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Identifying and Measuring the Contagion Channels at Work in the European Financial Crises*

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Abstract

We use impulse response functions computed from linear and nonlinear, Markov switching models to investigate the strength of four alternative contagion channels. These are the flight-to-quality, flight-to-liquidity, risk premium, and correlated information channels. We study the differences among estimates and impulse response functions across linear and nonlinear models to identify and measure cross-asset contagion. An application to weekly Eurozone data for a 2007-2014 sample, reveals that a two-state Markov switching model shows accurately estimated but economically weak contagion effects in a crisis regime. These results are mainly explained by a flight-to-quality channel. Furthermore, we extend our analysis the analysis to investigate whether European market may be subject to contagion when exposed to external shocks, such as those originated from the US subprime crisis.

Keywords: Contagion channels, Markov switching models, vector autoregressions, impulse response function, flight-to-quality, flight-to-liquidity, risk premium.

JEL codes: G12, E43, C32.

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1. Introduction

Did the recent, 2010-2011 European sovereign debt crisis represent a glamorous instance of cross-country, cross-asset contagion, in which shocks spread from low credit-quality government bonds to corporate bonds and stock markets? To what extent the observed behavior of nominal yields and spreads (over a risk-free rate) reflected spillovers that went over and beyond what ought to be expected in "normal" times, before financial markets are hit by sizeable shocks, such as the Greek's debt restructuring and Portugal and Ireland's recourse to the IMF and EU bailout funds? Can we establish a link between the asset-backed securities (ABS) crisis originating in the US in 2007-2009 and the subsequent sovereign jitters in Europe? In this paper, we use state-of-the-art econometric methods applied to weekly data to answer these questions and disentangle the channels of contagion at work in the Eurozone during the tumultuous years marked by the Great Financial Crisis in the US and then by the European debt crisis (2010-2011, see Lane, 2012).[1]

Before the recent crisis, the literature had mostly focused on the dynamics of cross-country contagion applied to homogeneous asset markets. For instance, typically papers had asked how large, negative shocks propagate across different national stock markets (see, e.g., Forbes and Rigobon, 2002, Markwat et al., 2009) or across international bond markets (see, e.g., Dungey et al., 2006). Instead, following the 2007-2009 US subprime crisis, researchers have displayed an increasing interest in cross-asset contagion. The subprime crisis represents an ideal episode to study this variant of contagion because the effects of negative shocks to the US ABS market spilled over to other markets. Indeed, a number of papers have attempted to investigate the dynamics of such cross-asset contagion (see, e.g., Longstaff, 2010, and Guo et al., 2011), in addition to cross-country contagion from US to foreign markets (see, e.g., Samarakoon, 2011).

Longstaff (2010) has captured the effects of the US crisis by studying the changes over time in the linkages between asset-backed CDO returns and Treasury, corporate bond, and stock market returns, along with changes in the VIX index in three different periods: the pre-crisis, the subprime crisis, and the global crisis periods. The results of his analysis show evidence of an increase in the linkages in 2007, that is, when the crisis started, compared to the pre-crisis and (subsequent) global

[1] In this paper we shall use the Europe and Eurozone interchangeably as if the two geopolitical entities were the same. Of course, in a technical sense, we refer to the latter entity. In particular, UK data are not considered as they probably do deserve separate attention, see Degryse et al. (2015).

crisis periods. Similarly to ours, the analysis conducted by Longstaff also aims at identifying the contagion mechanisms at work in the subprime crisis, distinguishing among flight–to–quality, flight-to-liquidity, risk premium, and correlated information channels (see Section 2.1 for a definition). However, Longstaff's approach relies on simple regressions that treat breaks as exogenously given, while we deal with the instability in the data through the use of (Markov) regime-switching models. In addition, we do not limit our analysis at identifying contagion episodes, but we also attempt to measure the effects of a shock to one asset class to different national and international markets through the use of impulse response functions (IRFs).

