coefficients presented in Table 6 might be spurious while some might be unbiased but their respective standard errors cannot be used for inference. The baseline model therefore must be corrected accordingly. The obvious solution would be to estimate an error correction model (ECM). An ECM can include stationary series, non-stationary series and cointegrated non-stationary series, but only as long as at least one cointegration relationship is present: By adding leads and lags of the non-stationary variables in first differences, any spurious correlation will be avoided (e.g., Enns, Masaki & Kelly, 2014). However, given the limited variation in the quarterly fiscal and economic fundamentals, using first differences to estimate the ECM would reduce the explanatory power of the model. Therefore, this paper instead estimates a dynamic ordinary least squares (DOLS) model as proposed by Stock and Watson (1993).

The DOLS method augments the cointegrating regression with leads and lags of the non-stationary variables so that the error term is orthogonal to all past changes of the explanatory variables. Newey-West (1987) standard errors are used to correct for heteroskedasticity and serial correlation in the error terms, ensuring that the t-statistics follow the normal asymptotic distribution. Even though the DOLS model has been used in the same fashion as the ECM, i.e. including cointegrated and non-cointegrated non-stationary series (e.g., Santis, 2012), to the best of the author's knowledge, it has been never explicitly shown to bee econometrically appropriate. Therefore, as part of this paper, a simple Monte Carlo simulation modelled after Enns et al. (2014) was conducted to confirm this. The results are documented in Section A3.

$$
\begin{aligned}
y_{i,t} - y_{G,t} = \alpha_i &+ \beta_{1,i} GROWTH_{i,t} + \sum_{j=-L}^{F} \gamma_{1,i,j} \Delta GROWTH_{i,t+j} + \beta_{2,i} ACCOUNT_{i,t} \\
&+ \sum_{j=-L}^{F} \gamma_{2,i,j} \Delta ACCOUNT_{i,t+j} + \beta_{3,i} BUDGET_{i,t} + \sum_{j=-L}^{F} \gamma_{3,i,j} \Delta BUDGET_{i,t+j} \\
&+ \beta_{4,4} DEBT_{i,t} + \sum_{j=-L}^{F} \gamma_{4,i,j} \Delta DEBT_{i,t+j} + \beta_{5,i} LIQUIDITY_{i,t} \\
&+ \beta_{6,i} UNCERTAINTY_t + \beta_{7,i} RISKAVERSION_t + \epsilon_{i,t}
\end{aligned}
\tag{6}
$$

The DOLS model includes six lags and leads for each non-stationary variable (i.e. $L = F = 6$), to obtain comparable results for each country. As in the baseline regression, the number of Newey West lags correcting for heteroskedasticity and serial correlation is set to eight. Table 9 presents the corresponding results. Since the leads and lags in differences only serve to avoid any spurious correlation and enable robust inference, they are not reported. In this DOLS framework the coefficients on the cointegrated variables represent the long-run adjustment whereas the other coefficients represent the short-run dynamics of the model.

Figure 5 displays the residuals of the regression presented in Table 9 and shows that they are indeed stationary and highly mean-reverting. It is furthermore worth noting, that the estimated coefficients of the DOLS model are very close to the stationary model in Table 6 but with slightly different significance levels. If there was no cointegrating relationship, the DOLS regression would indicate insignificant coefficients for the non-cointegrated variables. Both observations taken together suggest that the hypothesis of Section 4.2 regarding the presence of a cointegration relationship was indeed correct.

Almost all of the credit risk variables have the expected sign when significant: The coefficients on GDP growth are all negative, indicating that higher growth rates lead to lower spreads. Only two countries do not have the expected sign on the current account balance, the Netherlands and Austria. Those two countries were among the countries with the highest current account surpluses in recent years. This could suggest that in those countries a further increase is seen as sign of a weak domestic economy with fewer imports. Similarly, unlike in all the other countries, markets might believe that lower budget surpluses in Italy might lead to higher GDP growth which would motivate the positive coefficient. In case of the public gross debt ra-

Table 9: Dynamic OLS estimates

	ES	PT	NL	IT	IR	GR	FR	FN	BE	OE
					Part I: i Baseline without instrumenting for risk aversion					
Intercept	-1.248***	-3.937***	-0.493***	-11.038***	-0.576***	-10.879***	-1.276***	0.095**	-2.084***	-1.875***
GDP growth	-0.923***	-0.398***	-0.053***	-0.463***	-0.023***	-0.968***	-0.091***	-0.020***	-0.108***	0.009
Current account	-0.060***	-0.082***	0.019***	-0.265***	-0.044***	-0.080***	-0.057***	-0.017***	-0.037***	0.015***
Budget surplus	-0.032***	-0.033**	0.002	0.032**	-0.079***	-0.020	0.001	-0.009***	-0.013***	-0.005***
Gross debt ratio	0.034***	0.050***	0.007***	0.101***	0.013***	0.106***	0.018***	-0.001	0.021***	0.025
Liquidity	41.759**	20.281***	15.184***	61.948***	18.535***	5.513***	21.963***	15.280***	54.267***	14.097***
Uncertainty	0.002	0.001	0.001**	0.002	0.002	-0.009	0.000	0.001***	0.002	0.001
Risk aversion	-0.007***	-0.002	0.002**	-0.001	0.003	0.018**	0.004***	0.002***	0.000	0.003***
					Part II: Instrumenting for risk aversion					
Intercept	-2.363***	-3.967***	-.501***	-11.363***	-.267	-9.966***	-1.300***	.093*	-2.046***	-1.937***
GDP growth	-.567**	-.187	-.022	-.276*	-.018	-.815***	-.054*	-.013**	-.021	.024
Current account	-.146**	-.102***	.028***	-.254***	-.050***	-.106**	-.065***	-.017***	-.038***	.018***
Budget surplus	-.022	.006	.005	.036**	-.062**	.070	.003	-.010***	-.015***	-.007**
Gross debt ratio	.055***	.059***	.008***	.107***	.016***	.124***	.019***	.000	.022***	.026***
Liquidity	54.282***	20.676***	11.460***	66.016***	20.751***	5.440***	22.963***	15.122***	51.468***	11.958***
Euro Uncertainty	-.117*	-.102	-.012*	-.063	-.064	-.403	-.010	-.006	-.021	-.012*
$\widehat{EuroRiskaversion}$	.149*	.131	.019**	.083	.088	.520	.017*	.011	.030	.019**

