Risk and Uncertainty in Sovereign Debt Markets*

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Abstract

We assess the relationship between macroeconomic uncertainty and risk aversion and their role for the pricing of sovereign debt. A theoretical model of sovereign default is used to tell apart the effects of risk aversion and uncertainty for bond prices. We find that investors' risk aversion is positively affected by an increase in uncertainty, indicating toward uncertainty constituting a root cause for changes in risk attitudes. Building a structural VAR, we decompose credit default swaps (CDS) for Spain and Italy into three shocks: fundamental risk, risk aversion and an uncertainty. We find that shocks to macroeconomic uncertainty (1) have a significant and economically relevant impact on sovereign financing premia, (2) account for a share in sovereign CDS of up to 25 basis points at the onset of the European sovereign debt crisis, quantitatively comparable to the effect of increased risk aversion during this period, and (3) significantly increase international investors' risk aversion, in line with the predictions of the theoretical model.

Keywords: Uncertainty; Risk aversion; Sovereign debt; Markov-switching; Identification via heteroscedasticity

JEL Classifications: C32, D80, E43, G01, H63

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1 Introduction

Over the course of the European sovereign debt crisis of 2009-2012, affected governments in the euro area faced financial conditions in international capital markets that seemed to be inconsistent with public debt sustainability. In this context, a growing strand of empirical literature shows that public debt of distressed countries was priced at levels that cannot be explained by macroeconomic fundamentals alone.[1] Instead, other explanations, such as increased risk aversion, redenomination risk, contagion, or self-fulfilling default expectations were seen as drivers for the unexplained part in European bond returns.[2] This paper is part of the research effort attempting to understand the deterioration in financing conditions for European sovereigns.

We argue that macroeconomic uncertainty explains a relevant share in the sovereign financing premia, potentially having a similar impact on sovereign debt pricing as variations in investors' risk aversion and, thus, risk premia. Further, this paper shows theoretically and empirically that increases in uncertainty rises investors risk aversion, pointing toward uncertainty constituting a root cause for changes in risk attitudes. While variations in risk aversion have received much attention in explaining the observed worsening in the financing conditions of European sovereigns, this is — to our knowledge — the first paper to empirically investigate the relation between risk aversion and uncertainty as well as the impact of economic uncertainty on sovereign debt markets.

To explore the different nature of risk and uncertainty for the pricing of sovereign debt, we use a theoretical model. In this model of a small open economy that exists for two periods, a sovereign government is rolling over its accumulated stock of government debt. It cannot commit to repay its debt in the final period. It will do so depending on the realization of aggregate productivity. However, the law of motion of the aggregate productivity state is uncertain, similar to the case in Ilut and Schneider (2014). As a result, the payoffs from holding government debt turn out to be uncertain, too. International investors purchase the government

[1] See, among others, Aizenman et al. (2013), Grauwe and Ji (2012), or D'Agostino and Ehrmann (2014).

[2] Redenomination risk was at the center of the speech of ECB president Mario Draghi when he declared he would do *whatever it takes* to prevent that the risk of convertibility would drive up sovereign yields in the euro area ECB (2012). Kriwoluzky et al. (2014) and De Santis (2015) look systematically into redenomination risk. Hagen et al. (2011) and Bernoth and Erdogan (2012) document a sharp increase in risk aversion, Favero and Missale (2012) and Beirne and Fratzscher (2013) analyze contagion, while Bocola and Dovis (2015) address self-fulfilling default expectations. See below for a detailed discussion.

debt. Their preferences have two features. First, they exhibit constant relative risk aversion (CRRA) preferences with habit persistence. This yields variation in the intertemporal elasticity of substitution in consumption that affects their level of relative risk aversion, depending on the level of habit persistence. Second, investors exhibit preferences that make them sensitive toward the uncertainty surrounding the future aggregate productivity realization. Having multiple priors regarding the possible law of motion for productivity, maxmin preferences make them choose a worst case prior. We use the model to pin down analytically the three components, fundamental risk, risk aversion and uncertainty, in the arising asset pricing equation for risky government debt, which we subsequently identify in an empirical model.

Further, we show that a rise in uncertainty increases the degree of relative risk aversion of the investor. This is a result from the interaction between the worst case prior of investors and habit persistence. As investors expect lower levels of surplus consumption in the future, they tend to be more risk averse today. This mechanism is not described previously in the literature. In the empirical part, we provide evidence that an increase of uncertainty is typically followed by a rise in risk aversion in the data, as predicted by the model.

Based on the decomposition of sovereign financing premia of our theoretical model, we analyze the role of economic uncertainty for the pricing of sovereign debt empirically in a structural vector autoregressive (SVAR) model. Given the lack of identifying restrictions provided by economic theory, we exploit the statistical properties of the data in order to identify three shocks: A fundamental risk shock, a risk aversion shock, and an uncertainty shock. We deploy the structural model to empirically assess the relevance of macroeconomic uncertainty for the pricing of Spanish and Italian sovereign debt, decomposing their financing premia in contributions from the three shocks considered. We find that shocks to macroeconomic uncertainty have a significant impact on sovereign yields. They make up for close to 30 basis points in credit default swaps (CDS) at the onset of the European sovereign debt crisis, while their role diminishes as the sovereign debt crisis unfolds. Our model provides evidence for macroeconomic uncertainty to play a comparable role for the pricing of sovereign credit risk as time varying risk aversion. By jointly analyzing the effects of risk aversion and uncertainty, we achieve a comprehensive empirical identification of these two closely related concepts, which are often not clearly separated in the literature.

On the empirical side, this paper makes three main contributions. Firstly, we propose a novel high frequency measure of economic uncertainty. To this end, we apply the methodology put forward by Jurado et al. (2015) to a large set of Spanish and Italian equity returns. As a result, we obtain a weekly time series that reflects the underlying economic uncertainty faced by investors, entrepreneurs, and employees alike. Secondly, we propose an identification of fundamental risk, risk aversion, and uncertainty shocks within a Markov-switching structural vector autoregressive (MS-SVAR) model that makes use of the data properties following Rigobon (2003) and Lanne and Lütkepohl (2008). Such a statistical identification approach is particularly helpful as economic theory does not offer any structural restrictions that facilitate the disentanglement of risk aversion from uncertainty shocks. We label the statistically identified shocks by investigating their contribution to the forecast error variance of the endogenous variables in the SVAR model and confirm the labeling based on the heteroscedasticity pattern of the shocks. Thirdly, we quantify the share in sovereign yields over the most recent period of financial and fiscal distress in the euro area and find that uncertainty shocks account for a relevant share in the financing premium of the countries considered.

Relation with the literature. This paper is related to a growing literature on the determinants of sovereign yields. While Laubach (2009), Borgy et al. (2011) and Hilscher and Nosbusch (2010) find evidence for an important role of fiscal variables on government bond yields in US, European and emerging market data, respectively, fundamentals fall short in explaining the deterioration in sovereign financing conditions. There are four main explanations put forward in the existing literature for this overpricing of risk during the European sovereign debt crisis.

First, with the onset of the crisis, several papers investigate the time variation in investors' risk perception as an explanation of diverging European sovereign spreads. Hagen et al. (2011) find that markets turned more sensitive toward fiscal measures after the collapse of Lehman Brothers. Bernoth and Erdogan (2012) also find evidence for time-varying coefficients that determine the impact of fiscal variables for the pricing of sovereign debt and for investors' risk aversion in a semi-parametric approach. According to D'Agostino and Ehrmann (2014), time varying risk appetite of investors can explain some increase in European bond yields, but it still falls short of explaining the rise seen in French and Italian data over the crisis period. They conclude that observed yields are due to an overpricing of risk or possible concerns about redenomination of currencies. Further, Aizenman et al. (2013) and Haan et al. (2014) refer to overpricing in selected member countries of the euro area using cross-country panel data approaches.

Second, it is argued that the exit from the currency union would expose counterparties to the risk of redenomination into a new currency. Kriwoluzky et al. (2014) analyze this scenario against the Greek case. They use a structural macroeconomic model with exogenous probabilities of regime changes to decompose the sovereign yield of Greece. For identification, they exploit private versus public sector interest rates. They find that the exit risk can explain up to 10 percent in Greek bond yields. De Santis (2015) exploits the difference between credit default swap (CDS) contracts denominated in euro and US dollar to identify the redenomination risk contained in European sovereign yields. He finds that up to 50 percent, or 275 basis points, in Spanish yield spreads could be explained by changes in redenomination risk.

As a third explanation, Favero and Missale (2012) and Beirne and Fratzscher (2013) look into the role of cross-country contagion. Both papers find only a minor role for contagion over the course of the European sovereign debt crisis. In much more general terms, Longstaff et al. (2011) point toward a common factor in driving sovereign bond yields across the globe, along with risk aversion and liquidity factors.

Fourth, Bocola and Dovis (2015) quantify the share of non-fundamental self-fulfilling default expectations due to future inefficient roll-over crises in Italian data. They build a structural model of optimal sovereign default with sun-spot rollover crises and an endogenous maturity structure in order to decompose yields on Italian government bonds. They find that the non-fundamental share was important during 2011, but steadily declined throughout 2012, when yields were actually rising.

On the theoretical part, this paper is closest to models of sovereign default that look into the effects of risk aversion. Lizarazo (2013) shows how the introduction of risk aversion into investors' preferences makes the pricing of sovereign debt sensitive toward investors' stock of accumulated wealth. Borri and Verdelhan (2011) introduce time-varying risk aversion through habit in the utility function. Große Steffen (2015) analyses the effects of ambiguity aversion in a quantitative model of sovereign default. This paper is distinct in that it augments investor preferences by ambiguity aversion, risk aversion and habit persistence in consumption in order to achieve a clear separation of the concepts of risk and uncertainty in an encompassing framework.

The paper is structured as follows. The next section develops a small scale structural model in order to study the relation between macroeconomic uncertainty and the pricing of sovereign debt and motivate the following empirical analysis. Section 3 outlines the empirical setup for the analysis of economic uncertainty for sovereign financing premia. The data set is introduced in Section 4. This section also discusses in greater detail the construction of a high frequency index of macroeconomic uncertainty. Section 5 presents the empirical model used to analyze the relevance of uncertainty shocks for the pricing of sovereign debt, before results are presented in Section 6. Section 7 concludes.

2 A theoretical model

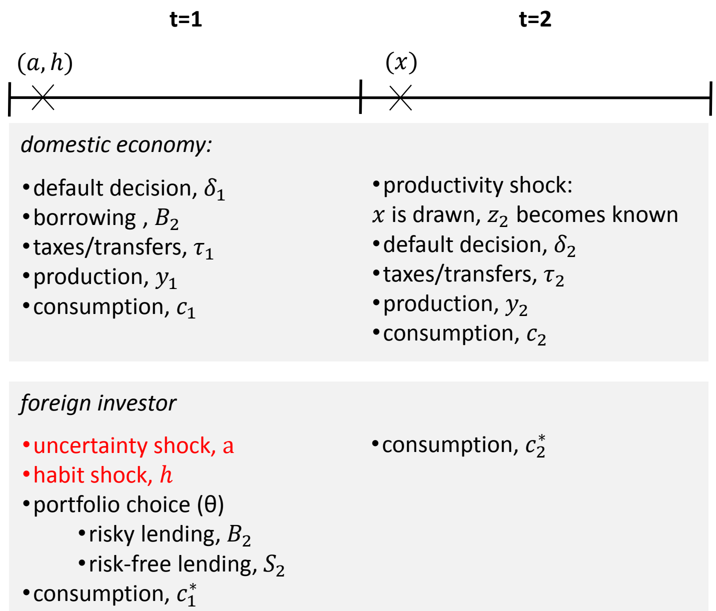
In this section, we develop a parsimonious model that jointly accommodates changes in uncertainty and risk aversion of investors holding defaultable government debt. The objective is to study the effects of theses two shocks in a unified framework for the pricing of risky government debt, motivating the decomposition of the sovereign yield spread in fundamental risk, time varying risk aversion, and uncertainty shocks. While the level of uncertainty is fully exogenous in this setting, one main result of the model is that risk aversion is endogenously affected by changes in macroeconomic uncertainty.

2.1 Environment

We start with a simplified model of optimal sovereign default in a production economy. The innovation lies in modeling the preference structure of international investors more exhaustively. Specifically, international investors are assumed to be simultaneously risk averse and ambiguity averse. Further, due to habit persistence in consumption of international investors, variation in the degree of relative risk aversion is introduced (Constantinides, 1990; Borri and Verdelhan, 2011).

We assume a small open economy over two periods, $t = 1, 2$. The economy is populated by three different agents: a representative household, a government, and a representative international investor.

Figure 1: Timing of events



Household. The household produces a final tradeable good y_t with constant labor input and subject to the aggregate level of productivity, thus $y_t = e^{z_t}F(l)$. The law of motion for aggregate productivity z_t is specified below. The household derives utility from consumption in each period, given by a quadratic utility function:

$$u(c_1, c_2) = \sum_{t=1}^{2} \left(c_t - \frac{\psi}{2}(c_t)^2 \right)$$

Final goods cannot be stored and therefore consumption is given by aggregate final good production net of government transfer payments or lump sum taxation, τ_t. Thus, the household is respecting a set of period t budget constraints of the form

$$c_1 = y_1 - \tau_1,$$

$$c_2 = y_2 - \tau_2.$$

Government. The government is a benevolent planner. Given limitations of private households to save or access international financial markets for consumption smoothing, the government provides an optimal tax- and transfer schedule that smooths private consumption, which are given by

$$\tau_1 = q_1 B_2 - B_1,$$
$$\tau_2 = -B_2.$$

In particular, the government may borrow from international investors in the form of one-period discount bonds, denoted by B_t. It enters the period $t = 1$ with the previously accumulated stock of debt B_1. Importantly, the government cannot commit to repay the debt when it becomes due in period t. Instead, it takes an optimal default decision. Default is a binary choice, denoted by $\delta_t \in (0, 1)$. When defaulting, the government suffers from an exogenous penalty, which comes in the form of a loss on aggregate output:

$$g(y) = \left\{ \begin{array}{lll} \hat{y} & \text{if} & z_t = z^h \\ y & \text{if} & z_t = z^l \end{array} \right\}$$

The penalty function $g(y)$ is pro-cyclical. The lower aggregate output, the lower is the output-loss conditional on defaulting. As a result and depending on the exact specification, default becomes more likely in times of below average production, thus during recessions, as shown by Arellano (2008). This is in line with the existing literature on endogenous output costs (Mendoza and Yue 2012; Engler and Große Steffen 2015), which is motivated by more frequent sovereign default episodes during recessions (Tomz and Wright, 2007).