Similarly to our paper, also Guo et al. (2011) have computed IRFs under a Markov switching Vector Autoregressive (MSVAR) model to investigate the effects of contagion among stock, real estate, credit default swap, and energy markets during the subprime crisis within two different states of the US economy: a "stable regime" (high mean returns, low volatility), and a "risky regime" (low mean returns and high volatility). They report significant contagion effects in the "risky regime", while these effects are weaker in the "stable" regime. This finding supports the idea that contagion episodes occur during crises. However, differently from our paper, they do not attempt to disentangle the channels through which shocks propagate within the financial system.[2]

Our objective is to carry out an analysis of cross-asset contagion in European financial markets similar to Guo et al., while using econometric tools that allow us to better identify and characterize the contagion dynamics. In addition, we extend our analysis to cross-country, cross-market contagion, as we investigate spillovers effects of the US subprime crisis to European financial markets. To this purpose, we perform two distinct simulations: of a shock to peripheral (an equally weighted portfolio of GIIPS, i.e., Greece, Ireland, Italy, Portugal, and Spain) sovereign yields similar to the one that occurred during the 2010-2011 European crisis; of a shock to US low-quality (Bbb, to mimic subprime features) ABS yields, similar to the one that was possibly imported in Europe between 2007 and 2008, after Lehman's demise. Indeed, we estimate the IRFs generated by such negative shocks to low-grade sovereign bonds and foreign ABS markets in a single-state

[2] There are few other papers that discuss contagion that, similarly to our, use regime-switching models to deal with the instability of the data. For instance, Kenourgios et al., (2011) have used MS to test for cross-country contagion among developed and emerging markets. Philippas and Siriopoulos (2013) have used MS models to show contagion effects in volatility that occurred beyond simple spillovers caused by integration across the EMU bond markets, during the sovereign debt crisis.

vector autoregressions (VAR) and in a multi-state MSVAR models to study the effects of these two shocks on European assets such as corporate bonds, equity and repo rates.

We find two key results that refer to a crisis shock to peripheral European yields. First, a two-state Markov switching (henceforth, MS) model that dominates simple VARs in a statistical perspective, gives evidence of accurately estimated but economically weak contagion effects, limited to a crisis regime. In particular, Bbb corporate bonds (both short- and long-term) and equity valuations (as captured by the dividend yield) are somewhat affected. Consistent with a common sense prior, the effect on core sovereign yields tends to be modest, it is not precisely estimated, and it declines to show negative effects (as a result of a "flight-to-quality" dynamics) rather quickly. Second, this mild evidence of contagion is mostly explained by a flight-to-quality channel being active in a few periods during 2010 and 2011, when – whilst the yields of investment grade, Aaa corporate bonds and (at least eventually) core European government bonds declined – the yields of "junk", Bbb corporates and on equities increased. There is some evidence of also flight–to–liquidity and correlated information channels (measured as the difference between single-state and MS IRFs) being at work during the European sovereign crisis. We obtain instead no evidence of the presence of a risk premium channel.[3]

A few more papers relate to ours and found evidence of contagion, especially in what they define to be a crisis state. De Santis (2014) has used VECM-based IRFs to demonstrate that Greek credit downgrades affected other European bond spreads over a sample September 2008 – August 2011, even though the effects turn out to be economically small. Arghyrou and Kontonikas (2012) have used monthly data on bond spreads over German bunds to document the existence of contagion effects in the sovereign debt crisis, particularly among EMU peripheral countries. Antonakakis and Vergos (2013) used VAR-based IRFs to show that shocks from the periphery have, on average, three times the destabilizing force on other countries than shocks coming from the core. Afonso et al. (2011) have examined whether sovereign yields and CDS spreads in one country react to rating

[3] When, in Section 5, we extend this analysis to a subprime shock originating in the US, there is evidence of spillover effects from a US ABS shock to European markets, but no evidence of contagion in an economic sense. A few European markets do react to US ABS shocks, but moving in an opposite direction. Similarly to ours, the analysis of bond yields by Caporin et al. (2012) points out a change in the intensity of the propagation of shocks in the 2008-2011 post-Lehman sample, but the coefficients actually declined, indicating the opposite of contagion.

announcements of other countries and concluded that there is evidence of contagion, especially from lower rated to higher rated countries.