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

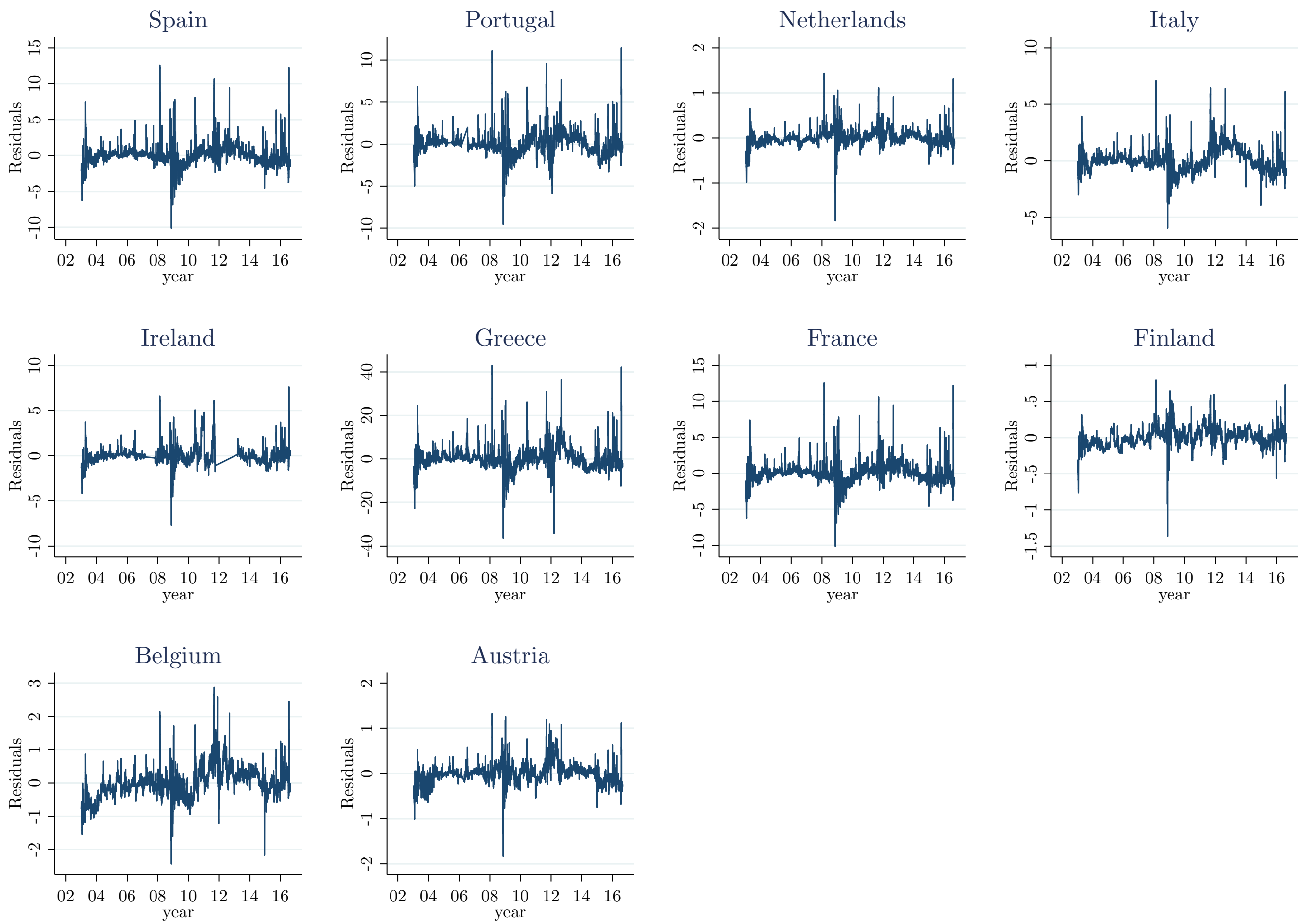

Figure 5: Residuals of the DOLS regression

tio all significant coefficients have the expected positive sign, implying that higher debt rates uniformly lead to higher spreads. The only outlier here is that markets do not seem to pay attention to the Finnish gross debt level, perhaps because for most of the time Finland had the lowest debt level of all countries, only rivalled by Ireland in the pre-crisis period. The bid-ask spread is highly significant for all countries and carries the expected positive sign, implying that higher bid-ask spreads (i.e. lower liquidity) leads to higher spreads. It is perhaps worth to notice the significant coefficients on a range of variables which were previously indicated to be non-stationary but not part of a cointegration relationship (in particular current account balances and gross debt ratios). Since the DOLS model should avoid any spurious correlations and Figure 5 confirms that the residuals of the model are indeed stationary, this result hints at insufficiencies in the unit root and cointegration tests which were used.

Regarding to the main variables of interest, uncertainty and risk aversion, the latter seems to be relatively more important: Uncertainty is only significant for the Netherlands and Finland. In those two cases the coefficients have the expected positive sign. Risk aversion on the other hand is significant in six out of ten countries, mostly with the expected positive sign. Only Spain is an outlier, with a highly statistically significant negative coefficient. This result is highly counter-intuitive as this implies that with growing risk aversion investors bought more Spanish government bonds, a behaviour not seen in safer countries like the Netherlands, France, Finland or Austria. One possible explanation is that increasing risk aversion quickly led to a relatively quick European policy response due to Spain's systemic importance, which drove down yields again.

Instrumenting risk aversion with the real interest rate does not change any coefficient sign on the country specific variables but renders some insignificant, in particular the coefficients on GDP growth (Portugal, Italy, Ireland, Belgium) and budget surplus (Spain, Portugal). The coefficients on the current account balance, the gross debt ratio and liquidity remain mostly unaffected. This could be explained by the fact that real interest rates and the budget surplus closely track the business cycle and instrumenting risk aversion with the real interest rates results in multicollinearity. This increases the standard deviation of those coefficients and thus makes them less significant. Instrumenting risk aversion does have a significant effect on the uncertainty estimates, rendering one insignificant (Finland), reversing the sign of one coefficient (Netherlands) and making two coefficients slightly significant which were not before (Spain and Austria). The significant uncertainty coefficients in the instrumental regression now all have the expected positive sign. The instrumented series itself also experiences some changes: In three cases the coefficients become insignificant (Greece, France, Finland) or less significant (Austria) and the coefficient on risk aversion in Spain now has the expected positive sign.

4.4 Rolling Estimation

In order to account for the likely fact that the relation between the explanatory variables and the spreads has not remained constant over the almost 14 years under observation, this paper conducts a rolling estimation of the DOLS model. The rolling estimation is conducted using the baseline model without instrumenting monetary policy because the coefficient on the instrumented variable has significantly higher standard errors, making any visual inspection of the rolling estimates nearly impossible. The stability of the coefficients from the instrumental variable regression will instead be tested in the robustness check section. The model is estimated on a window of 500 observations, roughly corresponding to two years, moving forward in steps of 20 observations at a time, which roughly corresponds to four weeks. This yields a total of 154 observations for each coefficient. However, due to space restrictions only the coefficients for uncertainty and risk aversion will be discussed. It is worth noting that the coefficients displayed in Figure 6 & 7 refer to the preceding two years - the first observation therefore is the estimated coefficient for the period between 2003 - 2005.

Figure 6 shows that the estimated coefficient for uncertainty in the Eurozone is relatively small and mostly insignificant during the pre-crisis period, except for Finland and Austria which have small but significantly positive coefficients on uncertainty during 2005 and 2007 (Finland)/2008 (Austria). The estimated coefficients become increasingly volatile with the onset of the crisis, but they are rarely significantly different from zero. Portugal (2012-2013), the Netherlands (2009-20013), Ireland (2013), Greece (2014), Finland (2012) and Belgium (2009-2013) are the exceptions to that. The Italian coefficient is particularly volatile and is significantly positive for various short intervals between 2011 and 2014. After 2014 a general downward trend is recognizable (significant for the Netherlands, Belgium and Austria), indicating that countries slowly reclaim their status as safe haven to economic uncertainty. This trend, however, reversed again between 2015-16 so that the coefficients now hover around zero.