Technology and uncertainty. The formalization of uncertainty closely follows Große Steffen (2015). In order to keep the model as parsimonious as possible, the productivity parameter z_t is assumed to feature two states, $z_t \in \{z^l, z^h\}$, with $z^h > 0$ and $z^l < 0$. We think of these two states as being sufficient in introducing risk and uncertainty about the future fundamental state of the economy. The values of z_t can be interpreted as recessions and booms, respectively. For simplicity, let aggregate productivity at $t = 1$ be deterministic and taking the lower value, $z_1 = z^l$. However, productivity at $t = 2$ is *uncertain* and can take either the low or the high value. This is decided from the realization of a random draw of the stochastic variable x, which is uniformly distributed on the interval $[x^*_{lb}, x^*_{ub}]$, i.e. $x \sim \mathcal{U}(x^*_{lb}, x^*_{ub})$. If the draw exceeds the threshold variable $\bar{x}$, the high productivity level realizes, thus $z_2 = z^h$. Agents know the threshold value $\bar{x}$ and that x is drawn from a uniform distribution. Thereby, future productivity is stochastic, hence *risky*. Additional uncertainty in the form of *ambiguity* enters into the law of motion of aggregate productivity. Specifically, we assume that the exact upper and lower bounds (x^*_{lb}, x^*_{ub}) of the distribution of x are unknown. In order to form expectations about the realization of future productivity, agents have multiple priors about these two parameters, which are specified next.

International investor. Government debt is purchased by a unit mass of identical international investors. The novelty here is to modify the preferences of these investors such that changes in relative risk aversion and the level of uncertainty regarding future productivity simultaneously matter for the portfolio decision of investors, and ultimately for the pricing of risky government debt in this framework. This is achieved by combining two concepts. First, the concept of habit persistence in the utility function allows for changes in the degree of relative risk aversion (Constantinides, 1990). Second, multiple-prior utility makes investor preferences sensitive toward changes in the level of uncertainty regarding future cash flows.[3] We refer to maxmin preferences as introduced by Gilboa and Schmeidler (1989) to formalize the multiple-prior model. International investors maximize their expected utility over the two periods, which yields the following optimization problem

$$\max_{\{c_1^*,\theta\}} \min_{\{supp^p(\mathcal{U}\in\mathcal{P})\}} \upsilon(c_1^*, h_1) + \beta E^p[\upsilon(c_2^*, h_2)]$$

$$\text{s.t.} \quad c_2^* = (W_1 - c_1^*)R^\omega + X$$

$$\begin{aligned} with \quad W_1 &\equiv X + S_0 + (1-\delta_1)B_1, \\ R^\omega &= (1-\theta)(1+r^f) + \theta E^p[1-\delta_2] \end{aligned}$$

where $\upsilon(c_t^*, h_t)$ denotes foreign investors' utility, β is the investor's discount factor, c_t^* is the amount the investor consumes, h_t denotes the level of habit persistence, X is exogenous income investors receive each period in the form of an endowment, S_t is a risk-free asset with a fixed and exogenously determined return r^f, and B_t denotes the risky government bond that trades at the endogenous price q_t. Maxmin preferences lead the investor to maximize expected utility, while simultaneously evaluating the minimum conditional on each prior $supp^p$ in the set $\mathcal{P}$.

[3]See Guidolin and Rinaldi (2010) for a survey on the effects of ambiguity aversion in asset pricing models.

The ambiguity averse investor then proceeds by acting as if there is just the worst case prior in the set $\mathcal{P}$.[4]

We follow Campbell and Cochrane (1999) with their specification of habit persistence in the constant relative risk aversion (CRRA) utility function, which yields time variation in risk aversion, using the following specification of investor preferences:[5]

$$\upsilon(c_t^*, h_t) = \frac{(c_t^* - h_t)^{1-\gamma} - 1}{1-\gamma}$$

Given the reduced time horizon in the analysis of two periods, we assume that the habit parameter h_t is fully exogenous and not determined by the history of aggregate consumption. Specifically, the habit parameter can take two values, $h_t \in \left\{h^l, h^h\right\}$. For simplicity, let $h_2 = h^l$. Here, we are mainly interested in the effects of a change in relative risk aversion in the initial period $t = 1$, which depends on the level of surplus consumption. Following Campbell and Cochrane (1999), we define the surplus consumption ratio of the investor as $\phi_t \equiv (c_t^* - h_t)/c_t^*$, such that the coefficient of relative risk aversion is given by

$$\eta_t = -\frac{c_t^* \upsilon_{cc}(c_t^*, h_t)}{\upsilon_c(c_{t,}^* h_t)} = \frac{\gamma}{\phi_t}. \tag{1}$$

If there is a higher realization of habit h_1, then this *ceteris paribus* leads to a reduction in surplus consumption and an increase in the local curvature of the utility function, which amounts to higher relative risk aversion of the international investor.

Next, it remains to describe what defines the set of priors considered by international investors regarding the ambiguous distribution of productivity shocks. We follow Große Steffen (2015) by assuming that there is an exogenous realization of uncertainty that pins down the set of prior beliefs about the true data generating process.[6] Specifically, investors are assumed to have *a priori* information about parameters of the distribution, denoted by $\tilde{x}^p_{lb}$ and $\tilde{x}^p_{ub}$. Then,

[4]The formation of a worst case prior is a result of the axiom of strict ambiguity aversion in Gilboa and Schmeidler (1989).

[5]There are several approaches to modeling habit persistence in the literature. In Abel (1990; 1999), utility is redefined as $\upsilon(c^*/h_t)$, which yields constant degree of risk aversion. Constant habit is modeled in Brunnermeier and Nagel (2008).

[6]This approach to modeling uncertainty is based on Ilut and Schneider (2014), adjusted to the simplified set-up described here.

an uncertainty realization from a known uniform distribution $a \sim \mathcal{U}[0, \bar{a}]$ pins down the set of prior beliefs about the true probabilistic model $\mathcal{U}(x^*_{lb}, x^*_{ub})$ as a symmetric interval around the *a priori* given parameters $\tilde{x}^p_{\cdot}$ according to

$$supp^p(\mathcal{U}) \in \mathcal{P} = \begin{cases} x^p_{lb} \in [\tilde{x}_{lb} - a, \tilde{x}_{lb} + a] \\ x^p_{ub} \in [\tilde{x}_{ub} - a, \tilde{x}_{ub} + a]\,. \end{cases}$$

Investors who exhibit the maxmin type of preferences then revert to the worst case prior, denoted by E^p_t, in order to sort out expectations. This means that they maximize the objective function conditional on every prior in the set, and then choosing the minimum from all possible outcomes. The first order conditions to the investors problem then amount to

$$\lambda_t = (c^*_t - h_t)^{-\gamma} - h_t \tag{2}$$

$$q_t = \beta \min_{\{supp^p(\mathcal{U} \in \mathcal{P})\}} E^p_t \left[\frac{\lambda_{t+1}}{\lambda_t} (1 - \delta_{t+1}) \right] \tag{3}$$

$$q^f = \beta \min_{\{supp^p(\mathcal{U} \in \mathcal{P})\}} E^p_t \left[\frac{\lambda_{t+1}}{\lambda_t} \right] \tag{4}$$

Using the definition of the covariance and dropping the minimization operator to simplify the notation, we can rewrite the asset pricing condition for risky government debt as

$$\begin{aligned} q_t &= \beta \frac{cov\left[\lambda_{t+1}, (1 - \delta_{t+1})\right] + E^p_t[\lambda_{t+1}] E^p_t[1 - \delta_{t+1}]}{\lambda_t} \\ &= \beta \frac{cov\left[\lambda_{t+1}, (1 - \delta_{t+1})\right]}{\lambda_t} + \beta \frac{E^p_t[\lambda_{t+1}] E^p_t[1 - \delta_{t+1}]}{\lambda_t} \end{aligned}$$

Now, substituting in the definition of the risk-free rate, i.e.

$$R^f = 1 / \left\{ \beta \left(E^p_t[\lambda_{t+1}] / \lambda_t \right) \right\}$$

the bond pricing condition can be rewritten as

$$q_t = \beta \frac{cov\left[\lambda_{t+1}, (1 - \delta_{t+1})\right]}{\lambda_t} + q^f \left(1 - E^p_t[\delta_{t+1}]\right) \tag{5}$$

The first term of this condition contains a risk-premium that is negative. Intuitively, if default is expected to happen and $(1-\delta_{t+1}) \rightarrow 0$, then this affects negatively the wealth of the international investor in the consecutive period, W_{t+1}, along with her consumption level. This, in turn, pushes up the marginal utility for future consumption, $\upsilon_{c(t+1)}$, which leads to the conclusion that $cov[\upsilon_{c(t+1)}, (1-\delta_{t+1})] < 0$. The lower asset price given from the covariance-term makes borrowing for the government more costly.[7]

The second term in the asset pricing condition for government bonds (5) captures the first order effect of uncertainty for the pricing decision of ambiguity averse international investors. Higher uncertainty leads to a more pessimistic worst case prior, hence to a higher default expectation $E_t^p[\delta_{t+1}]$.

However, condition (5) also shows a second order effect of uncertainty on the pricing of government debt. In particular, the covariance is affected by worst case prior beliefs about the repayment of government debt in the final settlement period. In order to show this, let's first define the covariance evaluated under the worst case prior as

$$cov^p(\lambda_{t+1}, (1-\delta_{t+1})) \equiv E_t^p\left[\lambda_{t+1}(1-\delta_{t+1})\right] - E_t^p\left[\lambda_{t+1}\right] E_t^p\left[1-\delta_{t+1}\right] \tag{6}$$

Here, we loosely follow Epstein and Schneider (2010), who show that in the classic mean-variance portfolio choice problem, ambiguity averse investors consider uncertainty in the co-variance matrix of assets for their decision. The evaluation of the covariance under the worst case prior is necessary as the underlying uncertainty is related to the fundamental state of the economy.

Next, we substitute back into equation (5) the definition of the covariance under uncertainty (6), which yields

$$q_t = \beta \frac{cov^p(\lambda_{t+1}, (1-\delta_{t+1}))}{\lambda_t} + q^f\left(1 - E_t^p[\delta_{t+1}]\right). \tag{7}$$

How does ambiguity aversion affect the covariance term of international investors? In order to

[7] As we know from the definition of the covariance,

$$cov[\lambda_{t+1}, (1-\delta_{t+1})] = E_t\left[\lambda_{t+1}(1-\delta_{t+1})\right] - E_t\left[\lambda_{t+1}\right] E_t\left[1-\delta_{t+1}\right].$$

For the covariance to exist, both its elements need to be stochastic variables. In fact, both arguments are dependent on the exogenous variables of the model, which are given by TFP (z_2), uncertainty (a) and eventually habit (h), see Figure 1.

discuss the differences between an ambiguity averse investor and the case without uncertainty, let us define the operator E_t as the expectation under the assumption that the level of realized uncertainty is zero, $a = 0$. Given the specific assumption in the model about the symmetry of the set of prior beliefs, this case coincides with a case where the investor uses a uniform weight for every prior in the set $\mathcal{P}$.[8] This case is typically understood as the standard assumption in the subjective expected utility paradigm. Therefore, we use it as a useful benchmark to illustrate the effects of uncertainty and ambiguity aversion in the model.

There are two counteracting effects that must be considered. On the one hand the expected cash flow under the worst case prior is lower than under subjective expected utility, as discussed before, such that $E_t^p[1-\delta_{t+1}] < E_t[1-\delta_{t+1}]$. On the other side is the expectation for marginal utility of consumption in the consecutive period higher under the worst case prior, such that $E_t^p[\lambda_{t+1}] > E_t[\lambda_{t+1}]$. This is the case since a higher probability of default of the government on its outstanding bonds will make consumption more valuable. The overall effect of ambiguity aversion on the financing premium is therefore *a priori* unclear.

Further, a change in uncertainty leads to higher risk aversion. We summarize this important result in a proposition.