On the opposite, a few papers have questioned the existence of contagion in the European sovereign crisis. For instance, Beirne and Fratzscher (2013) have shown that a sharp rise in the sensitivity of government bond yields to fundamentals is the main explanation for the rise in yields and CDS spreads during the crisis. By contrast, they indicate that regional contagion has been less important. Kalbaska and Gatkowski (2012) have studied contagion among European sovereign bonds using CDS data and found that GIIPS countries triggered very little or no contagion among the Euro area countries during the 2005-2010 period.

The rest of the paper is structured as follows. Section 2 reviews our methodology by providing details on the econometric models, on the nature of the contagion channels, and on the way in which IRFs may be used to identify alternative channels. Section 3 introduces the Eurozone data. Section 4 reports the main empirical results. Section 5 expands the analysis to investigate whether European market may be subject to contagion when exposed to external shocks, such as those originated from the US subprime crisis in 2007-2008. Section 6 concludes.

2. Research Design

2.1. Alternative channels of financial contagion

The literature on financial contagion provides a number of definitions of the phenomenon. In essence, four strands of research map into four distinct and yet complementary (and, to some extent, overlapping) characterizations (see Forbes, 2012): (i) episodes in which following a shock, its transmission is in excess of what can be explained by fundamentals (see, e.g., Bekaert, Harvey, and Ng, 2005, Corsetti et al., 2005, Pritsker, 2001); (ii) the transmission of shocks is different from regular adjustment typical in quiet "regimes", as in Forbes and Rigobon (2002) who capture this phenomenon through increased correlations during times of distress (see also Beirne and Fratzscher, 2013); (iii) the events constituting contagion represent negative extremes of the empirical distribution and imply excessive tail co-dependence (see, e.g., Longin and Solnik, 2001); and (iv) the transmission of shocks is sequential from the epicenter to markets that are subsequently hit by contagion, for example in a causal sense. Yet, there is no agreement about which ones of these four criteria are necessary or sufficient to characterize a contagion event. In our paper, we shall adopt a view of contagion that reflects features (ii)-(iv). On the opposite, we shall not take a stand of what "normal" transmission justified by fundamentals may represent, because when it comes to asset prices little is known about what fundamentals really are and what portion of the volatility of returns these can explain (see Beirne and Fratzscher, 2013).

Recently, researchers have made progress by isolating and measuring the strength of four distinct propagation channels: the correlated information, the flight–to–liquidity, the flight–to–quality, and the risk premium channels. Because our empirical analysis adopts this perspective, in what follows we describe these channels. Under the *correlated information channel*, a shock to one market provides information that is then incorporated also by the equilibrium prices of other assets that are not directly affected by the shock. The idea behind this contagion mechanism is that price changes in one market may be perceived as relevant by investors for the valuation of other markets. Because investors immediately adjust their beliefs, prices in other markets change as well. Such a channel has also been dubbed "wake-up call" (see e.g., Bekaert et al., 2014) and, starting with the seminal paper by King and Wadhwani (1990), it has been used to explain episodes of (almost) simultaneous drops in asset prices in different countries.

Under the *flight-to-liquidity channel*, following a shock to one market, agents' preferences shift towards more liquid securities (see e.g., Beber et al., 2009). For instance, Brunnermeier and Pedersen (2009) have developed a model in which the negative spiral in market liquidity that follows a shock originates from variations in traders' funding liquidity.[4] Vayanos (2004) has considered the flight–to–liquidity phenomenon under the perspective of fund managers that execute portfolio strategies taking into account the risk of withdrawals by the individuals investing in their funds. Therefore, during periods of increased uncertainty and volatility, managers are less willing to hold illiquid securities which leads to an upward adjustment in the premium investors recognize to illiquid financial instruments.