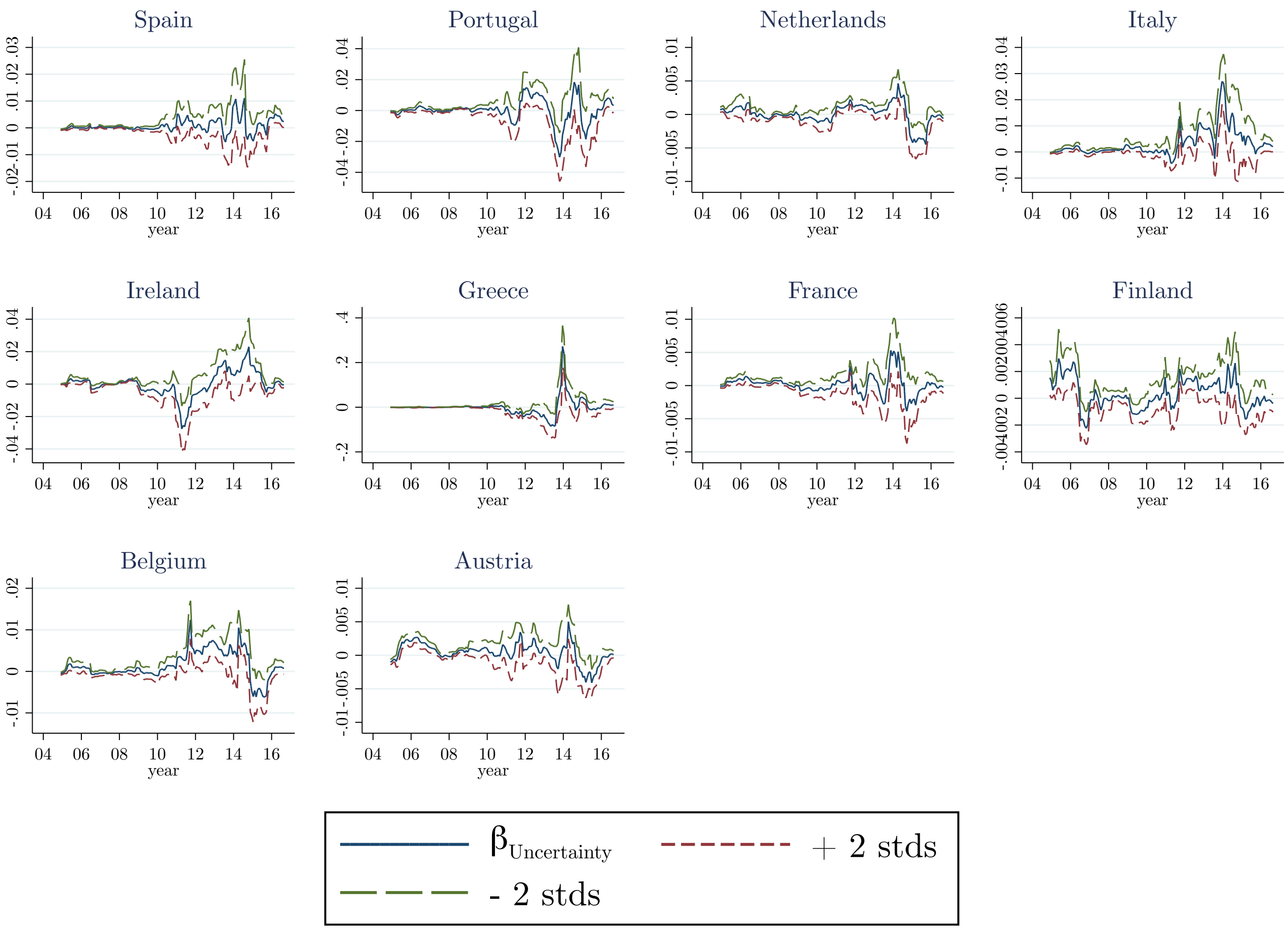

Figure 6: Rolling estimation of the uncertainty coefficient

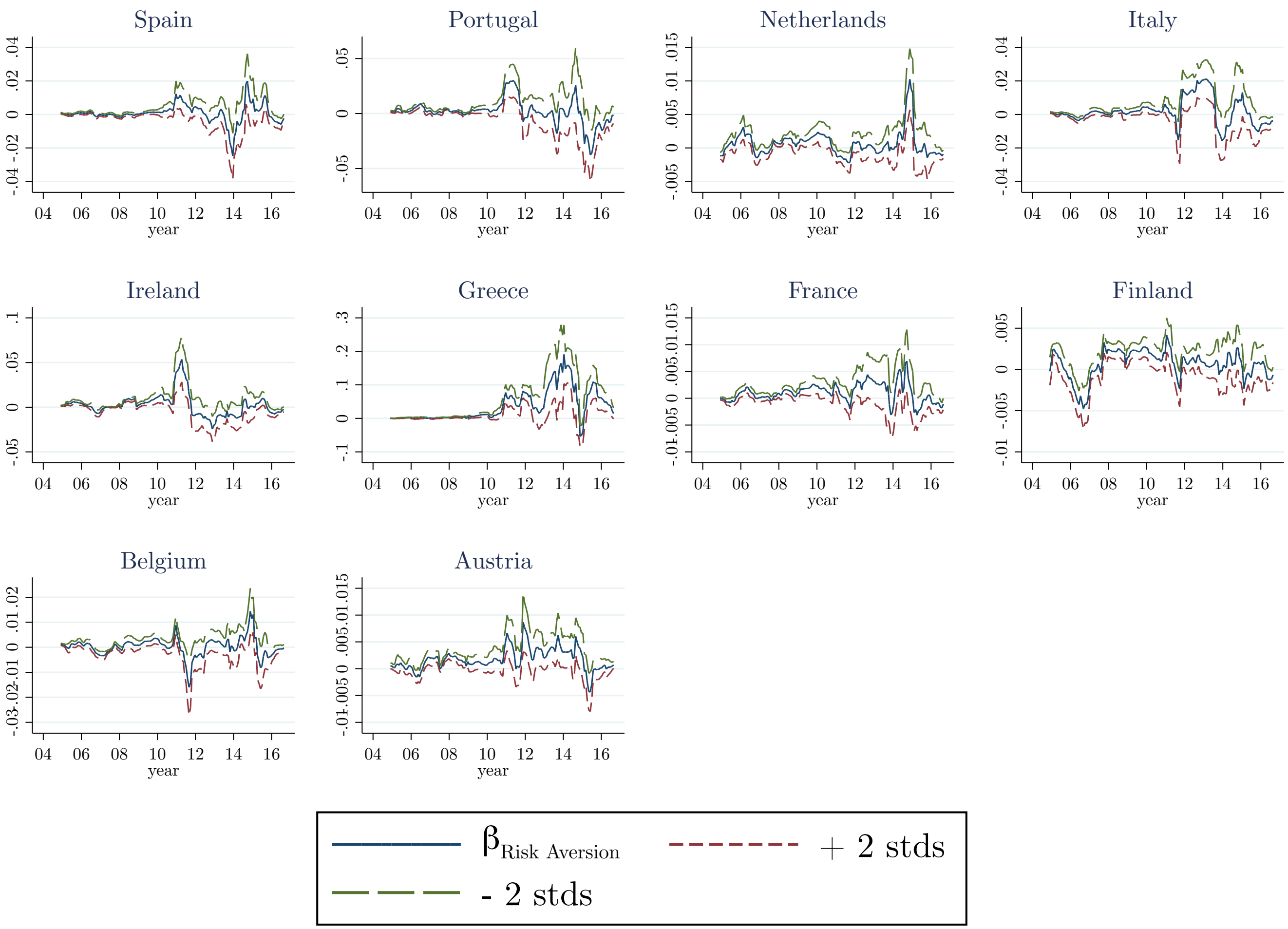

Figure 7: Rolling estimation of the risk aversion coefficient

The coefficient on risk aversion largely follows the same pattern (see Figure 7), with Finland and Austria again being the only countries with slightly significant positive coefficients during the pre-crisis period. During the crisis, the coefficient becomes positive for Portugal, Italy, Greece and Finland. Like the uncertainty coefficient, a general downward trend is observable after 2014 which then levels off again. Notably, the crisis states Spain, Portugal, Italy and Ireland seem to have a slightly negative statistically significant coefficient, whereas the Greek coefficient has been significantly positive since 2011 with two brief interruptions.

The start of the downward trend in both series corresponds to the general downward trend in spreads beginning in mid 2012. This suggests that the reduction in spreads is partially driven by the diminishing impact of the underlying uncertainty and risk aversion series.