Proposition 1. *Given that investors' attitudes toward risk and uncertainty can be jointly characterized by (i) risk aversion, (ii) habit persistence in consumption, and (iii) ambiguity aversion, an increase in uncertainty leads to higher coefficient of relative risk aversion.*

Proof. We need to show that the following condition holds:

$$-\frac{u''(E_t^{p,1}[W_2])}{u'(E_t^{p,1}[W_2])}E_t^{p,1}[W_2] < -\frac{u''(E_t^{p,2}[W_2])}{u'(E_t^{p,2}[W_2])}E_t^{p,2}[W_2],$$

where the level of uncertainty a_1 on the left hand side is lower than the level of uncertainty a_2 on the right hand side. Using the notation of equation (1), this condition can be rewritten as

$$E_t^{p,1}[\eta_{t+1}] = \frac{\gamma}{E_t^{p,1}[\phi_{t+1}]} < \frac{\gamma}{E_t^{p,2}[\phi_{t+1}]} = E_t^{p,2}[\eta_{t+1}].$$

Given that expected surplus consumption under the worst case prior is decreasing in uncertainty,

[8] Note that this would formally violate the assumption of ambiguity aversion, as such an investor would feature subjective probability measures to form expectations.

or $\partial E_t^p[\phi_{t+1}]/\partial a < 0$, this condition is fulfilled, such that the following relationship holds:

$$\frac{\partial E_t^p[\eta_{t+1}]}{\partial a} > 0 \tag{8}$$

□

The content of Proposition 1 is illustrated in Figure 2, where the wealth of the investor in period $t = 2$ under default and repayment are plotted ($W_{2,d}$, $W_{2,pay}$), along with the expected utility of the international investor. If uncertainty is positive ($a > 0$), then the investors' expected utility changes from $E_t[\upsilon(W_2)]$ to $E_t^p[\upsilon(W_2)]$. Due to the first-order effect of uncertainty through ambiguity aversion, expected utility is moving from point A to point B. It is through the presence of habit persistence in consumption that this change is accompanied by an increase in risk aversion of the investor.[9]

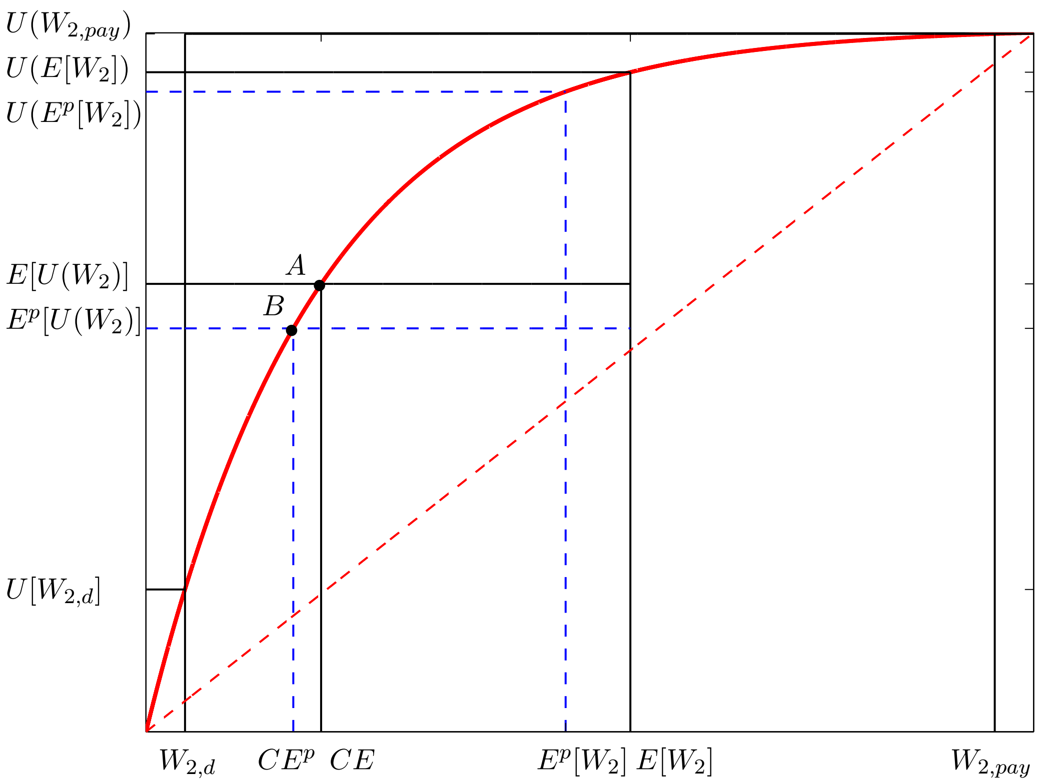


Figure 2: Effect of uncertainty on risk aversion

In the next section, we provide a numerical example in order to give an illustration of the

[9]A further detailed illustration of each of these investors' characteristics is provided in the Appendix B.

different effects that uncertainty and risk imply for the pricing of government debt.

2.2 Numerical illustration

We solve the model numerically in order to illustrate the two different effects of risk aversion and uncertainty. Equation (3) contains the results of various calibrations. First, we assume that investors are risk neutral and are not exposed to uncertainty ($\gamma = a = 0$). This would imply that the government can borrow within the risky borrowing region at a constant price given by $q_t = (1 - \pi^{z_l})/(1 + r^f)$. Note that the probability of default in this simple model collapses to the probability of a low productivity state (π^{z_l}). Since there is no penalty in the low productivity state, the government will default for sure in this state.

Next, introducing risk aversion makes investors charge a risk premium for holding risky government debt that increases in the investors' exposure. Habit persistence in consumption increases the risk premium, which is a well established result in the literature (Campbell and Cochrane, 1999).

The novel part in this model is to combine risk aversion with risk aversion. As illustrated in equation (3), introducing ambiguity aversion and habit persistence gives two effects. First, there is the well-known first-order effect of ambiguity on the pricing of risky asset which is characterized by the upward shift in the curve that characterizes the yields on government debt. Second, there is a second-order effect of ambiguity aversion that leads to a stronger increase in the yields on sovereign debt due to a higher level of risk aversion from the worst case prior from equation (7).

3 Empirical setup

The remainder of the paper, building upon the model outlined above, is concerned with the empirical assessment of the effect of economic uncertainty on the pricing of sovereign debt. Let us rewrite equation 7 such that the decomposition in fundamental risk, risk aversion and uncertainty on the government bond price becomes more evident. We use the property that for positive values of uncertainty we always have $E_t[\delta_{t+1}] < E_t^p[\delta_{t+1}]$, where E_t denotes the rational expectations operator using subjective expected utility and the assumption that all priors in

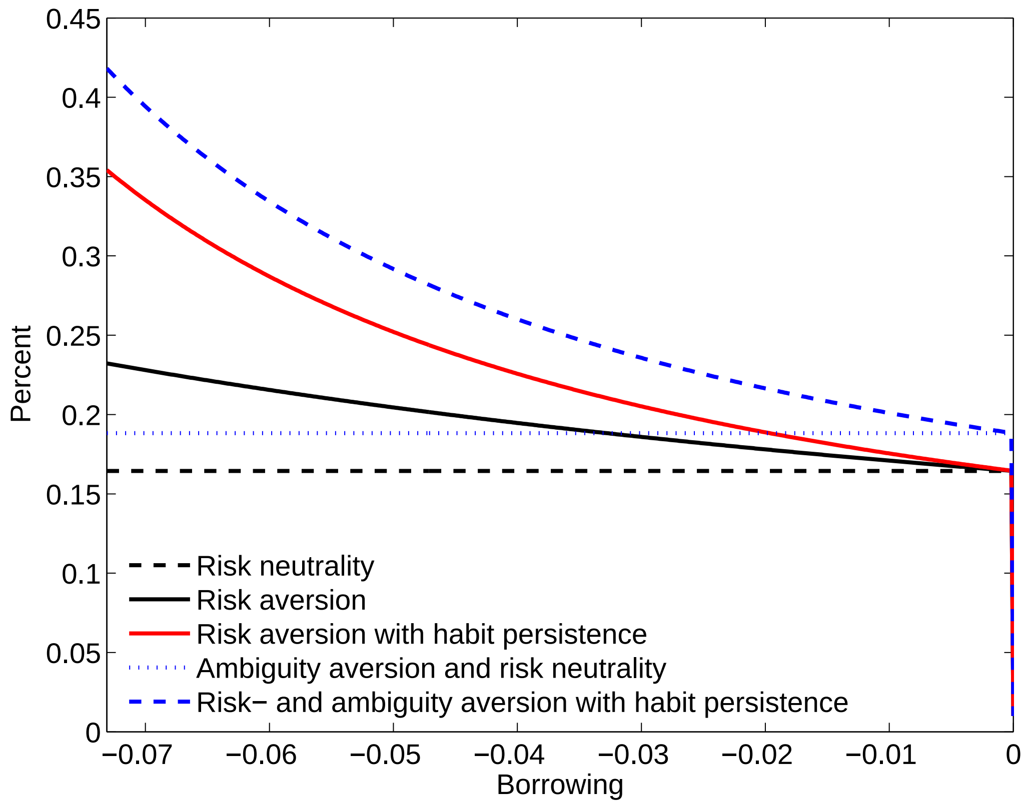


Figure 3: Bond returns from the model under risk aversion and ambiguity

the set $\mathcal{P}$ have identical weights, as discussed in the previous section. We then arrive at the following decomposition for the asset pricing equation:

$$q_t = \; q^f - \underbrace{\left(q^f E_t[\delta_{t+1}]\right)}_{\text{fundamental risk}} + \underbrace{\frac{\beta}{\lambda_t} cov(\cdot)}_{\text{risk premium}} + \underbrace{\frac{\beta}{\lambda_t}\left\{cov^p(\cdot) - cov(\cdot)\right\} + q^f\left(E_t[\delta_{t+1}] - E_t^p[\delta_{t+1}]\right)}_{\text{uncertainty premium}}, \quad (9)$$

where $cov(\cdot) = cov(\lambda_{t+1}, (1-\delta_{t+1}))$ and $cov^p(\cdot) = cov^p(\lambda_{t+1}, (1-\delta_{t+1}))$. Equation (9) is our point of departure for taking the model to the data. Note that the model implies financing premia to increase in fundamental risk, risk aversion, and uncertainty. Further, the bond price is inversely related to the yield. We make use of a trivariate Markov-switching in heteroscedasticity vector autoregressive model (MSH-VAR) containing a measure of the sovereign financing premium as well as measures of aggregate risk aversion and macroeconomic uncertainty and model the sovereign yield driven by the three (unobservable) terms in equation (9): a fundamental risk shock, a risk aversion shock and an uncertainty shock. Finally, we decompose the sovereign financing premium requested by market participants into contributions from these three shocks. In addition we use the model to evaluate the response of the measure of risk aversion to an uncertainty shock as an empirical assessment of the validity of Proposition 1.

The choice of model has the advantage of allowing to make use of the statistical properties of the data in order to identify the shocks of interest following the identification procedure pioneered by Rigobon (2003). In the context of VAR models, the properties of the data allow for the identification of orthogonal structural shocks within the model under certain conditions (Lanne and Lütkepohl, 2008). Making use of the statistical properties of the data for identification of orthogonal shocks is particularly helpful as economic theory does not provide a set of restrictions — neither exclusion restrictions on the short or long run impact matrix nor more agnostic sign, shape or magnitude restrictions — that would enable us to disentangle the three shocks of interest: a fundamental risk shock, a risk aversion shock, and a macroeconomic uncertainty shock.

While the statistical properties of the data help to uncover orthogonal shocks that are unique up to sign and column rotations from the model, they do not deliver any labeling of the shocks that would make them economically interpretable. Our strategy to label the set of shocks is twofold. Firstly, we draw on the information contained in the forecast error variance decomposition: The shocks explaining most of the variance in the risk aversion and the uncertainty measure are labeled risk aversion shock and uncertainty shock, respectively. The remaining shock, expected to dominate the variation in the financing premium, is labeled the fundamental risk shock. Secondly, in order to further back the economic interpretation of the shocks we follow the more narrative approach by Rigobon (2003) and exploit patterns in the series of the structural shocks uncovered from the MSH-SVAR model for a consistency check of the labeling. The two subsequent sections discuss details of the construction of the uncertainty index and the remainder of the data set as well as the specification of the MSH-SVAR in further detail before turning to the results.

4 Data

This section provides an overview of the data used in the subsequent analysis. It discusses in greater detail the construction of the uncertainty index and the measure of aggregate risk aversion, as well as describes the vector of exogenous variables — mainly related to unconventional monetary policy action — that are controlled for in the empirical analysis.

The vector of endogenous variables consists of a proxy for the sovereign financing premium, a measure of aggregate risk aversion, and an uncertainty proxy. The analysis covers Italy and Spain, two countries that exhibited a particularly strong deterioration in their sovereign financing conditions throughout the financial and sovereign debt crisis. Neither received any financial assistance from the European Financial Stabilization Mechanism (EFSM) or its successor, the European Stability Mechanism, that could potentially distort the estimation of the effect running from the uncertainty in the economy to the pricing of sovereign debt, discussed in Section 2. Limited by the availability of data, the sample spans from 2004 to 2015.[10] We use data of weekly frequency in order to average out noise in higher frequency, for example daily data, and to make the estimation of the model computationally feasible. Figure C.10 in the Appendix C plots the endogenous variables in the VAR model explained in detail below.

We proxy for the sovereign financing premium with credit default swaps (CDS), following Aizenman et al. (2013). CDS, usually traded over-the-counter, are derivatives that function similar to credit insurances. The seller of a CDS insures the buyer against the default of the creditor such that the price of CDS mirrors the financing premium of the underlying asset over a safe asset. The advantage of using CDS rather than yield spreads on sovereign bonds is that the CDS markets usually are more liquid and, hence, deliver more accurate measures of financing premia (Longstaff et al., 2011). Fontana and Scheicher (2010) find that price discovery for Spanish and Italian sovereign debt actually takes place in the sovereign CDS markets rather than in sovereign bond markets during the financial crisis. We obtain sovereign CDS data for Spain and Italy at five year maturity from Bloomberg.

[10]The limiting factor is the availability of sovereign credit defaults swaps data at weekly frequency. The sample spans from 01/12/2004 for Italy and 04/12/2004 for Spain to 04/20/2015 and includes 589 (Italy) and 576 (Spain) weekly observations.