A third channel of contagion, generally referred to as *flight–to–quality channel* in the literature (see e.g., Caballero and Kurlat, 2008), identify episodes in which, following a shock to one

[4] In particular, trading in financial markets requires capital and traders can use securities as collateral to borrow funds, although the amount obtained is subject to a haircut (or margin) applied to the value of the collateralized assets. Brunnermeier and Pedersen argue that, when a shock to one market leads to an increase in the volatility of asset prices, the margins required by lenders will increase as well. This reduces the availability of funding to traders and forces them to trim their positions in capital-intensive securities.

market, investors attempt to substitute risky assets in their portfolios with safer ones. Finally, under the *risk premium channel,* shocks to one market lead to a generalized increase in the risk aversion of market participants. This generates, in turn, an increase in the risk premium of all assets. Of course, this mechanism requires either preferences or the quantities of undiversifiable risk to be time-varying. For instance, Longstaff (2010) has explained the effects that negative returns in one market have on subsequent returns in other markets by way of time-varying risk premia. Kyle and Xiong (2001) have proposed a theoretical framework in which the source of the increase in market risk aversion, and thus of contagion through the increase of asset risk premium, is a net worth effect through the balance sheet of financial intermediaries.

2.2. Econometric models and their use in disentangling contagion channels

We use two sets of econometric models to assess the strength of the four alternative channels introduced above, standard single-state vector autoregressive (VAR) and multi-state, MSVAR models. Under both models we compute impulse response functions (IRFs). In particular, the comparison between regime-specific IRFs computed under a MSVAR model and the single-state ones allow us to disentangle the correlated information channel. Indeed, such channel can be seen as a simultaneous switch of several markets or countries to an extreme state of poorly performing markets and high volatility, which fails to be captured by the single-state model. To quantify and discuss the rest of the channels, we mainly rely on regime-switching IRFs only. As it is well known, adopting MS models as a working tool offers a range of benefits as they are able to capture features of the series that a single–state VAR fails to feature, including fat tails, heteroskedasticity, skewness, and time–varying correlations (Ang and Timmermann, 2012).

The standard (single-state) VAR model is defined as follows:

$$y_t = A_0 + \sum_{j=1}^{p} A_j y_{t-j} + u_t, \quad u_t \sim IID\ N(0, \Sigma_u), \tag{1}$$

where p indicates the number of lags, $y_t = \left(y_{1,t}, \dots, y_{N,t}\right)'$ is a $N \times 1$ random vector of endogenous variables, $A_0 = \left(a_{1,0}, \dots, a_{N,0}\right)'$ is a $N \times 1$ vector of intercepts, A_i for $i = 1, .., p$ are the $N \times N$ vector autoregressive coefficient matrices, and $u_t = \left(u_{1,t}, \dots, u_{N,t}\right)'$ is a N–dimensional white noise innovation process, such that $E(u_t) = 0, E(u_t {u'}_t) = \Sigma_u$, and $E(u_t {u'}_s) = 0$ for $s \neq t$. The Reader interested in a more in-depth review of a standard VAR models may refer to Enders (2008).