4.5 Robustness Checks

This paper uses six different specifications of the model to test whether the findings are robust, three of which concern the credit risk variables and three of which concern the measure of the monetary policy stance: First, GDP growth is dropped from the model to avoid any possible multicollinearity between that measure, the current account balance, the budget surplus and the gross debt ratio, because GDP growth determines the denominator of all three variables. Secondly, the squared debt to GDP ratio is also included to control for non-linear relationships between levels of gross debt and spreads (as, for example, Bayoumi et al. (1995) find for U.S. states). Thirdly, the credit risk variables are replaced with the forecasts published by the European Commission (see Table 4). Since both the squared debt to GDP ratio and the EC's forecasts follow the same underlying process as the fiscal and economic fundamentals, it can be assumed that the forecasts equally follow a unit root process and are cointegrated. Fourthly, the nominal interest rate in first differences replaces the real interest rate as measure of the monetary policy stance. Fifthly, monetary policy shocks (see Figure 4) are used instead of the real interest rate as instrument for risk aversion. This should rule out any remaining doubts regarding the endogeneity of monetary policy and risk aversion since it only returns the response of risk aversion to structural disturbances by monetary policy. Lastly, risk aversion is instrumented by both the real interest rate in differences and monetary policy shocks. The results of the robustness checks (presented in Table A3 - A8) are as follows:

- The exclusion of GDP growth changes little regarding the country specific variables, only in case of Spain and the Netherlands some changes occur. The only two significant uncertainty coefficients in the baseline (Netherlands and Finland) become less significant. The risk aversion coefficient for Spain becomes insignificant while the coefficient for Italy becomes slightly significant with the expected positive sign. Instrumenting risk aversion, however, renders almost all those coefficients insignificant with the sole exception of Austria, where the coefficient on risk aversion is still slightly significant. The much less significant role of the instrumented risk aversion series compared to the baseline with GDP growth suggests that the part of risk aversion explained by the real interest rate alone is relatively small compared to the part explained jointly by the real interest rate and GDP growth.

- The small, yet in all but one country significant coefficient on the squared debt ratio suggests that higher public debt levels have a diminishing marginal effect. In two countries however, the Netherlands and Finland, the coefficient on the squared deficit ratio is positive, albeit very small. At the same time, the coefficient on the debt ratio for both countries has turned negative, indicating that higher debt levels were accompanied with lower spreads, although the marginal effect for higher public debt ratios is slightly declining. Both countries generally are perceived as save borrowers and therefore could have benefited from (relatively) declining spreads as their debt levels rose during the crisis. The inclusion of the squared debt to GDP ratio however comes at the cost of rendering some coefficients on the budget surplus (Spain, Portugal, Austria) and the current account surplus (Portugal, France) insignificant. In two cases the coefficients on the budget surplus become significant but with a positive sign (Greece, France). Uncertainty is significant in more cases than before (Italy, Spain). Risk aversion on the other hand becomes insignificant for Spain and Greece, while becoming slightly significant for Italy and Ireland. The estimated coefficients on uncertainty and risk aversion change little, so the changes can be mainly explained by differences in the coefficients' standard errors. Instrumenting risk aversion leads to few changes in the estimated coefficients compared to the baseline model, so that the changes in significance in the Netherlands and Italy (uncertainty), as well as Italy and France (risk aversion) can also be explained by changes in their standard errors.
- Estimating the model with forecasts instead of historical numbers

switches some signs on the budget surplus coefficients. It could be the case the in Spain, Portugal, Italy and France investors expected looser fiscal policy to have a stimulating effect on the economy. Curiously, however, is the now positive sign on the growth coefficient in case of Ireland and Belgium as well as the negative sign on debt to GDP ratio in Belgium. This could imply that markets see lower growth rates and higher debt levels in those two countries respectively as contributing to their long-term debt sustainability, perhaps due to an increased likelihood of European intervention. Another explanation would be that the longer run forecasts are systematically overly optimistic (e.g., Frankel, 2011): Since each of the three forecasts per year year is closer to the forecasted period (i.e. the upcoming year), any overly optimistic earlier forecast would likely be adjusted downwards, leading to a negative market reaction. The coefficients on uncertainty in the Netherlands and Finland become less significant compared to the original model specification, instead the coefficient in Italy and Austria become (slightly) significant. Whereas risk aversion is not a significant factor for France and Austria any more, it now is for Portugal. Both uncertainty and risk aversion have the same sign as in the original specification. Instrumenting risk aversion leads to an even higher number of countries for which uncertainty (seven instead of four) and risk aversion (seven instead of six) is a determining factor than in the baseline, even though only in case of Austria the coefficient is highly significant.

- Instrumenting risk aversion with the nominal interest rate renders a few country specific variables insignificant, especially for Italy and Belgium. Furthermore, it leads to more (slightly) significant coefficients on uncertainty (six instead of three) and risk aversion (seven instead of three) than in the baseline model. However, none of the coefficients is highly significant and only in case of the Netherlands, France and Finland (risk aversion) they even exceed the 5 percent significance threshold. Furthermore, all but two coefficients on GDP growth become insignificant which suggests that the nominal interest rate introduces a high degree of multicollinearity with the business cycle, which was partially avoided by the real interest rate.

- Using monetary policy shocks as measure of the monetary policy stance and as instrument for risk aversion does not suggest any impact of monetary policy on spreads via its impact on risk aversion. In case of

Spain, the Netherlands and Austria some country specific coefficients change in significance but there are no changes in the signs of the coefficients.

- Instrumenting risk aversion with both the real interest rate and monetary policy shocks barely changes the estimated coefficients from Table 9 - Section 2, but makes the coefficient on both uncertainty and risk aversion for Portugal slightly more significant. The uncertainty coefficient for Austria, however, is not significant at all any more. Overall, the results of this specification are hardly different from the original instrumental variable regression, further emphasizing that monetary policy shocks do not exert any influence.

Furthermore, it might be the case that the relationship between the spreads and the explanatory variables is not constant over time, as many earlier findings suggest. This paper focusses its attention on the stability of the uncertainty and risk aversion parameters estimated in the instrumental regression as they are the main variables of interest. This is done by estimating the model recursively, using the same window and step size as the rolling estimation in Section 4.4. The resulting series of coefficients is presented in Figure A1 & A2. Extreme spikes in the standard errors of the coefficients in case of Spain, Portugal, the Netherlands and Austria in late 2008 make it hard to draw any inferences about the stability of the parameters from visual inspection alone. Visual inspection of Italy, Ireland, Greece, France, Finland and Belgium however does not suggest any structural breaks in the relationship and neither does closer inspection of the data for the remaining countries.

5 Discussion

The different model specifications tested in Section 4.3 and Section 4.5 at some times yield very different outcomes. However, a few results generalise over various specifications: The coefficients on the debt to GDP ratio and the liquidity term are almost always significant with the expected sign, which corresponds to the existing literature. The sign on GDP growth is also constant across model specifications but its significant varies.

The signs of the coefficients on the budget deficit and the current account surplus are surprisingly sensitive to the model specification. A similar pattern, however, can also be observed in Giordano et al. (2013), where the coefficient on the current account balance changes signs depending on the estimation method applied to the data. Just as in the model specifications of this paper, they find that the gross debt ratio and GDP growth always have the expected sign. One source of misspecification here could come from the use of revised data in this paper. However, this should not bias the coefficient estimates if the initial data is not systemically over-optimistic or over-pessimistic.