4.1 A high frequency measure of macroeconomic uncertainty

In the construction of a measure of economic uncertainty, we face two main challenges. Firstly, the empirical model relies on the identification of different volatility regimes, which requires sufficient number of observations. As we aim at decomposing sovereign financing premia over the course of the recent period of fiscal stress, there is a natural limit to the number of available observations. A solution is to aim for a high frequency measure. Secondly, as the concept of uncertainty implemented in the theoretical model in Section 2 refers to the production outlook specific to the economy, the measure should be country specific and talk about the uncertainty of the production outlook.

In the construction of a high frequency, country specific, measure of economic uncertainty we follow the approach proposed by Jurado et al. (2015) in order to construct a proxy for country specific economic uncertainty.[11] They extract an uncertainty proxy at monthly frequency from a large dataset of macroeconomic variables by determining the common variation in the unforcastable component of those data. Their procedure involves three steps. In a first step forecast errors are obtained based on conditional mean forecasts from factor augmented autoregressive models. In a second step stochastic volatility in the forecast errors allows the extraction of uncertainty of variable y_j at time t for horizon h of a single series, which is defined by

$$\mathcal{U}^{y}_{j,t}(h) \equiv \sqrt{E\left\{\left[y_{jt+h} - E\left(y_{j,t+h}|I_t\right)\right]^2 |I_t\right\}},$$

where I_t denotes the information available at time t. A crucial assumption in their setup is that every series in the dataset features time varying volatility, which generates the time variance in uncertainty. In a third step the uncertainty estimates for the single variables in the dataset are aggregated to an economy-wide index of uncertainty. Appendix A provides a more detailed description of the construction of the uncertainty index in Jurado et al. (2015).

As we are aiming at a higher frequency, we depart from their approach in the use of the underlying dataset and apply their methodology to a large dataset of equity returns in order to construct a country specific high frequency measure of economic uncertainty.[12] The construc-

[11] Unfortunately, there is no stock market implied volatility available for Spain and Italy. Alternative measures, such as the disagreement or subjective uncertainty of professional forecasters would be a natural candidate for uncertainty regarding the production outlook, but are only available at lower, that is, monthly frequency.

[12] Jurado et al. (2015) argue that the construction of an uncertainty measure for macroeconomic data needs

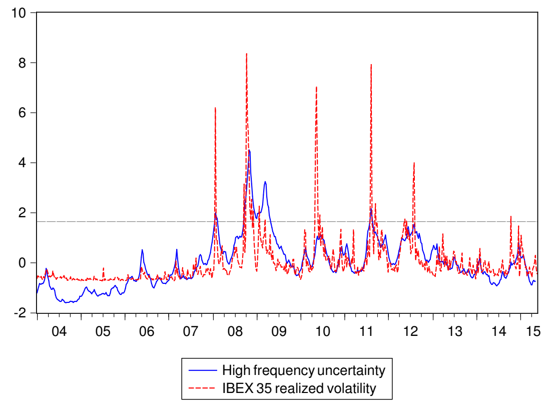
tion of the fundamental macroeconomic uncertainty measure builds upon a large set of weekly equity return data (total return index) from Thomson Reuters Datastream covering 1492 equity return index series for Spain and 1928 for Italy.[13] This dataset contains many degenerate series, as certain stocks are not traded continuously or traded only for a short period within the sample. We therefore remove all series that are either not traded or not present in the market for more than one third of the sample. In addition we control for equity splits and other outliers by removing observations associated with changes in the index above four standard deviations. After cleaning the dataset and filling the remaining gaps by means of a dynamic factor model along the lines of Schumacher and Breitung (2008), the dataset compiles 201 and 543 equity return series for Spain and Italy, respectively.

Figure 4 plots the high frequency uncertainty measures for Spain and Italy against stock market volatilities and the low frequency macroeconomic uncertainty measures based on monthly macroeconomic data. This type of comparison is particularly informative as we deploy similar data as the former and the same methodology as the latter. While the stock market volatility exhibits numerous large and significant jumps throughout the sample, but low persistence, the low frequency macroeconomic uncertainty peaks only once significantly at the onset of the financial crisis and is quite persistent. The measure proposed here, resembles those properties rather well, although it is constructed based on equity data: It features much stronger persistence than the stock market volatility and exhibits two significant peaks, the larger one at the onset of the financial crisis in late 2008 and one during the unfolding of the sovereign debt crisis in 2011.

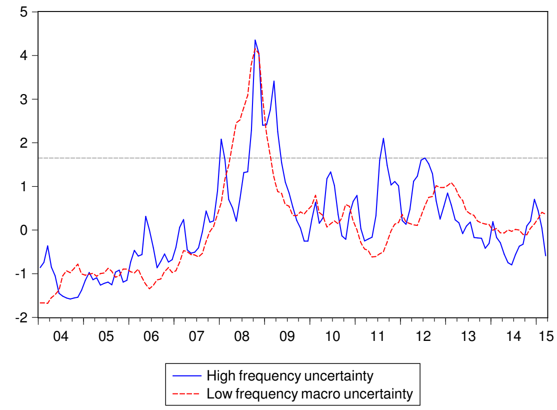
The high frequency measures of economic uncertainty for Spain and Italy for the period from 1990 to 2015 are plotted in Figure C.9 in the Appendix C. Both indexes clearly indicate their largest peaks in economic uncertainty around the most recent financial crisis in the years 2008 and 2009, but also around the introduction of the euro in 1999, while for Italy the rise

to take into account the forecastable component in macroeconomic data and deploy diffusion index forecasts for this purpose. Accounting for the forecastable component in equity returns may seem less urgent, given that equity returns are harder to forecast than macroeconomic time series. However, Ludvigson and Ng (2007) document that diffusion index models, as the one used in the construction of forecast errors here, provide forecasts for equity returns that are superior to using simple historical averages.

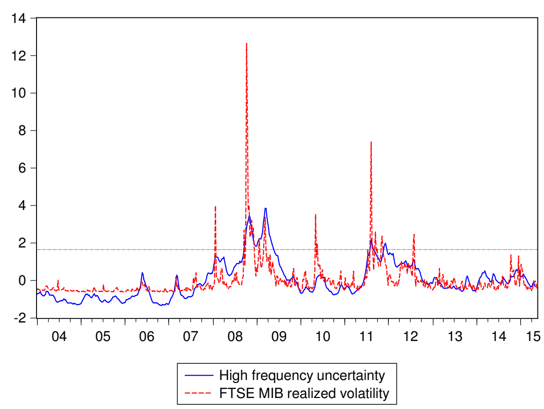
[13] The total return index is a better measure of the performance of a stock and its underlying company in that it includes not only the capital gain on the stock, but also returns related to dividend payments, the value of rights issues, special dividends, and stock dilutions.



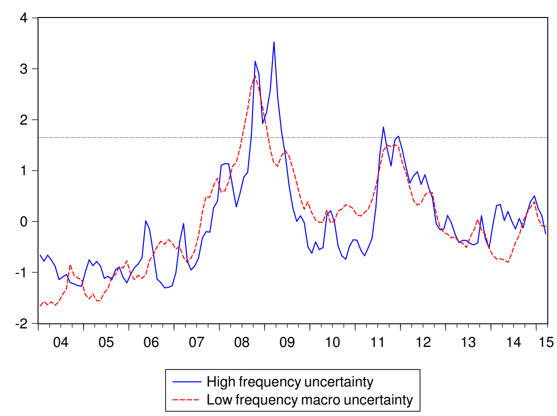
(a) Spain, comparison with stock market volatility



(b) Spain, comparison with low freq. macro uncertainty



(c) Italy, comparison with stock market volatility



(d) Italy, comparison with low freq. macro uncertainty

Figure 4: Comparison of our high frequency uncertainty measure with realized stock market volatility and low frequency macroeconomic uncertainty, data standardized and horizontal dashed line indicating 1.65 standard deviations.

in uncertainty surrounding the sovereign debt crisis 2011/2012 is more pronounced than for Spain.

Table 1 compares the constructed high frequency measure of macroeconomic uncertainty (aggregated to monthly frequency) to a number of alternative indicators of economic uncertainty at lower frequencies, that is, to (1) a measure of the degree of disagreement of professional forecasters,[14] (2) a low frequency macro data uncertainty index that we construct from a set

[14]Forecast disagreement captures the interdecile range of the distribution of point forecasts over GDP growth provided by a panel of professional forecasters, where the data is taken from Consensus Economics and Focus

of 51 monthly macro series for Spain and 84 monthly macro series for Italy; and (3) the news based policy uncertainty index provided by Baker et al. (2013) and weekly realized stock market volatility.[15] The aggregated high frequency measures correlate significantly with the set of alternative measures at lower frequency and the realized stock market volatility. The lowest correlation is found with the policy uncertainty measure. As it aims at capturing a somewhat different concept of uncertainty related to political decision processes, the lower correlation seems plausible and expected.

Table 1: Correlation of our uncertainty measure with alternative measures

	Spain	Italy	frequency	sample
Forecast Disagreement	0.46***	0.33***	monthly	2007M01 — 2014M08
Uncertainty based on monthly macro data	0.37***	0.46***	monthly	1990M07 — 2015M05
Policy uncertainty index (Baker et al., 2013)	0.27***	0.33***	monthly	1997/2001M1 — 2015M05
Realized stock market volatility	0.65***	0.66***	weekly	2000W1 — 2015W19
Uncertainty measure (Spain)	1	0.81***	monthly	1990M02 — 2015M05

Notes: The weekly uncertainty measure based on equity returns is aggregated to monthly frequency where necessary for comparison. *** Indicates significance at the 1 percent level.

The validity of the constructed measure as a high frequency indicator of macroeconomic uncertainty critically depends on the closeness of the link between equity markets and the real economy. We argue that under the assumption of efficient markets our equity based measure of economic uncertainty reflects the fundamental macroeconomic uncertainty in the economies, for investors, entrepreneurs, employees and other stakeholders of the considered companies alike.

Overall we take the strong and significant correlation among our high frequency uncertainty measure and alternative uncertainty measures together with the evidence from the graphical comparison in Figure 4 as reassuring in that we well capture economic uncertainty at weekly frequency.

4.2 A measure of risk aversion

In order to construct a measure of risk aversion of (international) investors, we borrow from the recent literature that computes the variance premium from options implied volatility indexes

Economics.

[15] Realized stock market volatility is taken from the Oxford-Man Institute's 'realized library' and aggregated to weekly frequency by averaging.

(Bollerslev et al., 2009, 2011; Bekaert et al., 2013; Bekaert and Hoerova, 2014). Option implied volatility indexes, for example, the CBOE volatility index (VIX), may be decomposed into one part capturing expected market volatility and a second part capturing risk aversion. We make use of such a decomposition in order to obtain a proxy for the risk aversion of international investors, as discussed in the model framework in Section 2. We base our measure of global risk aversion on the VIX, the options implied volatility index of the S&P500. We follow Bekaert et al. (2013) in constructing forecasts for the realized volatility based on a linear model incorporating the squared VIX and the past realized variance as predictors.[16] As in Bekaert et al. (2013), we winsorize the data prior to the estimation.[17] The difference between the squared VIX and the estimated conditional variance constitutes the proxy for risk aversion among international investors.

4.3 Exogenous controls

The vector of exogenous control variables in the MSH-VAR model contains the short term US nominal interest rate in order to control for global opportunity costs, bid-ask spreads controlling for time varying liquidity premia and a range of non-standard monetary policy measures, as we want to make sure our estimates are not affected by the extraordinary monetary policy action taken during the sample period. The monetary policy measures include dummy variables for the announcements of the Securities Markets Programme (SMP), the Long Term Refinancing Operations (LTROs) and the Outright Monetary Transactions (OMT) as well as variables capturing the volumes of their implementation[18]. The CDS Bid-Ask Spread is computed from bid and ask prices according to the formula $spread = \frac{p^{ask} - p^{bid}}{(p^{ask} - p^{bid})/2}$. Sources are Thomson Reuters Datastream, the ECB for data on unconventional monetary policy, and Bloomberg for the CDS prices, the bid-ask spreads are based upon.

[16]The data on realized variances for five minute windows is taken from the Oxford-Man Institute's 'realized library'.

[17]Winsorization eliminates outliers in the distribution by replacing values in the tails with those of the respective percentiles. The underlying data, that is the VIX and realized stock market volatility, are winsorized at the one percent level.

[18]We include volumes for the SMP and LTROs with 6-12 and 36 months maturity as additional exogenous variables.

5 The MSH-SVAR

The reduced form VAR model used for the empirical analysis is described by

$$y_t = \nu + A_1 y_{t-1} + A_2 y_{t-2} + \cdots + A_p y_{t-p} + \Gamma_0 x_t + \Gamma_1 x_{t-1} + \cdots + \Gamma_n x_{t-n} + \Xi d_t + u_t, \qquad (10)$$

where y_t is the vector of K endogenous variables, x_t contains the N exogenous variables and A_i's and Γ_j's are matrices that hold the respective coefficients with $i = 1, \ldots, p$ and $j = 1, \ldots, n$. ν is a vector of constant terms and d_t holds the L dummy variables with Ξ being its respective coefficient matrix. u_t represents the vector of reduced form error terms with $E[u_t] = 0$ and $E[u_t u_t'] = \Sigma_u(S_t)$. In addition we assume that the conditional distribution of u_t is normal, hence, $u_t|S_t \sim \mathrm{N}(0, \Sigma_u(S_t))$, that is, following Lanne et al. (2010) and Lütkepohl and Netšunajev (2014) the distribution of the reduced form error term is assumed to depend on a discrete Markov process S_t that can take on M values representing different regimes, $S_t = 1, \ldots, M$. While the model allows for Markov switching in the covariance of the residuals the parameters governing the first moments of the model are restricted to be constant over the sample.