Instead, a k–regimes Markov switching VAR (henceforth, MSVAR) process with heteroskedastic components, compactly MSIAH(k, p) (Markov switching intercept autoregressive heteroskedasticity), is defined as follows:

$$y_t = A_{0,S_t} + \sum_{j=1}^{p} A_{j,S_t} y_{t-j} + \Omega_{S_t}^{1/2} e_t, \qquad e_t \sim IID\ N(0, I_N), \tag{2}$$

where $S_t = 1,2, \dots, k$, k is the number of regimes, p is the number of VAR lags, A_{0,S_t} is the $N \times 1$ vector collecting the k regime–dependent intercepts, and $A_{1,S_t} \dots A_{j,S_t}$ are the regime–dependent $N \times N$ autoregressive coefficient matrices. $\Omega_{S_t}^{1/2}$ is a lower triangular matrix and represents the factors applicable to the regime S_t in a state–dependent Choleski decomposition of the covariance matrix Ω_{S_t}. In our specification of MS models, we assume that alternative states are possible, that is, $k > 1$, and that regimes are hidden, meaning that at all times, investors fail to observe S_t. Moreover, in MSVAR models, the state S_t is assumed to be generated by a discrete–state, homogeneous, irreducible, and ergodic first–order Markov chain with transition probabilities

$$Pr\left(S_t = j | \{S_j\}_{j=1}^{t-1}, \{Y_\tau\}_{\tau=1}^{t-1}\right) = Pr(S_t = j | S_{t-1} = i) = p_{i,j} \in (0,1), \tag{3}$$

where $p_{i,j}$ is the generic $[i, j]$ element of the $k \times k$ transition matrix P with elements

$$p_{i,j} = Pr(S_{t+1} = j | S_t = i), \qquad \sum_{j=1}^{k} p_{ij} = 1 \quad \forall_{i,j} \in \{1, \dots, k\}. \tag{4}$$

The elements of the main diagonal of the transition matrix, $p_{i,i} = Pr(S_{t+1} = i | S_t = i)\ \forall_i \in \{1, \dots, k\}$, estimate the probability to remain in regime i in two consecutive periods and allow us to capture a persistence in the data that is not linear.[5]

MS models are estimated by maximum likelihood (MLE) and estimation is performed through the Expectation–Maximization (henceforth, EM) algorithm proposed by Hamilton (1990). Given the matrix Y_{t-1}, which collects lagged values of the variables, and a regime ξ_t, the density function of y_t conditional on the realization of the regime k is Gaussian:

$$p(y_t | S_t = i, Y_{t-1}) = ln(2\pi)^{-1/2} \ln|\Omega|^{-\frac{1}{2}} exp\{(y_t - y_{k,t})' \Omega_k^{-1} (y_t - y_{k,t})\} \tag{5}$$

If we consider that the information set available at time $t - 1$ includes only the pre–sample values collected in Y_{t-1}, the sample observations, and the states of the Markov chain up to S_{t-1}, then the

[5] The model in (2) requires us to estimate a large number of parameters, in particular, if the number of variables included in the system is large. As an alternative, it is possible to estimate models that require a lower number of parameters than a fully-fledged MSIAH(k,p) framework. For example, in a MSIH(k,0) (Markov switching intercept heteroskedasticity) we have $p = 0$ and only the intercepts and the covariance matrix of the error terms are regime–dependent. Our specification search selects instead a MSIH(k, p), with $p > 0$ but the VAR coefficients matrices not linked to the state variable.

conditional density of y_t is a mixture of normal distributions:

$$p(y_t|S_{t-1} = i, Y_{t-1}) = \sum_{j=1}^{k}\sum_{i=1}^{k} p_{i,j}\left(ln(2\pi)^{-1/2}\ln|\Omega|^{-1/2}exp\left\{\left(y_t - \bar{y}_{k,t}\right)'\Omega_k^{-1}\left(y_t - \bar{y}_{k,t}\right)\right\}\right). \quad (6)$$