Furthermore, across the various specifications, risk aversion is usually (more) significant for more countries than uncertainty is and, in absolute terms, has larger coefficients. This suggests that risk aversion plays a more important role than uncertainty and extends the results of Bekaert and Hoerova (2014) from stock returns to bond spreads: They find that risk aversion is a good predictor of stock returns while uncertainty is not. However, the result of S. R. Baker et al. (2015b), who find that their alternative series of economic policy uncertainty triggers bond market jumps, is not readily transferable to spreads. Unlike for the credit risk variables, the signs of the uncertainty and risk aversion coefficients are not sensitive to the specification of the model. The behaviour of both coefficients over time resembles the findings of Bernoth and Erdogan (2012) to some extend: During the pre-crisis period both coefficients are largely insignificant and then become significant from 2006 on, which is on average two to three years earlier than what the rolling estimation suggests. Unfortunately, their dataset only extends to early 2010 - comparing the two approaches to estimate time varying coefficients with an updated dataset would be a possible topic for another study.

The only occasion when the coefficients on risk aversion and uncertainty change their signs is when risk aversion is instrumented. In that case, they all have the expected positive sign, suggesting that monetary policy has a slight negative effect on spreads. The significance of that effect however is sensitive to the model specification and the measurement of the monetary policy stance - using monetary policy shocks instead of the real or nominal interest rate does not attest to any impact of monetary policy on spreads via risk aversion. Even though the start of downward trend in the estimated coefficients for uncertainty and risk aversion roughly corresponds to the timing of Mario Draghi's "whatever it takes" speech, those results suggest that monetary policy at best has a marginal effect on risk aversion and thereby on spreads. They of course do not rule out the possibility that monetary policy exerts some influence over spreads. However, the results of this paper suggest that this influence does not stem from the ECB's impact on risk aversion. For example, the correlation between short-term interest rates and spreads found by Manganelli and Wolswijk (2009) is likely to come from a different source than through their impact on risk aversion: The correlation between their measure of risk aversion (the corporate AAA - 10-year Treasury bond spread) and the interest rate is likely to be driven by an omitted factor, for example, the general business cycle, which is not fully captured by credit ratings, their credit risk variable.

Given the sensitivity of the budget deficit and the current account to the model specification, the estimated relationships should be taken with a pinch of salt. The consistency of the uncertainty and risk aversion estimates across models on the other hand strengthens the predictions of the model. Furthermore, the behaviour of the estimated coefficients over time in general fits the overall trend in spreads: Small and insignificant coefficients before the crisis, a big increase and volatile behaviour during the crisis, followed by a downward trend after 2012. But their varying significance, their relatively low absolute magnitude and the fact that both series are now below their pre-crisis levels while spreads remain elevated, suggest that risk aversion and uncertainty are not the main factors that can explain the surge in sovereign Euro area spreads. The decline of spreads and risk aversion might be better explained by, for example, the increased (and institutionalised) willingness of European states to bail out crisis countries and tackle problems in the financial sector.

From a policy perspective those results would imply that priority should be put on improving sovereign debt sustainability by reducing overall debt levels and increasing economic growth. This is in line with some of the previous literature, for example Santis (2012). Furthermore, political activism targeting the ECB for its supposed suppression of spreads by increasing risk appetite on capital markets seems misplaced.

There is one main downside to the model estimated in this paper: The estimated coefficients on risk aversion and uncertainty are very model sensitive and often surprisingly insignificant for the crisis countries, were one would expect them to have the strongest influence. On the one hand, this result is not completely without precedent in the literature: Barrios et al. (2009) find an impact of risk aversion on Belgian, French, Italian and Portuguese spreads, but not on Greek and Spanish spreads. However, it still is very counterintuitive that risk aversion should be significant for the Netherlands but not for Italy or Ireland, just to name two examples.

This might be explained by the fact that the uncertainty and risk aversion estimates derive from European stock markets and, therefore, might by an insufficient proxy of the compensation required to hold sovereign risk: Barrios et al. (2009) demonstrate that general risk aversion (measured by a PCA involving the AAA-BBB corporate bond spread, the VSTOXX and the exchange rate volatility in the Euro-Yen exchange rate) and sovereign risk (measured by a PCA on European sovereign spreads) moved closely together until 2009, after which risk aversion levelled off again while sovereign risk continued to increase. It could be the case that the risk aversion indicator derived in this paper does not fully capture the risk compensation required by markets to hold European sovereign risk. Instead, it could be argued that the risk compensation on stocks of European multinationals decoupled during the crisis from that on bonds of European sovereigns. Testing this theory would exceed the scope of this paper and thus would be a good starting point to extend the analysis of the topic raised here.

Finally, it should be noted that the results of this paper are probably not readily transferable to different circumstances: A supranational, industrialised currency union being hit with one of the worst crisis in history is probably a very unique setting. But even within the Eurozone it is hard to transfer those results to the countries that have not been part of this dataset: Those countries are relatively small, were only recently integrated into the European project and, therefore, might be structurally different. Therefore, the results of this paper should be rather seen as an attempt to explore the phenomena observed during the crisis and provide evidence to assess the actions of the ECB.

6 Conclusion

This paper provided an econometric model to disentangle the effects of uncertainty and risk aversion on Euro area sovereign bond yields vis-à-vis Germany, which were often confounded in previous studies. The model accounted for the non-stationary characteristics and the long-run cointegration relationship of its components in a dynamic OLS model whose applicability was ensured in a Monte Carlo simulation. Furthermore, the model incorporated the time varying nature of the relationships estimated by conducting a rolling estimation. The general results for country specific risks are in line with the existing literature. They also indicate that risk aversion was a more important factor than uncertainty and fit the narrative that the diminishing impact of risk aversion and uncertainty contributed to the decline in spreads. However, even though risk aversion and uncertainty have been on historical lows since 2012, spreads have remained elevated. This suggests that they are not the main source of divergence between Euro area sovereign yields. Furthermore, monetary policy affects spreads via risk aversion on the margin at best, as the instrumental variables estimates are highly sensitive to the model specification and the measurement of the monetary policy stance.

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A1 Horserace

Table A1: Model statistics

Coefficient	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
Constant	0.0005***	0.0004***	0.0003**	0.0012***	0.0003**	0.0004***	0.0010***
$VSTOXX^2$	0.4017***	0.4479***	0.5245***	-	0.5561***	0.4360***	-
$RVAR^{(22)}$	-0.0169	-0.0417***	0.0491	0.6803***	-	-	0.2656**
$RVAR^{(5)}$	0.5070*	0.8614**	-	-	-	-	0.8627**
$RVAR^{(1)}$	2.8115***	-	-	-	-	3.9356***	6.1340***

Note: Parameters are estimated using the full sample and twelve Newey-West (1987) lags.
***$p < 0.01$, **$p < 0.05$, *$p < 0.10$

Table A2: Model rankings

Model	RMSE	MAE	MAPE	Avg. Score
Model 1	**25.609**	**13.837**	**45.365**	**28.270**
Model 2	25.754	13.893	45.583	28.410
Model 3	26.168	14.084	46.003	28.752
Model 4	29.250	16.248	57.902	34.467
Model 5	26.200	14.128	46.232	28.853
Model 6	25.746	13.907	45.586	28.413
Model 7	26.200	14.128	46.232	28.853

Note: Statistics are based on the in-sample errors of the regression estimated using the full sample. For better comparability RMSE and MAE values have been multiplied by 10,000.