The uncorrelated structural shocks, given by ε, map into the reduced form residuals as

$$u_t = B\varepsilon_t, \qquad (11)$$

via the matrix B of impact effects (see Lütkepohl, 2005, Chapter 9). Since the distribution of the residuals is governed by a Markov process, we have $\mathrm{var}(u_t|S_t) = \Sigma_u(S_t) = B\Lambda(S_t)B'$ with $E[\varepsilon_t] = 0$ and $E[\varepsilon_t \varepsilon_t'] = \Lambda(S_t)$. $\Lambda(S_t)$ is a diagonal matrix satisfying the orthogonality condition of the structural shocks. The variances, i.e. the diagonal elements, in the first state are normalized to unity, such that $\Lambda(1) = I_K$.

The assumption on the constancy of B may be challenged and newer literature is adopting more flexible models with state dependent impact matrices (Bacchiocchi and Fanelli, 2015; Podstawski and Velinov, 2016). However, the feature of interest in the current setup is the ability to make use of the statistical properties of the data in order to identify a set of structural shocks, rather than the analysis of a potential state dependency of the shock transmission.

Given the assumption on the constancy of the structural impact matrix B, the setup allows

— assuming that the diagonal elements of Λ are distinct (Lanne et al., 2010) — for the uncovering of a set of orthogonal structural shocks from the reduced form VAR model that are consistent with the statistical properties of the data. Given distinct diagonal elements in the covariance matrix of the structural shocks, the structural impact matrix B is unique up to sign and column permutations.

We exploit this feature of the MSH-VAR setup for the identification of the structural shocks of interest. Making use of the statistical properties of the data for identification of orthogonal shocks is particularly helpful for our endeavor as economic theory does not provide a set of restrictions — neither exclusion restrictions on the short and long run impact matrices nor more agnostic sign, shape and magnitude restrictions — that is suited to identify of the three shocks of interest: a fundamental risk shock, a risk aversion shock, and an economic uncertainty shock.

Under the assumption that the conditions for identification via heteroscedasticity are met by the models — an assumption that we address in the subsequent section — the MSH-SVAR model leaves us with three orthogonal shocks, ε_1, ε_2 and ε_3. These shocks need to be labeled and, hence, endowed with economic interpretation. We turn to this issue in Section 6.2.

6 Results

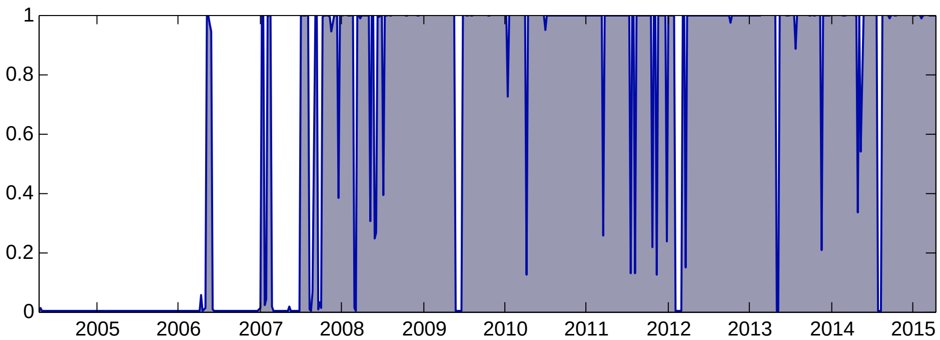
This section presents the results from the MSH-SVAR model for Spain and Italy. Based on information criteria for the linear model, we introduce two lags for the Italian model and three for the Spanish model. In order to keep the model parsimonious and since we mainly use the state switching property of the model for the purpose of identification, we resort to choosing two states for the Markov process.[19] Before turning to the analysis of the impulse responses and the historical decomposition, we report the state probabilities of the Markov process and a number of results related to the identification of the shocks.

[19]We attempt to investigate the data's preferences for higher order Markov-switching models by estimating the MSH-SVAR with three and four states. Maximizing the likelihood becomes increasingly more complicated with an increasing number of states. Given the sufficiency of two states to identify the three shocks in the model, we resort to two states for the Markov process.

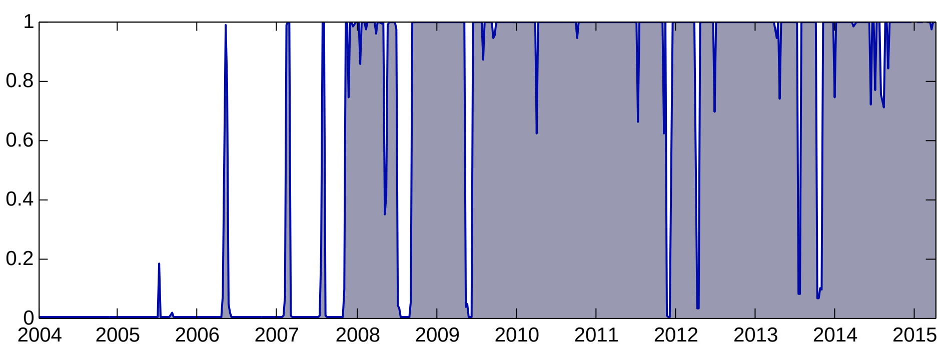
6.1 State probabilities

The smoothed state probabilities provide a first assessment of the suitability of the specified MSH-VAR model. Figure 5 reports the state probabilities for the high volatility state for the Spanish and Italian model.

Both models seem to clearly capture the crisis dynamics over the course of time, as they switch states around the emergence of the financial crisis in 2007/2008. The Markov-switching model identifies a low volatility state roughly before default of Lehman Brothers and a high volatility state afterwards, in which both economies remain for the remainder of the sample. The state probabilities clearly reflect the heteroscedasticity pattern in the data (see Figure C.10 in the Appendix C). All three endogenous variables exhibit low volatility in the period up to 2007 and higher volatility afterwards.



(a) Spain



(b) Italy

Figure 5: Smoothed state probabilities for the high volatility state

Figure C.11 in the Appendix C plots the reduced form residuals and the standardized reduced form residuals from both MSH-VAR models. The standardization takes into account the two volatility regimes, i.e. the diagonal elements of $\Sigma_u(1)$ and $\Sigma_u(2)$, and allows for an informal assessment of the fit of the model. The distinct heteroscedasticity pattern present in the non standardized residuals is tempered substantially by the standardization, indicating the models success in capturing the heteroscedasticity in the residuals. Although higher order MSH-VAR models may capture even more of the heteroscedasticity pattern, we conclude that the above state probabilities resemble the recent crisis dynamics, i.e. the heteroscedasticity pattern in the data, in a convincing manner. Next, we discuss issues related to the identification of the model and move from the reduced form to the structural model.

6.2 Identification

In a MSH-VAR model with two states the reduced form variances may be decomposed such that $\Sigma_{u1} = BB'$ and $\Sigma_{u2} = B\Lambda_2 B'$, where $\Lambda_2 = diag(\lambda_{21}, \ldots, \lambda_{2K})$, and all diagonal elements are positive. In order to make use of the statistical properties for the identification of the structural shocks we require the λ_{2i}s representing the variances of the structural shocks to be distinct (Herwartz and Lütkepohl, 2014; Lanne et al., 2010).

Table 2: Variances of structural shocks in the high volatility regime

	λ_{21}	λ_{22}	λ_{23}
ES	5246.125	23.471	5.543
	(10.845)	(4.807)	(0.750)
IT	1339.912	13.881	2.877
	(14.847)	(1.969)	(0.533)

Notes: Standard errors in parentheses.

Table 2 reports the estimated the relative variances of the structural shocks, λ_{2i}, for both models for the second state. Recall that the variance of the structural shocks is normalized to unity for the first state, i.e. $m = 1$. Clearly, the second state is the one exhibiting higher volatility and, indicating turbulent or crisis times. The point estimates exhibit some distance between the λ_{2i}s, while the error bands point toward a reasonable estimation precision and are small enough for the distributions to hardly overlap in the tails.[20] Overall, the point estimates and standard errors of the λ_{2i}s strongly indicate that identification is achieved based on the properties of the data.

6.3 Labeling the uncorrelated shocks

So far we have identified a set of three orthogonal structural shocks — ϵ_1, ϵ_2 and ϵ_3 — that we would like to label as fundamental risk, risk aversion and uncertainty shocks in order to make them economically interpretable. For the labeling we make use of the fact that proxies for two of the structural shocks we aim to identify are included in the vector of endogenous variables. We label the shock with the maximum contribution to the forecast error variance of the risk aversion proxy to be the risk aversion shock and the one with the maximum contribution to the forecast error variance of the uncertainty measure to be the uncertainty shock. The remaining structural shock is labeled fundamental risk shock. Table C.3 and C.4 report the forecast error variance decomposition (FEVD) for the first state at different horizons and allow a clear labeling based on the rational discussed above.

Figure C.12 in the Appendix C plots the three structural shocks and provides an opportunity for a further assessment of the labeling based on the forecast error variance decompositions. The dynamics of the shocks look quite similar among the models: The fundamental risk shock exhibits highest volatility during the sovereign debt crisis emerging around 2011/12 and has a very distinct pattern of heteroscedasticity. The risk aversion shock exhibits the strongest impulses during the unfolding of the financial crisis in 2008 — a pattern also found by Guiso et al. (2013) based on survey data of customers of Italian banks and in line with the general notion of countercyclical risk aversion (Cohn et al., 2015). In addition, this pattern seems to match the dynamics of alternative proxies for risk aversion such as the Baa-Aaa corporate bond spread provided by Moody's, which similarly jumps during this time period. Finally, the uncertainty shock is less clustered among the time dimension than the other two, but still exhibits phases of higher volatility during both, the financial crisis and the European sovereign

[20] There is no formal test available in order to assess the distinction of the λ_{2i}s. While previous research uses Likelihood-ratio and Wald tests to formally test for distinct λ_{2i}s, newer research indicates — according to personal communication with Helmut Ltkepohl — that the assumptions regarding the distributions of the test statistics may be inaccurate.

debt crisis broadly in line with the literature on economic uncertainty (Bloom, 2014). Overall the dynamics of the structural shocks strongly support the labeling based on the forecast error variance contributions.

In addition to the labeling of the structural shocks the FEVD provides first insights into the role of the three shocks for the sovereign financing premium. Clearly, the fundamental risk shock dominates the variations in CDS, but risk aversion and uncertainty shocks make up for a substantial share of the variation in CDS in both models, increasing in the forecast horizon. Also note that the risk aversion measure seems to contain a significant uncertainty component. We take this as first evidence of an impact of uncertainty shocks on investors' risk aversion in line with the predictions of Proposition 1. As opposed to that, the uncertainty measure seems rather well described by its own shock — and less affected by the fundamental risk and the risk aversion shock — judged by the forecast error variance decomposition.

Based on the identification and the labeling of the structural shocks, we turn to the impulse responses analysis and the historical decomposition of the financing premia in the subsequent sections.

6.4 Impulse responses

Impulse responses from both models are plotted in Figure 6. Confidence sets are based on a fixed design wild bootstrap in order to re-sample without corrupting the heteroscedasticity properties of the data.[21] The responses of the CDS spreads are in line with the theoretical prediction of the model presented in Section 2. All three shocks, the fundamental risk shock, the risk aversion shock, and the uncertainty shock impact positively on the CDS, increasing the borrowing cost for the sovereign. Among the three shocks, the fundamental risk shock has the largest short run impact on the sovereign financing cost, followed by the uncertainty and the risk aversion shock — in line with the findings from the forecast error variance decomposition in Tables C.3 and C.4 in the Appendix C. Note that the uncertainty shock impacts sovereign financing costs even stronger than the fundamental risk shock at longer horizons in the low volatility regime, although the fundamental risk shock becomes by far the strongest driver of

[21] For further details on the bootstrapping procedure see Podstawski and Velinov (2016).

sovereign CDS in the high volatility regime plotted in Figure C.13 in the Appendix C.[22]

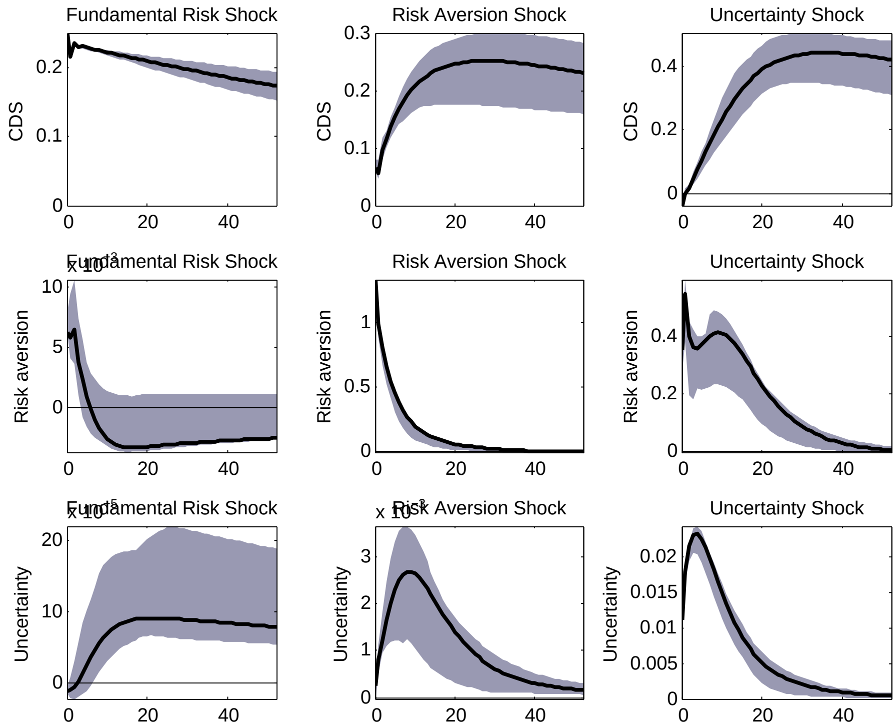
Quantitatively a one standard deviation uncertainty shock increases CDS by 0.5 to 1 basis points in the first regime of the MSH-SVAR model and between 1 and 1.5 basis points in the second regime. The effect of a uncertainty shock on CDS is somewhat comparable to that of a risk aversion shock, both with a somewhat larger impact in the high volatility regime. A fundamental risk shock of one standard deviation, however, has an impact effect of about 20 basis points in the high volatility regime, clearly dominating the risk aversion and uncertainty shocks. Overall the impulse responses for both economies are qualitatively and quantitatively very similar. They only feature slight differences in the persistence of the responses of the CDS to the three shocks.