The information about the Markov chain is collected in the vector ξ_t. Because at time $t-1$ the only information available is the realized time series, the unobserved regime vector ξ_t needs to be estimated alongside the parameters. The corresponding estimates are collected in the vector $\hat{\xi}_{t|\tau}$,

$$\hat{\xi}_{t|\tau} = \begin{bmatrix} \Pr(S_t = 1|Y_\tau) \\ \vdots \\ \Pr(S_t = k|Y_\tau) \end{bmatrix} \quad (7)$$

to include the probabilities of being in regime k given the information set Y_τ. If we collect the densities of y_t conditional on S_t and Y_{t-1} in the vector η_t, the conditional probability density of y_t given Y_{t-1} in (6) can be written as $p(y_t|Y_{t-1}) = \eta_t' P' \hat{\xi}_{t-1|t-1}$, where $\eta_t \equiv [p(y_t|\xi_t = 1, Y_{t-1})$ ... $\mathrm{p}(y_t|\xi_t = k, Y_{t-1})]'$. Following the same derivation applied to the single observation y_t, we derive the conditional probability density of the whole sample. The EM algorithm can be used to carry out an iterating process to jointly estimate the parameters and the Markov state probabilities.

According to the general definition, an IRF represents the difference between the conditional expectation of y_{t+h} at time t in case y_t has been subject to a shock and the conditional expectation of y_{t+h} at time t in case y_t has not been subject to any shock. In practice, we can define the h–step ahead IRF as follows[6]

$$IR_{\Delta u}(h) = E[Y_{t+h}|y_t(\omega')] - E[Y_{t+h}|y_t(\omega)], \quad (8)$$

where the sample path $y_t(\omega')$ differs from the sample path $y_t(\omega)$ because the initial value of y_t has been subject to a shock Δu (see Potter, 2000). This general definition can be extended and adapted to a MS framework. In this case, we obtain the following representation:

$$IR_{\Delta u}(h) = E[Y_{t+h}|\xi_t, u_t + \Delta\, u_t; Y_{t-1}] - E[Y_{t+h}|\xi_t, u_t + \Delta\, u_t; Y_{t-1}]. \quad (9)$$

The h–step ahead IRF thus depends on the state prevailing at time t, when the shock occurs. However, when computing IRFs in a MS framework we need to deal with the additional issue that regimes are latent and therefore the prevailing state at time t is unobservable. For this reason, we compute regime–dependent IRFs assuming that the regime prevailing at the time the shock occurs is known. Both reduced-form VAR and MSVAR models are subject to identification problems. Therefore, we apply a Choleski decomposition to the regime–dependent covariance matrices (as

[6] For concreteness, we focus on the case of a MSIH(k, p) model, the one selected by our specification search.

we do in the case of single –state model).[7] In addition, to take into account the uncertainty of the estimated values, for each IRF we also construct the appropriate confidence intervals through Monte Carlo simulation techniques.[8]

To identify each financial contagion channel, we perform qualitative comparisons of the IRFs obtained from the different models. The idea is to exploit the different information captured by the MSVAR model for yields, as well as the single–state VAR, to distinguish different contagion mechanisms. The first channel that we address is the flight–to–liquidity channel. Based on the definition discussed in Section 2.1, the evidence of this channel being at work can be retrieved by observing the differences in the estimated responses to a shock between liquid and illiquid assets. Under the risk premium channel, contagion occurs because shocks to one market lead to an increase in the risk aversion of financial market participants, thus triggering an upward adjustment of the risk premia on all the risky assets in the economy. Accordingly, for this channel to be a driver of contagion we would expect positive and significant estimated responses of the spreads and, consistently, of the yields of all the other assets to a shock in one market. Under to a fight–to–quality channel, following a shock to one market, investors attempt to sell risky assets and purchase safer assets. Consequently, the price of the former declines (yield rises), while the price of the latter increases (yield falls). The evidence of a flight-to-quality being a contagion channel is given by the presence of an increase of the price (decrease of the yields) of the safest assets (e.g., Aaa graded bonds) accompanied by a decrease of the price (increase in yields) of low quality (e.g. Bbb graded bonds) assets. In contrast, under a risk premium channel, the risk premium of all assets is supposed to increase. In order to disentangle the effects of this channel, we repeat the analysis that we performed for the yields for European yield spreads. Finally, under the correlated information channel, contagion occurs because the negative shock to one market conveys information that investors perceive as relevant for the pricing of other assets. This generates an immediate effect because the evidence of a shift to a crisis state in one market triggers an adjustment of investors'

[7] A Choleski triangular factorization allows to solve the identification problem without imposing structure. Because it forces asymmetries in the model, the ordering of the variables becomes crucial. To control for this drawback, we apply different orderings to the series and verify that the results are stable.