A2 Robustness Checks

Table A3: Robustness check I

	ES	PT	NL	IT	IR	GR	FR	FN	BE	OE
Part I: Baseline without instrumenting for risk aversion										
Intercept	-.754*	-4.062***	-.637***	-11.491***	-.597***	-14.620***	-1.292***	.112**	-2.101***	-1.857***
Current account	.012	-.082***	.016***	-.290***	-.047***	-.140***	-.062***	-.020***	-.0422***	.015***
Budget surplus	-.086***	-.023	.006**	.050***	-.080***	-.068***	.001	-.010***	-.015***	-.004
Gross debt ratio	.018***	.051***	.010***	.104***	.013***	.128***	.018***	-.001*	.021***	.024***
Liquidity	53.478**	21.338***	19.591***	62.868***	18.523***	5.830***	20.503***	15.335***	53.311***	13.861***
Uncertainty	.001	.000	.001*	.001	.002	-.011	-.000	.001**	.002	.001
Risk aversion	.000	.002	.002***	.005*	.003	.024***	.004***	.002***	.000	.003***
Part II: Instrumenting for risk aversion										
Constant	-1.251*	-4.032***	-.586***	-11.495***	-.285	-12.866***	-1.298***	.114**	-2.032***	-1.901***
Current account	-.032	-.094***	.023***	-.290***	-.053***	-.160***	-.066***	-.020***	-.038***	.018***
Budget surplus	-.075***	-.003	.006***	.050***	-.062**	.052	.001	-.010***	-.016***	-.008**
Gross debt ratio	.029***	.056***	.009***	.104***	.016***	.146***	.018***	-.001	.022***	.026***
Liquidity	56.541***	21.207***	15.294***	62.972***	20.821***	5.666***	21.156***	15.306***	50.857***	10.909***
Euro uncertainty	-.046	-.065	-.008	-.001	-.065	-.490	-.005	.000	-.022	-.014
$\widehat{Euroriskaversion}$	.061	.085	.013	.008	.091	.638	.010	.004	.031	.022*

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using the real interest rate in first difference.
$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

Table A4: Robustness check II

	ES	**PT**	**NL**	**IT**	**IR**	**GR**	**FR**	**FN**	**BE**	**OE**
					Part I: Baseline without instrumenting for risk aversion					
Intercept	-11.795***	-10.190***	.713***	-80.074***	-2.522***	-64.445***	-9.928***	.851***	-20.205	-4.057***
GDP growth	-.981***	-.375***	-.060***	-.469***	-.028***	-.578***	-.109***	-.014***	-.129***	.010
Current account	-.241***	.027	.019***	-.116***	-.091***	-.058***	-.004	-.017***	-.036***	.016***
Budget surplus	.002	-.017	-.001	.041***	-.046***	.075***	.016***	-.011***	-.012***	-.003
Gross debt ratio	.333***	.229***	-.036**	1.318***	.089***	.934***	.243***	-.034***	.386***	.082**
Gross debt ratio^2	-.002***	-.001***	.000**	-.005***	-.000***	-.003***	-.001***	.000***	-.002***	-.000
Liquidity	20.642*	18.561***	15.643***	41.724***	18.633***	5.347***	19.955***	17.519***	51.911***	14.139***
Uncertainty	.003**	.003	.001*	.007***	.001	-.005	.000	.001**	.002	.001
Risk aversion	-.003	-.002	.002**	-.005*	.004*	.012	.003***	.002***	-.000	.003***
					Part II: Instrumenting for risk aversion					
Constant	-13.496***	-10.337***	.417	-72.852***	-2.245***	-58.004***	-9.790***	.887***	-20.493***	-4.500**
GDP growth	-.766***	-.178	-.029	-.281**	-.020	-.479**	-.095***	-.008	-.038	.026
Current account	-.312	.010	.028***	-.120***	-.109***	-.087**	-.008	-.017***	-.037***	.020***
Budget surplus	.010	.019	.002	.044***	-.012	.153*	.016***	-.011***	-.014***	-.006
Gross debt ratio	.375***	.240***	-.025	1.191***	.104***	.865***	.240***	-.035***	.393***	.094**
Gross debt ratio^2	-.002***	-.001***	.000	-.005***	-.001***	-.003***	-.001***	.000***	-.002***	-.000
Liquidity	26.436**	18.890***	12.001***	48.551***	22.295***	5.278***	20.357***	17.475***	48.809***	11.939***
Euro uncertainty	-.072*	-.093	-.011	-.059*	-.118	-.387	-.003	-.005	-.022	-.012*
$\widehat{Euroriskaversion}$	.096*	.122	.018**	.082*	.157	.502	.008	.010	.032	.020**

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using the real interest rate in first difference.
***$p < 0.01$, **$p < 0.05$, *$p < 0.10$

Table A5: Robustness check III

	ES	**PT**	**NL**	**IT**	**IR**	**GR**	**FR**	**FN**	**BE**	**OE**
Part I: Baseline without instrumenting for risk aversion										
Intercept	-.108	-.080	-.181**	-8.050***	-1.251***	-9.220***	-.906***	.168***	.600***	-.180
GDP growth (forecast)	-.508***	-.766***	-.011	-.532***	.133***	-.107	-.166***	-.056***	.192***	-.068**
Primary budget surplus (forecast)	.101***	.201***	-.009	.389***	-.181***	-.385***	.027***	-.016***	-.181***	-.051***
Gross debt ratio (forecast)	.029***	.028***	.005***	.070***	.017***	.101***	.016***	-.000	-.006***	.006***
Liquidity	46.023**	15.223***	6.156**	81.178***	24.166***	4.978***	22.543***	18.374***	34.304***	8.583***
Uncertainty	.003	.003	.001*	.005*	.001	.008	.001	.001**	.002	.002***
Risk aversion	-.005*	-.009***	.002***	-.002	.004	.015*	.001	.002**	.003**	.001
Part II: Instrumenting for risk aversion										
Constant	-.437	-.237	-.179**	-9.914***	-1.568***	-11.332***	-1.380***	.111*	.601***	-.643**
GDP growth (forecast)	-.233	-.429*	-.010	-.167	.298***	.479	-.017	-.020	.120***	.061
Primary budget surplus (forecast)	.064	.201***	-.009	.405***	-.184***	-.596***	.040***	-.023***	-.182***	-.062***
Gross debt ratio (forecast)	.034***	.034***	.005***	.087***	.022***	.128***	.022***	.001	-.006***	.011***
Liquidity	55.647***	17.719***	5.929	92.238***	24.834***	4.995***	22.204***	18.067***	34.119***	6.973***
Euro uncertainty	-.095*	-.131*	.001	-.090*	-.091**	-.353	-.027**	-.009*	.001	-.018**
$\widehat{Euroriskaversion}$	.123*	.165*	.002	.121*	.126**	.482	.038**	.015**	.005	.028***

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using the real interest rate in first difference.
$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

Table A6: Robustness check IV

	ES	**PT**	**NL**	**IT**	**IR**	**GR**	**FR**	**FN**	**BE**	**OE**
Constant	-3.153***	-3.987***	-.510***	-14.298***	-.685***	-10.098***	-1.402***	.088	-1.401	-2.151***
GDP growth	-.315	-.055	.008	1.408	-.024***	-.841***	.104	.001	1.494	.074
Current account	-.207***	-.114***	.037***	-.150	-.042***	-.102***	-.097***	-.017***	-.059	.028**
Budget surplus	-.016	.030***	.007*	.075	-.084***	.056	.010	-.010***	-.054	-.017*
Gross debt ratio	.071***	.064***	.008***	.162***	.012***	.121***	.023***	.002	.037	.032***
Liquidity	63.163***	20.922***	7.836	102.750*	17.753***	5.452***	27.166***	14.768***	3.007	4.687
Euro uncertainty	-.201*	-.166*	-.024**	-.638	.026	-.340*	-.052**	-.021*	-.420	-.054
$\widehat{Euroriskaversion}$	.260*	.214*	.035**	.839	-.027	.440*	.071**	.030**	-1.401	.075*

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using the nominal interest rate. The baseline model without instrumenting for monetary policy is not reported because the estimates are the same as in Section 1 of Table 9.
$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

Table A7: Robustness check V

	ES	**PT**	**NL**	**IT**	**IR**	**GR**	**FR**	**FN**	**BE**	**OE**
Constant	-.920*	-3.977***	-.487***	-10.759***	-.274	-10.880***	-1.244***	.095**	-2.083***	-1.763***
GDP growth	-1.027***	-.118	-.072**	-.623***	-.018**	-.996***	-.141***	-.022***	-.107	-.0173
Current account	-.035	-.108***	.013	-.275***	-.050***	-.077***	-.047***	-.017***	-.0367***	.009
Budget surplus	-.034***	.018	.001	.028*	-.062***	-.031	-.001	-.009***	-.013***	.000
Gross debt ratio	.028***	.061***	.007***	.095***	.016***	.104***	.017***	-.001	.021***	.021***
Liquidity	38.078*	20.804***	17.524***	58.458***	20.704***	5.522***	20.627***	15.339***	54.228***	17.909***
Euro uncertainty	.037	-.135	.009	.057	-.062	.032	.013	.004	.001	.023
$\widehat{Euroriskaversion}$	-.052	.174	-.009	-.073	.087	-.037	-.014	-.001	.000	-.027

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using monetary policy shocks. The baseline model without instrumenting for monetary policy is not reported because the estimates are the same as in Section 1 of Table 9.