The impulse response also allow for empirically assessing the prediction of Proposition 1; that is — given ambiguity averse investors with preferences further featuring varying relative risk aversion in wealth, and thus ambiguity — a positive response of investors' risk aversion to uncertainty shocks. Indeed, in both models we find a strong positive response of the measure of risk aversion to uncertainty shocks. This is in line with the large fraction of the forecast error variance of the risk aversion measure driven by uncertainty shocks. We take this as strong evidence in support of Proposition 1.

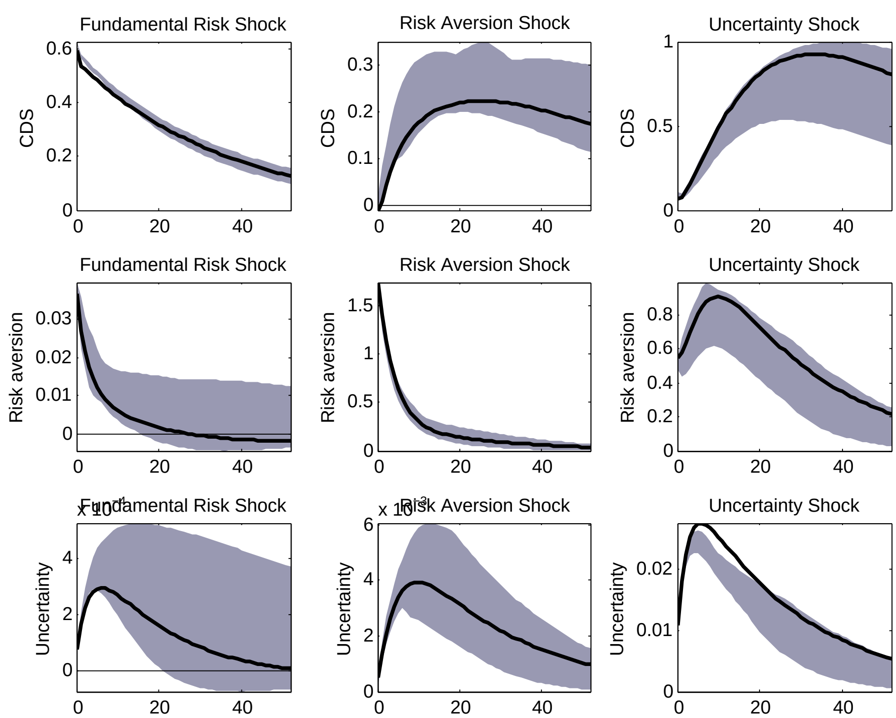
6.5 Historical decomposition

In order to assess the contributions of the three shocks to sovereign CDS, we conduct a historical decomposition based on the structural shocks. The series of structural shocks is constructed based upon equation (11) using the structural impact matrix B and the observable reduced form residuals. The Wald decomposition of the model described by equation (10) allows for expressing the endogenous variables at time t as a linear combination of initial values and structural shocks in the past according to

[22] Qualitatively, the impulse responses in the first regime and the second regime are identical by construction. The only difference stems from the higher variances of the structural shocks that maps into the set of impulses.



(a) Spain



(b) Italy

Figure 6: Impulse responses with 68% confidence intervals based on 1000 bootstrap replications, low volatility state

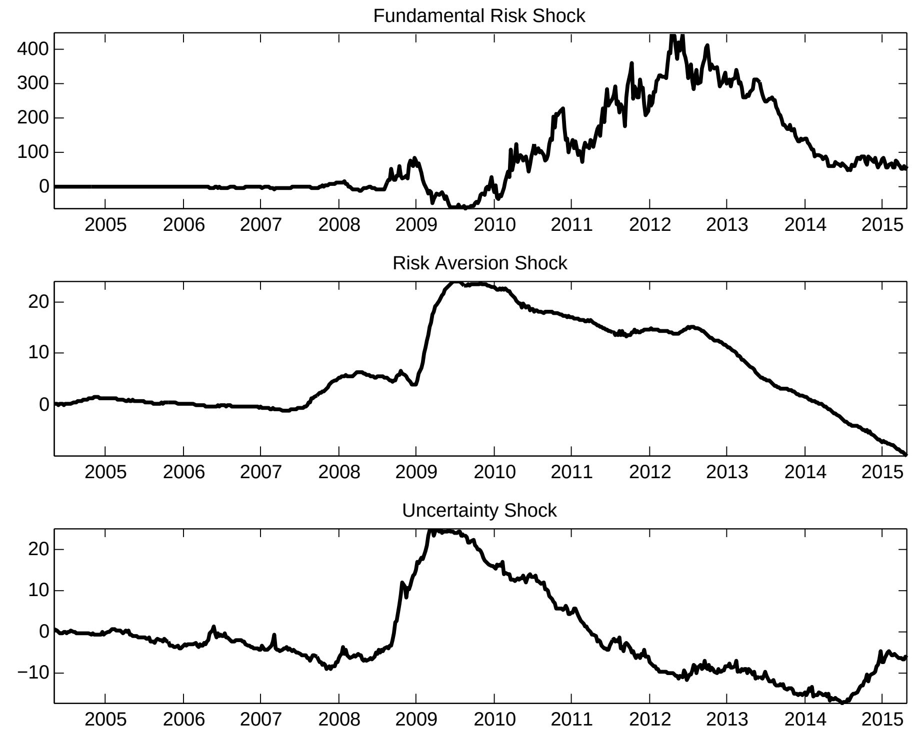
$$
\begin{aligned}
y_t &= \sum_{i=0}^{t-1} \phi_i u_{t-i} + A_1 y_0 + \ldots + A_p y_{-p+1} \\
&= \sum_{i=0}^{t-1} \phi_i B \varepsilon_{t-i} + A_1 y_0 + \ldots + A_p y_{-p+1},
\end{aligned}
$$

where ϕ_i is the matrix of impulse response coefficients $\phi_i = \sum_{j=1}^{i} \phi_{i-j} A_j$ with $i = 1, 2, \ldots$, $\phi_0 = I_K$ and $A_j = 0$ for $j > p$ (see Lütkepohl, 2005, Chapter 2) and we ignore the deterministic and exogenous terms in equation (10) that are not relevant for the impulse responses. The historical decomposition of the sovereign CDS variable into the three structural shocks is reported in Figure 7. It clearly supports the hypothesis of risk aversion and uncertainty shocks being relevant drivers of sovereign financing premia, notwithstanding the fact that fundamental risk shocks account for by far the largest share in CDS spreads, especially in later stages of the sovereign debt crisis.

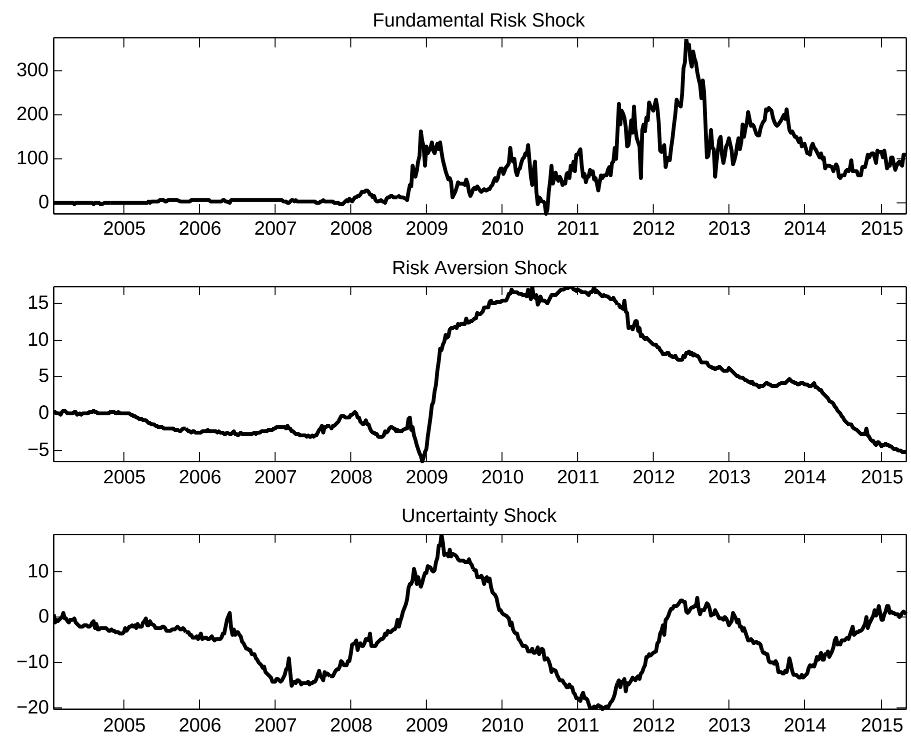
The decomposition indicates large positive contributions of both increased risk aversion and uncertainty to the financing premium of the Spanish and the Italian sovereign. Shocks to economic uncertainty make up for up to about 25 basis points in Spanish CDS spreads and 15 basis points in Italian CDS spreads, playing quantitatively a similar role to changes in investors risk aversion. The Spanish sovereign faced an increase in CDS of up to 25 basis points driven by an increase in risk aversion. Similarly, the Italian financing premium increased by up to 15 basis points due to rising risk aversion among international investors.

Our findings are in line with the literature that documents a crucial role of time variations in risk aversion for the pricing of sovereign debt.[23] Barrios et al. (2009) argue that general risk perception and its interaction with macroeconomic fundamentals are strong driving forces of yield spreads in the euro area. Sgherri and Zoli (2009) provide evidence that 15 to 30 basis points of the increase in Spanish and Italian yield spreads from late 2008 to early 2009 are attributable to increased risk aversion. Caceres et al. (2010) also find a positive contribution from global risk aversion to changes in Spanish yield spreads during 2009 but not for changes

[23] Note that we do not decompose CDS into a default-risk component and a risk premium, as do, for example, Longstaff et al. (2011), who find about one third of the CDS spread to be associated with the risk premium. Instead we assess the effects of changes in the risk aversion over time, against the backdrop of a steady-state risk aversion in the VAR model considered.



(a) Spain



(b) Italy

Figure 7: Historical decomposition of sovereign CDS of Italy and Spain

in Italian yield spreads.

While the sharp increase in financing premia for the sovereigns was fueled by risk aversion and uncertainty at the onset of the sovereign debt crisis, the picture is different for later stages of the crisis. Compared to the large contribution by fundamental risk shocks, there is only a minor impact from risk aversion and uncertainty shocks on sovereign financing premia, especially in 2012 when market participants became increasingly concerned about redenomination risk. This form of risk refers to the exit of member countries from the monetary union and the redenomination of their public and private liabilities. As we do not explicitly account for this specific form of exchange rate risk, it is part of the fundamental risk shock identified in our setup. In line with the large share of fundamental risk driving CDS upward during 2012, De Santis (2015) finds that close to half of the sovereign yield spreads were accounted for by redenomination risk in the first quarter of 2012.

Overall, we find clear evidence in support of a channel running from macroeconomic uncertainty to sovereign financing premia discussed in Section 2: Exogenous variations in uncertainty increase sovereign yields and make up for a non-negligible share in sovereign CDS. At the onset of the sovereign debt crisis in the euro area uncertainty shocks accounted for up to 15 basis points in Italian and up to 25 basis points in Spanish CDS spreads, an effect that is quantitatively comparable to the premium originating in rising risk aversion among international investors in the context of the global financial crisis.

7 Conclusion

In this paper we theoretically and empirically separate the effects of risk and uncertainty in the pricing of risky assets using the example of sovereign debt markets.

First, we build a simple model of optimal sovereign default that allows us to distinguish between the effects of risk aversion and uncertainty aversion for the pricing of risky government debt. In order to arrive at an analytical decomposition of the price for public debt, we assume that the investor is (i) risk averse, (ii) ambiguity averse, and (iii) has habit persistence in consumption. We show that risk aversion and ambiguity aversion lower bond prices and, hence, increase sovereign yields. Further, the model features an endogenous relationship between uncertainty and risk aversion: An increase in uncertainty affects the worst case prior of ambiguity averse agents, which feeds into higher levels of risk aversion.

Second, we take this theoretical decomposition of prices for government debt to the data. In order to jointly analyze the contributions from risk aversion and uncertainty for the financing premia faced during the European sovereign debt crisis empirically, we set up a structural VAR model and exploit the statistical properties of the data in order to identify three shocks: A fundamental risk shock, a risk aversion shock, and an uncertainty shock. Within this framework, we assess the relevance of risk aversion and macroeconomic uncertainty as drivers of the pricing of sovereign debt for Spain and Italy. We find that shocks to macroeconomic uncertainty (1) have a significant impact on sovereign yields; (2) make up for a non-negligible share in Spanish and Italian sovereign yield spreads of up to 25 basis points at the onset of the sovereign debt crisis that is quantitatively comparable to the effect of increased risk aversion during this period; and (3) significantly increases international investors' risk aversion, in line with the predictions of the theoretical model.