[8] For instance, in the case of $p = 1$, we assume that the matrix of the VAR coefficients A_1 follows an asymptotic multivariate normal distribution $A_1 \sim N(\widehat{A_1}, \Sigma)$, where $\widehat{A_1}$ represents the estimate of the true but unknown coefficient matrix and Σ is a diagonal matrix consisting of the squares of the standard errors of the estimates. The matrix A_1 is the extracted from this distribution a large number of times and is used to compute the IRFs and to construct the upper and lower bounds of the 95 per cent confidence interval.

beliefs concerning other asset prices. We therefore identify this form of contagion with the non–linear and instantaneous effect captured by the MSVAR, due to the possibility that the components of yield series may move in the same direction when a shift to a given regime occurs. In particular, because we aim at modelling contagion during financial crises, we are interested in the effect generated by a shift to the crisis regime. The time–invariant nature of the single–state VAR instead fails to capture a similar effect. Therefore, we can isolate and measure the contribution of the correlated information channel through the difference between the values of the IRFs computed under the MSVAR and the single–state VAR frameworks. Because this effect is immediate, we shall limit our discussion to the values of the IRFs estimated in the two frameworks to the first few weeks after a shock hits in the crisis state.

3. The Data

We collect weekly data to span a March 23, 2007 - December 19, 2014 sample.[9] The sample mainly includes European yield series (i.e., sovereign and corporate bonds, repo contracts, and stock yields). The sovereign bond yields are collected in two equally weighted portfolios (as in Beirne and Fratzscher, 2013) concerning core vs. periphery/low-credit quality (high credit risk) countries, respectively. The yields concern 10-year government bonds. The core countries are Austria, Belgium, France, Finland, Germany, and the Netherlands. The periphery consists of Greece, Ireland, Italy, Portugal, and Spain (the so called GIIPS). The repo rate concerns transactions in which the sale of long-term German bunds at a certain price is combined with the agreement to repurchase them at a higher price at maturity, and repo rates paid by borrowers are given by the difference between these two prices. When needed, we use such a repo rate as the euro-denominated riskless rate. Corporate bonds data are by Bank of America-Merrill Lynch (BAML) and concern high quality (Aaa) short/medium (up to 5 year maturities) vs. long (10 year and longer maturities) portfolios and low quality (Bbb) short vs. long (identical definitions apply) EMU corporate portfolios, for a total of eight different yield series.[10] Equity market data consists of Eurostoxx 600 dividend-to-

[9] Weekly data strike a balance between the availability of a sufficient number of observations and tolerating an intermediate amount of conditional time variation in second moments. Interestingly, a good fraction of papers concerning crisis shocks transmission within the Euro-system borders are based on weekly time series, see, e.g., Holló et al. (2012) and references therein.

[10] All data are from Datastream. In the case of Irish government bonds, we cannot download market yields for the period October 2010 - March 2013 as trading did not occur. For this period, GIIPS equal-weighted portfolio only includes the remaining four peripheral countries.

price yields.[11] Finally, one of the exercises that we perform also include high- vs. low-credit quality US-issued ABS, of which we measure the yields using two indices also calculated by BAML, Aaa–rated ABS and instruments in the Aa–Bbb bracket.[12]