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

Table A8: Robustness check VI

	ES	**PT**	**NL**	**IT**	**IR**	**GR**	**FR**	**FN**	**BE**	**OE**
Constant	-2.288***	-3.968***	-.500***	-11.336***	-.269	-10.085***	-1.298***	.093*	-2.049***	-1.933***
GDP growth	-.591***	-.183	-.025	-.292*	-.018*	-.839***	-.057*	-.014**	-.026	.023
Current account	-.140**	-.102***	.027***	-.255***	-.050***	-.102***	-.064***	-.017***	-.038***	.017***
Budget surplus	-.023	.006	.005	.036**	-.062***	.057	.002	-.010***	-.015***	-.007**
Gross debt ratio	.054***	.059***	.008***	.106***	.016***	.122***	.019***	-.000	.022***	.026***
Liquidity	53.446***	20.682***	11.832***	65.677***	20.742***	5.451***	22.872***	15.134***	51.635***	12.100**
Euro uncertainty	-.109*	-.103*	-.010*	-.056	-.063	-.346	-.009	-.005	-.020	-.011
$\widehat{Euroriskaversion}$	.139*	.133*	.017**	.076	.088	.448	.015*	.010	.028	.018**

Coefficient estimates from a DOLS regression using eight Newey-West (1987) lags and six leads and lags of the non-stationary variables in first differences. Risk aversion in the Eurozone is instrumented using the real interest rate in first difference as well as the monetary policy shock series. The baseline model without instrumenting for monetary policy is not reported because the estimates are the same as in Section 1 of Table 9.

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

Table A9: Unconventional monetary policy announcements

Date	**Program type**	**Description**
22.08.07	Liquidity Support	Supplementary 3m LTRO
06.09.07	Liquidity Support	Supplementary 3m LTRO
08.11.07	Liquidity Support	Renewal of 3m LTRO
12.12.07	Liquidity Support	US dollar liquidity operations
07.02.08	Liquidity Support	Renewal of 3m LTRO
02.05.08	Liquidity Support	Liquidity support for funding markets
11.03.08	Liquidity Support	Supplementary 6m LTRO, renewal of 3m LTRO
28.03.08	Liquidity Support	Liquidity support for funding markets
31.07.08	Liquidity Support	Renewal of 3m LTRO
04.09.08	Liquidity Support	Renewal of supplementary LTROs
26.09.08	Liquidity Support	Liquidity support for Dollar funding markets
29.09.08	Liquidity Support	Special term refinancing operation
07.10.08	Liquidity Support	Allotment increase for the 6m LTRO
09.10.08	FRFA	Fixed-rate full allotment for MRO
13.10.08	FRFA	Fixed-rate full allotment for U.S. Dollar
15.10.08	FRFA	Fixed-rate full allotment for LTRO
07.05.09	Liquidity Support	Supplementary 1y LTRO
04.06.09	CBPP	CBPP 1
09.06.09	Liquidity Support	-
02.07.09	CBPP	CBPP 1
09.05.10	FRFA	Reactivation of the FRFA procedure
10.05.10	SMP	Detail of SMP announced
04.08.11	Liquidity Support	Supplementary 6m LTRO
08.08.11	SMP	SMP extended to Italy & Spain
06.10.11	CBPP	New CBPP announced
03.11.11	CBPP	CBPP 2
01.12.11	VLTRO	Draghi's speech at EU prliament
08.12.11	VLTRO	Announcement 3y LTRO
21.12.11	VLTRO	Round 1 of 3y LTRO
29.02.12	VLTRO	Round 2 of 3y LTRO
26.07.12	OMT	Whatever it takes speech
02.08.12	OMT	OMT mentioned at conference
06.09.12	OMT	OMT announced
04.07.13	FG	Forward guidance
09.01.14	FG	Forward guidance reiterated
06.03.14	FG	Forward guidance reiterated
05.06.14	VLTRO	ABSPP, announcement of 4y TLTRO

... table A9 continued

Date	**Program type**	**Description**
22.08.14	APP	Draghi's speech at Jackson Hole
04.09.14	APP/CBPP	Announcement ABSPP & CBPP 3
02.10.14	APP/CBPP	ABSPP and CBPP3
06.11.14	APP	Hint at PSPP
21.11.14	APP	Draghi's speech at the EBC
22.01.15	APP	Announcement of PSPP
05.03.15	APP	PSPP details
03.12.15	APP	Modification & expansion of the PSPP
10.03.16	APP	CSPP, announcement of new 4y TLTRO

Source: Kilponen et al. (2012), Szczerbowicz (2015) and Dewachter et al. (2016).