The results underline the relevance of macroeconomic uncertainty for the determination of asset prices and as a potential amplifier during times of economic crises. The theoretical and empirical connection between macroeconomic uncertainty and risk aversion documented in this paper might be an interesting avenue for future research.

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A Construction of the uncertainty measure

The following summary is based upon Jurado et al. (2015, section 3.1), which the reader is referred to for further details. Recall that uncertainty of variable y_j at time t for horizon h of a single series defined by

$$\mathcal{U}^y_{j,t}(h) \equiv \sqrt{E\left\{\left[y_{jt+h} - E\left(y_{j,t+h}|I_t\right)\right]^2|I_t\right\}}.$$

Assume that X_{it} contains the set of predictors used for forecasting and is representable by the following factor structure

$$X_{it} = \Lambda_i^{F\prime} F_t + e_{it}^X,$$

where F_t contains the latent factors, Λ_t the loadings and e_{it}^X the idiosyncratic errors. Forecasts for the series y_t are conducted using the following factor augmented autoregressive model

$$y_{jt+1} = \phi_j^y(L)y_{jt} + \gamma_j^F(L)\hat{F}_t + \upsilon^y_{jt+1},$$

where $\phi_j^y(L)$ and $\gamma_j^F(L)$ are polynomials in the lag operator L, $\hat{F}_t$ are estimates of F_t. The one-step-ahead prediction errors for each variable y_j and each factor F_t are allowed to feature time varying volatility, i.e. $\upsilon^y_{jt+1} = \sigma^y_{jt+1}\epsilon^y_{jt+1}$ and $\upsilon^F_{kt+1} = \sigma^F_{kt+1}\epsilon^F_{kt+1}$, an assumption that is crucial for the time variation in uncertainty. The forecasts $E\left[y_{jt+h}|I_t\right]$ are obtained from the factor augmented autoregressive (FAVAR) model, written in companion form as

$$\begin{pmatrix} \mathcal{F}_t \\ Y_{jt} \end{pmatrix} = \begin{pmatrix} \Phi^F & 0 \\ \Lambda_j' & \Phi_j^Y \end{pmatrix} \begin{pmatrix} \mathcal{F}_{t-1} \\ Y_{jt-1} \end{pmatrix} + \begin{pmatrix} \mathcal{V}_t^F \\ \mathcal{V}_{jt}^Y \end{pmatrix},$$

or written more compactly as

$$\mathcal{Y}_{jt} = \Phi_j^{\mathcal{Y}}\mathcal{Y}_{jt-1} + \mathcal{V}_{jt}^{\mathcal{Y}},$$

where $\mathcal{F}_t$ collects the factors and additional predictors used for forecasting, Y_{jt} represents the set of variables that are to be forecasted, Λ_j' and Φ_j^y are collections of the coefficients in the matrix polynomial lag operators from the single factor augmented forecasting model. The optimal forecast is given by the conditional mean

$$E_t\mathcal{Y}_{jt+h} = (\Phi_j^{\mathcal{Y}})^h\mathcal{Y}_{jt}$$

and the forecast error variance at horizon $h = 1$ is given by

$$\Omega^{\mathcal{Y}}_{jt} = E_t(\mathcal{V}^{\mathcal{Y}}_{jt+1}\mathcal{V}^{\mathcal{Y}\prime}_{jt+1}).$$

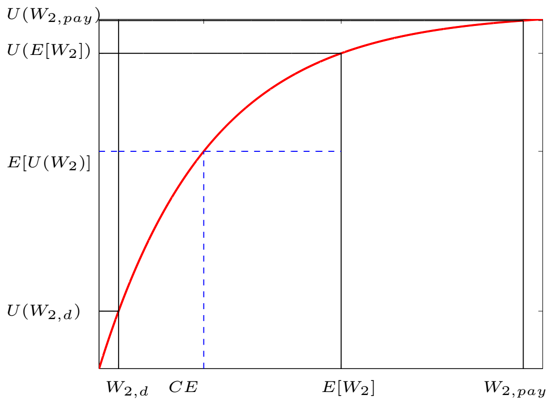
In order to obtain the expected forecast uncertainty of a single variable y_{jt+h} based on the information set at time t, we select a single entry of the forecast error variance matrix $\Omega^{\mathcal{Y}}_{jt}$ using 1_j as a selection vector

$$\mathcal{U}^y_{j,t}(1) = \sqrt{1_j'\Omega^{\mathcal{Y}}_{jt}1_j}$$

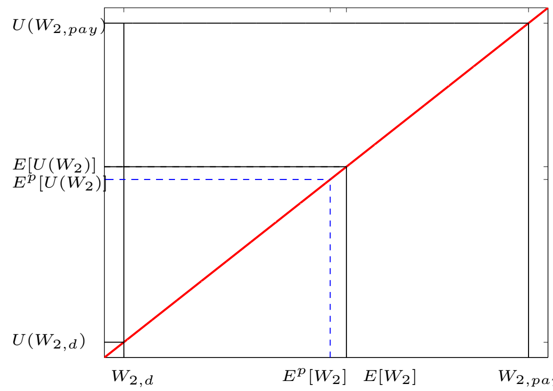
Aggregate uncertainty is finally computed as the average of the individual forecast uncertainties

$$\mathcal{U}^{y}_{j,t}(1) = \frac{1}{N}\sum_{j=1}^{N_y}\mathcal{U}^{y}_{j,t}(1)$$

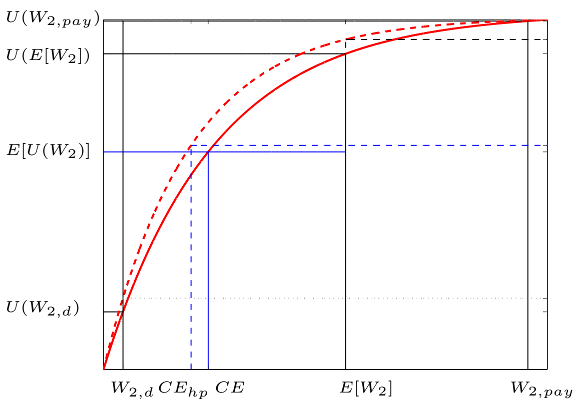
B Risk and uncertainty attitudes



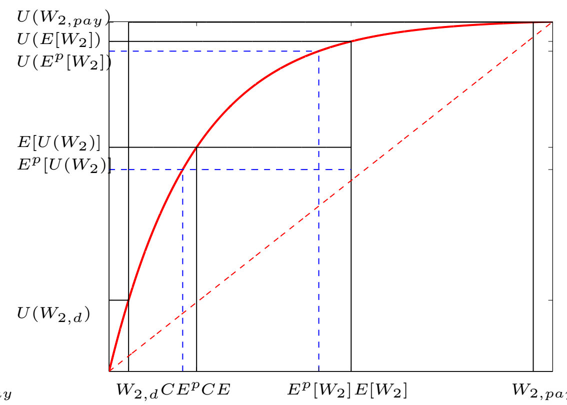
(a) Risk aversion



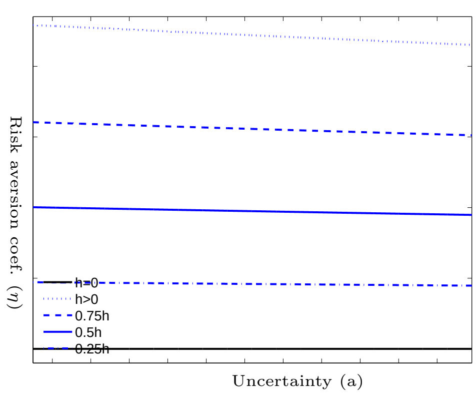
(b) Ambiguity aversion and risk neutrality



(c) Habit persistence



(d) Risk, ambiguity, and habit persistence



(e) Coefficient of relative risk aversion (η)

Figure B.8: Risk attitudes, utility on y-axis and wealth on x-axis wherever not noted otherwise.

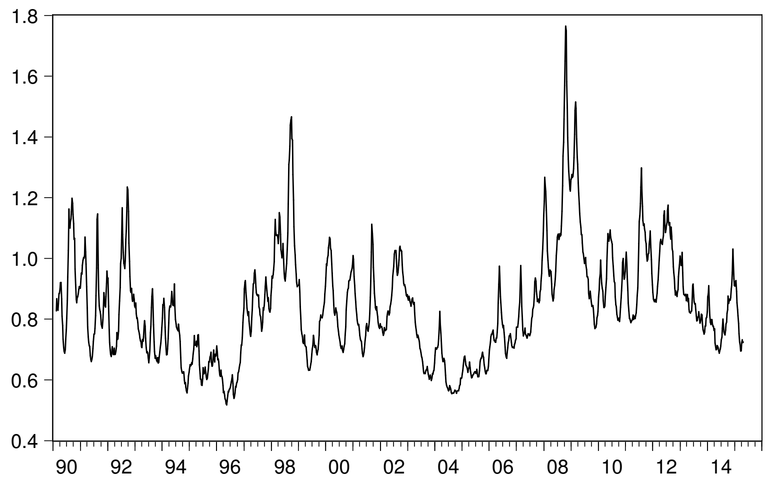
C Tables and figures

Table C.3: Forecast error variance decomposition from the MSH-SVAR for Spain (low volatility state)

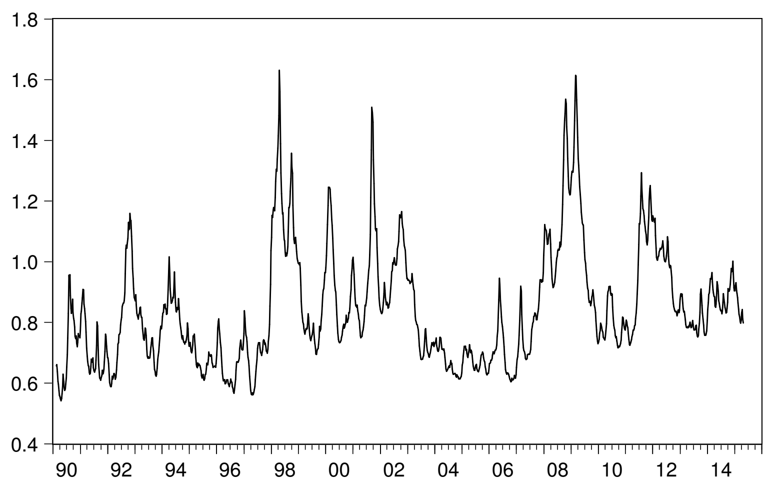
Variable	Horizon	ε_1	ε_2	ε_3
CDS	1	**0.92**	0.06	0.02
	5	**0.82**	0.15	0.03
	10	**0.60**	0.24	0.16
	20	**0.35**	0.25	0.40
Risk Aversion	1	0.00	**0.94**	0.06
	5	0.00	**0.83**	0.17
	10	0.00	**0.74**	0.26
	20	0.00	**0.63**	0.37
Uncertainty	1	0.00	0.00	**1.00**
	5	0.00	0.00	**1.00**
	10	0.00	0.01	**0.99**
	20	0.00	0.02	**0.98**

Table C.4: Forecast error variance decomposition from the MSH-SVAR for Italy (low volatility state)

Variable	Horizon	ε_1	ε_2	ε_3
CDS	1	**0.99**	0.00	0.01
	5	**0.93**	0.01	0.06
	10	**0.75**	0.04	0.22
	20	**0.41**	0.06	0.54
Risk Aversion	1	0.00	**0.91**	0.09
	5	0.00	**0.79**	0.21
	10	0.00	**0.61**	0.39
	20	0.00	**0.42**	0.58
Uncertainty	1	0.00	0.00	**1.00**
	5	0.00	0.01	**0.99**
	10	0.00	0.02	**0.98**
	20	0.00	0.02	**0.98**