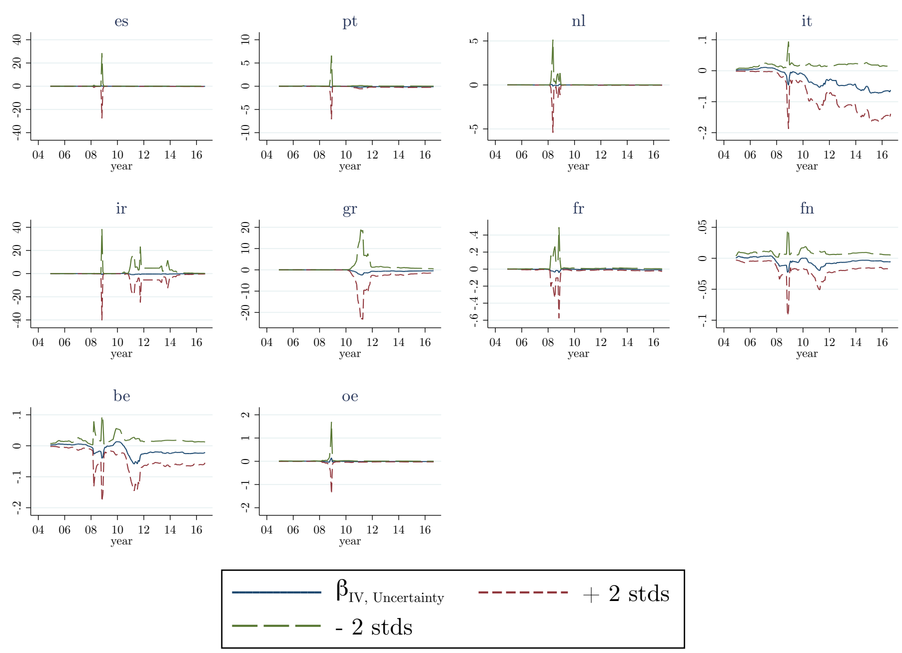

Figure A1: Recursive estimation of the uncertainty coefficient

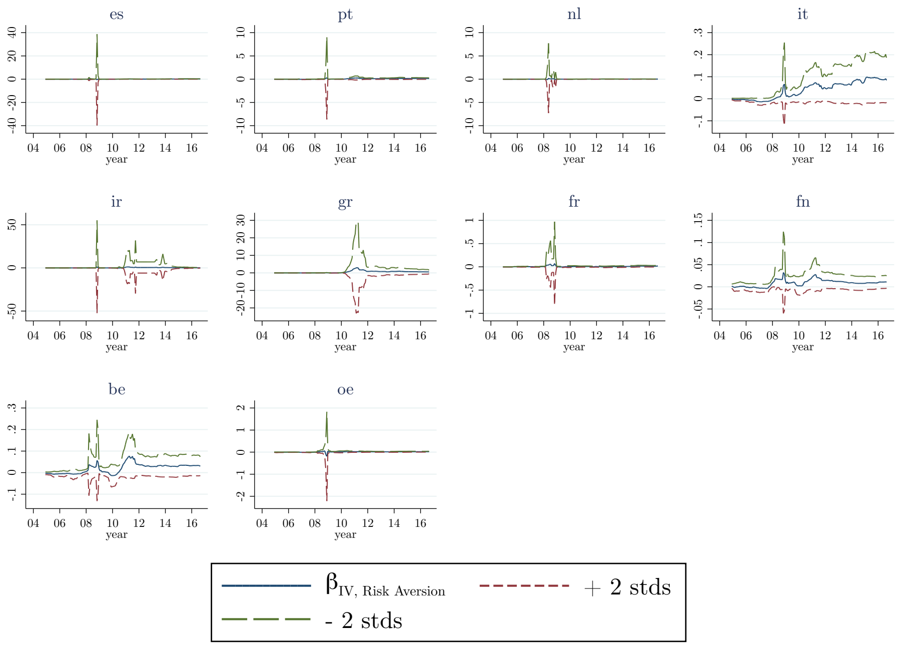

Figure A2: Recursive estimation of the risk aversion coefficient

A3 Monte Carlo Simulation

To ensure the applicability of the DOLS model used in this paper, a simple Monte Carlo simulation with four different data generating processes (DGP) is conducted.

- **Case 1:** Three non-stationary series that are unrelated to each other. The DGP is defined as follows:

$$\begin{aligned} Y_t &= Y_{t-1} + \epsilon_{1,t} \\ X_{1,t} &= X_{1,t-1} + \epsilon_{2,t} \\ X_{2,t} &= X_{2,t-1} + \epsilon_{3,t} \end{aligned} \tag{7}$$

- **Case 2:** Three non-stationary series that are related to each other. The DGP is defined as follows:

$$\begin{aligned} Y_t &= 0.5X_{1,t} + 0.5X_{2,t} + \epsilon_{1,t} \\ X_{1,t} &= X_{1,t-1} + \epsilon_{2,t} \\ X_{2,t} &= X_{2,t-1} + \epsilon_{3,t} \end{aligned} \tag{8}$$

- **Case 3:** Three non-stationary series of which two form a cointegration relationship. The DGP is defined as follows:

$$\begin{aligned} Y_t &= 0.5X_{1,t} + \epsilon_{1,t} \\ X_{1,t} &= X_{1,t-1} + \epsilon_{2,t} \\ X_{2,t} &= X_{2,t-1} + \epsilon_{3,t} \end{aligned} \tag{9}$$

- **Case 3:** Three series following a unit root of which two of them form a cointegration relationship. Additionally, a fourth series is created which is stationary and related to the dependent variable. The DGP is defined as follows:

$$\begin{aligned} Y_t &= 0.5X_{1,t} + \epsilon_{1,t} \\ X_{1,t} &= X_{1,t-1} + \epsilon_{2,t} \\ X_{2,t} &= X_{2,t-1} + \epsilon_{3,t} \\ W_t &= 0.6W_{t-1} + 0.6\Delta Y_t + \epsilon_{4,t} \end{aligned} \tag{10}$$

In each case, 800 replications for a sample size of $T = 1,000$ are simulated - Figure A3 presents an exemplary simulation. All $\epsilon_{i,t}$ are constructed as independent and identically distributed random variables with zero mean, variance of one and covariance of zero between each other. For each case the Engle-Granger test for cointegration is conducted first (note that the stationary variable W_t in case 4 is not part of the Engle-Granger test). Subsequently, a DOLS model with two leads and lags of each non-stationary variable in differences is estimated and the standard errors of the coefficients are corrected using three Newey-West (1987) lags. The first 100 observations are dropped from both regressions in order to allow for a sufficient burn in of the DGP. Table A10 displays the rejection rates of the Engle-Granger test and the rejection rates of the hypothesis that the coefficients on the non-stationary variables are equal to zero.

Table A10: Rejection rates

	EG Test[a]	$X_{1,t}$[b]	$X_{2,t}$[c]	$W_{1,t}$[d]
Case 1	95.25%	82.10%	83.25%	-
Case 2	0.00%	100.00%	100.00%	-
Case 3	0.00%	100.00%	5.00%	-
Case 4	0.00%	100.00%	8.75%	100.00%

[a] Percentage of simulations in which the null hypothesis of no cointegration is rejected; critical value of -3.77 (Table B.9, pg. 766 Hamilton, 1994).

[b] Percentage of simulations in which the null hypothesis of $\hat{\beta}_1 = 0$ is rejected at the 5% threshold (i.e. $|t| > 1.964$).

[c] Percentage of simulations in which the null hypothesis of $\hat{\beta}_2 = 0$ is rejected at the 5% threshold (i.e. $|t| > 1.964$).

[d] Percentage of simulations in which the null hypothesis of $\hat{\beta}_3 = 0$ is rejected at the 5% threshold (i.e. $|t| > 1.964$).

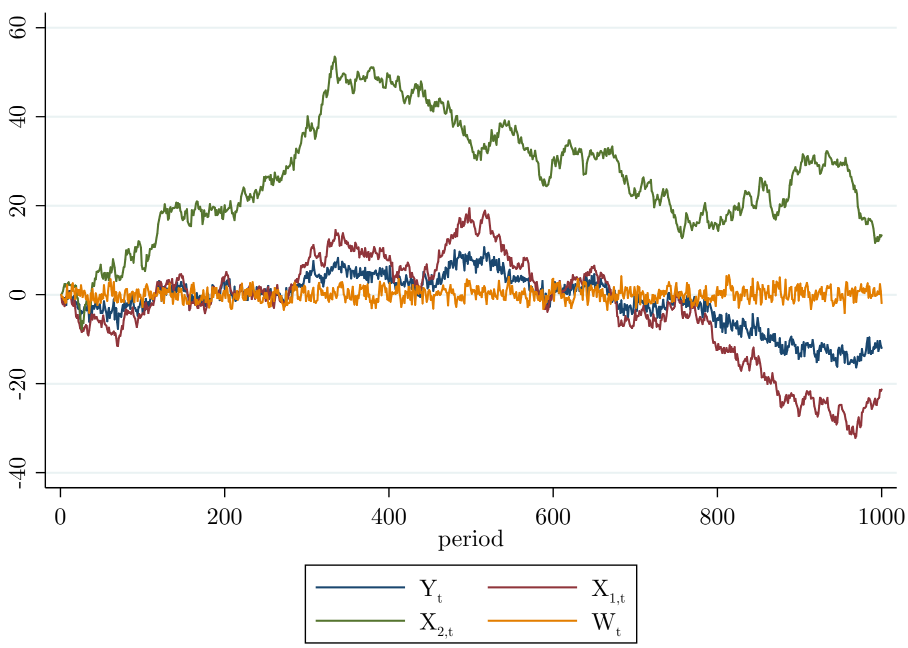

Figure A3: Simulation example