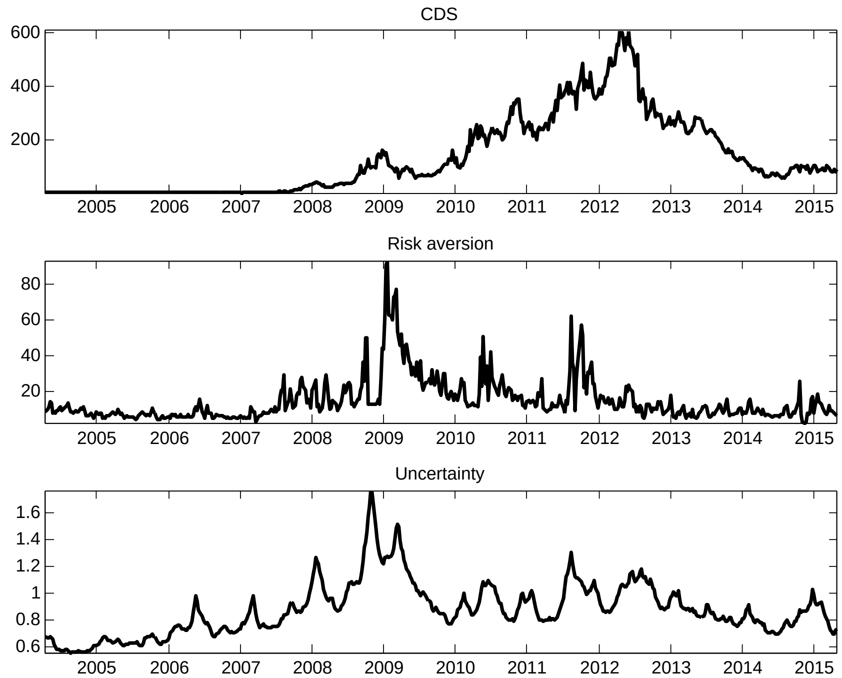


(a) Spain

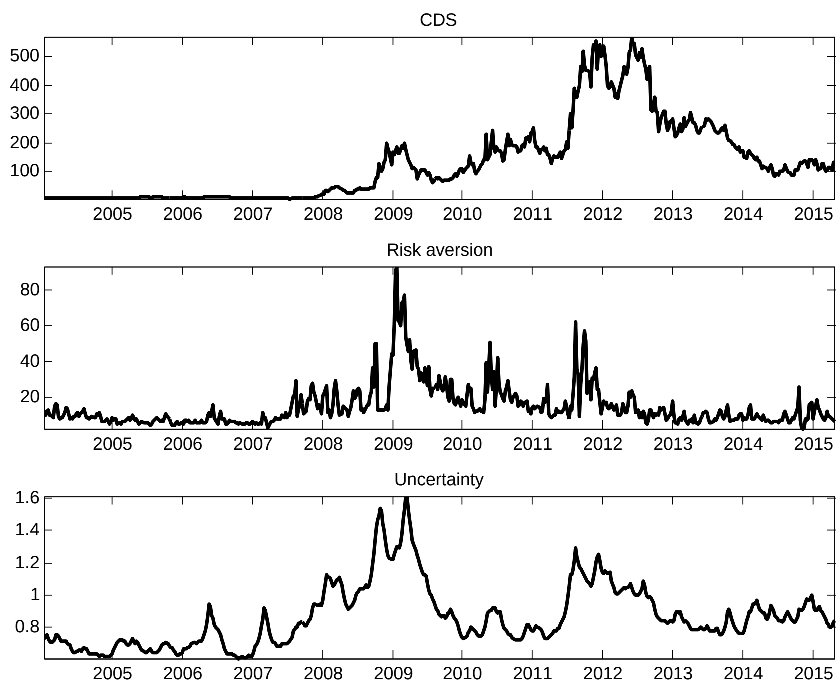


(b) Italy

Figure C.9: Weekly uncertainty index based on the method by Jurado et al. (2015) applied to a large dataset of equity returns.

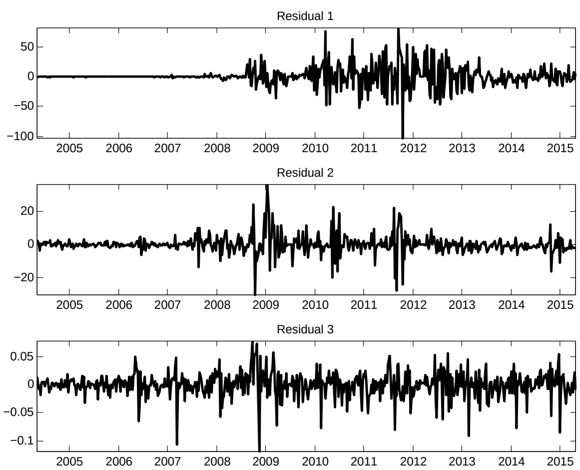


(a) Spain

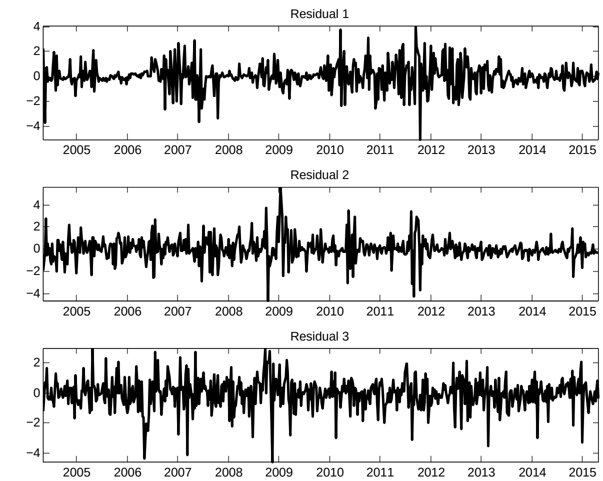


(b) Italy

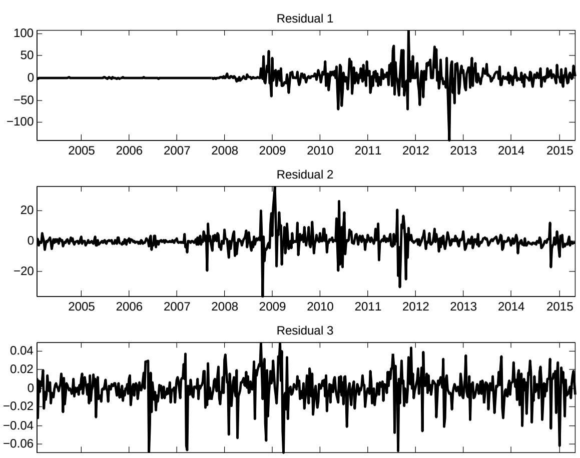
Figure C.10: Endogenous variables entering the MSH-VAR models



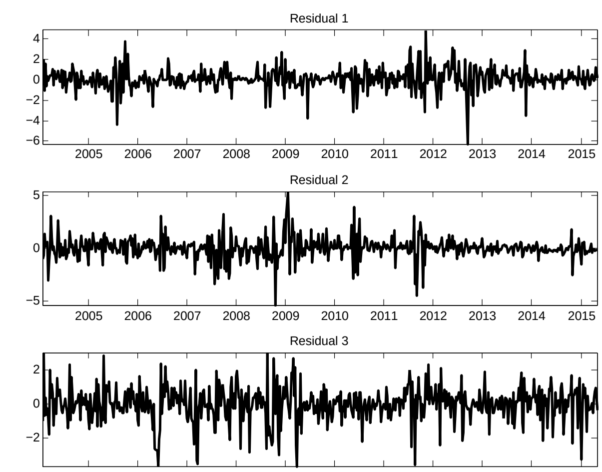
(a) Spain, residuals



(b) Spain, standardized residuals

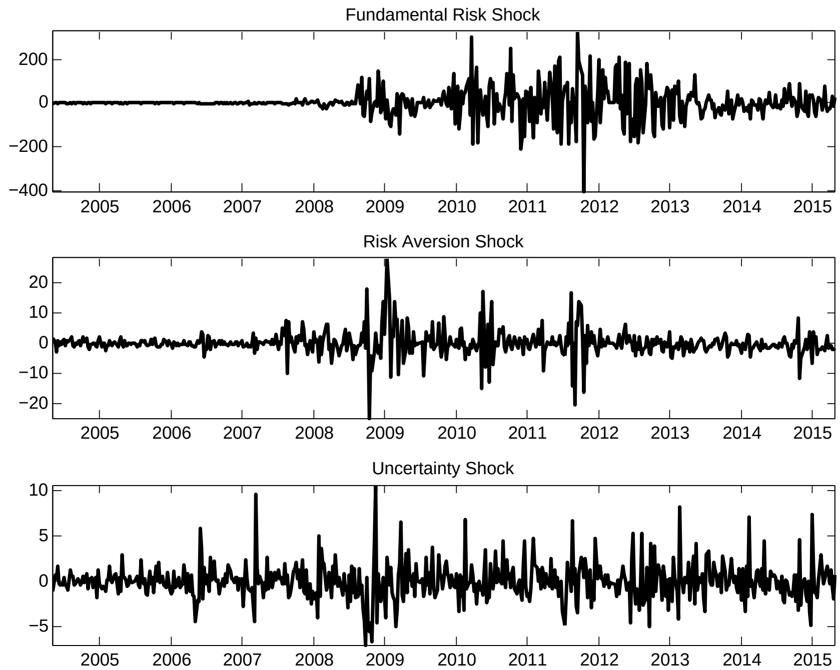


(c) Italy, residuals

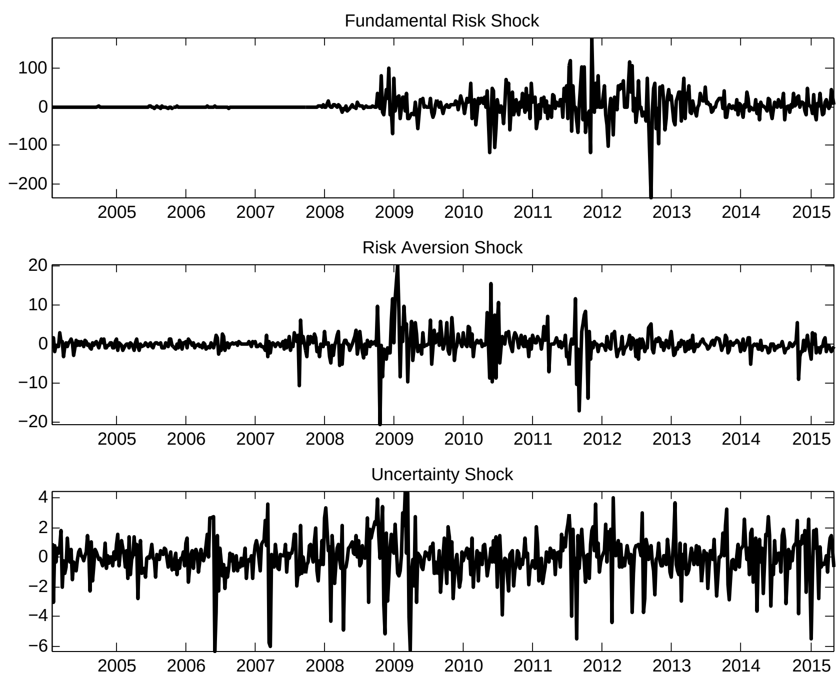


(d) Italy, standardized residuals

Figure C.11: Residuals and standardized residuals from the MSH-SVAR models

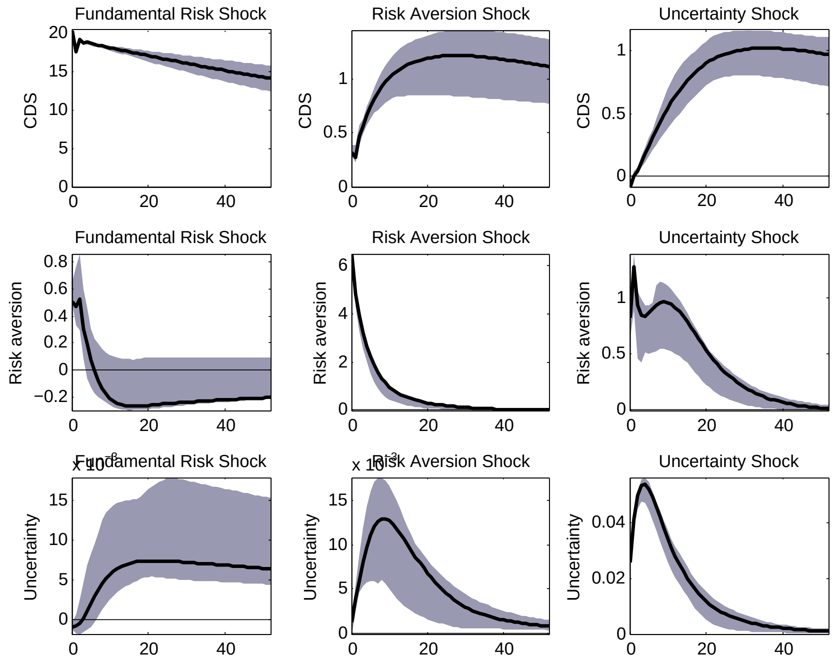


(a) Spain

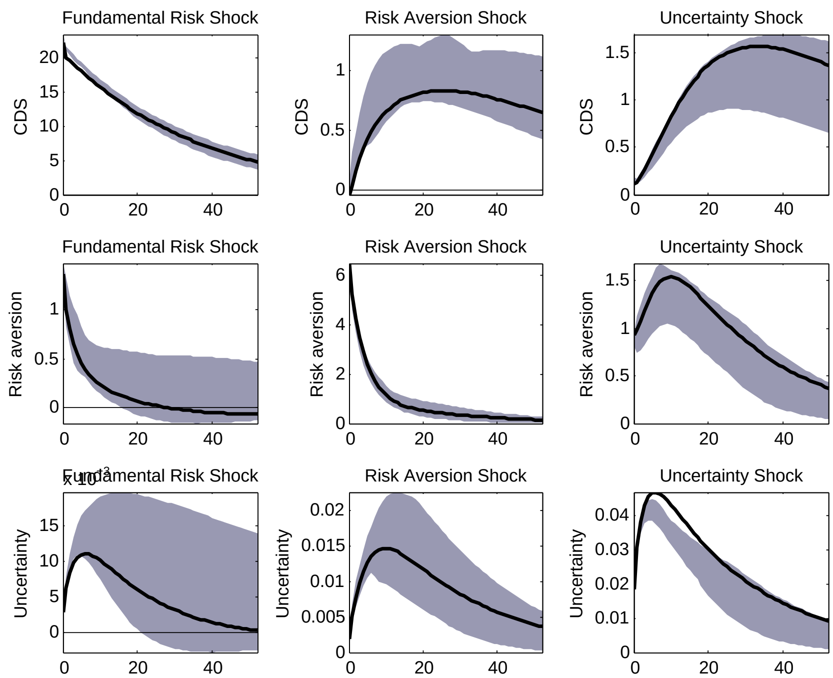


(b) Italy

Figure C.12: Structural shocks uncovered from the MSH-SVAR model



(a) Spain



(b) Italy

Figure C.13: Impulse responses with 68% confidence intervals based on 1000 bootstrap replications, high volatility state