Table 1 reports summary statistics in panel A and correlations in panel B. Data fail to show any surprising features: GIIPS yields are on average much higher (6.37%) than core yields. In particular, GIIPS yields peak at 15.25% during the sovereign crises, while core sovereign yields reach a maximum of 4.83%, meaning that during the crisis, spreads between GIIPS and core yields must have climbed even higher and beyond the 324 bps that Table 1 implies on average. Peripheral government yields are also characterized by a volatility that is almost three times larger than core's. As one would expect, high quality corporate bonds always give yields on average higher than low quality ones, both for short-term (4.64% vs. 2.65%) and long-term securities (5.39% vs. 3.80%); these differences in means are matched by differences in volatilities. All series are characterized by pervasive non-normalities, although in this case more as a result of the widespread right-skewness of the data, than of their fat tails. As discussed in Guidolin (2011), this is consistent with our speculation that data may contain regime shifts. Finally, in panel B of Figure 1, most series are positively and significantly correlated, with the only exception of GIIPS that seem to follow a process of their own and to negatively correlate with investment grade corporate yields, and the German bund repo rate, that is, with the highest quality assets. This may represent an early indication of flight-to-quality effects which we better investigate later on.

4. Key Empirical Findings

4.1. No European Contagion in Single-State Models

We start our analysis by asking whether, in the face of the peripheral sovereign yield shocks recorded during 2010 and 2011 in Europe, contagion channels similar to the ones that we have characterized in Section 2 and that have been proposed in the literature were active. A positive evidence would validate the notion that the four channels described above represent a general structural feature of the way financial systems absorb and propagate shocks.

A range of standard information criteria (IC) that trade-off in-sample fit with model

[11] Equity data are converted into 12-week rolling window 3-month dividend yields. All of our series measure ex-ante asset returns. Hence we use bond yields and the equity dividend yield following Campbell and Shiller's paradigm (e.g., Shiller, 2007, relates the dividend yield to nominal and real interest rates).

[12] We do not resort to lower grade ABS data because before 2007 they were rare and time series unreliable.

parsimony (hence, likely out-of-sample performance), provide heterogeneous indications as to the appropriate number of VAR lags. The latter range from 10 in the case of the Hannan-Quinn IC, to four in the case of the Akaike IC, to two in the case of the most parsimonious Schwarz, Bayesian IC. However, only models with three or less lags guarantee saturation ratios – defined as the ratio between the total number of observations across all time series to be used in estimation and the total number of parameters to be estimated – of approximately 20 or more. We therefore select a VAR*(2)* model. An Appendix available upon request reports specification search results and parameter estimates for this model. We note that only approximately one-third of the estimated VAR coefficients are statistically significant in tests of 10% size or lower. Therefore, in this perspective, there is relatively weak evidence of predictability of fixed income and equity yields in Europe, as this can be captured from simple VAR models. Yet, the R-squares that we have obtained range between 97.4 and 99.5 percent, even in adjusted terms, which is an indication of the almost perfect fit to the data provided by a *VAR(2)*. The most predictable series – in terms of yielding statistically significant coefficients (as all adjusted R-squares are systematically high) – are Bbb corporate yields and sovereign bonds. Importantly, we also find that the system is stable (and therefore stationary) as the largest root of the characteristic polynomial has a modulus of 0.991.

Figure 1 shows the IRFs computed from the *VAR(2)* model in Table 2. A comment comes rather naturally from observing the plots: apart from the high persistence of the shock onto peripheral yields (which is consistent with an own first-order serial VAR coefficient of 0.964), all other yields are scarcely affected and only short-term, investment grade corporate yields decline as a result *of a peripheral sovereign shock* (but only by 2-3 bps) in a statistically significant manner between weeks 6 and 16 after the shock. All other reactions to the primitive impulse are instead imprecisely estimated. Therefore, a single-state VAR model implies the absence of any contagion effects deriving from a sovereign debt crisis. This flies directly in the face of the empirical evidence that most investors, policymakers, and commentators have experienced between the Summer of 2010 and the Fall 2011 when a state of turmoil engulfed European financial markets, with clear evidence of contagion and spillover effects (see e.g., Arghyroua and Kontonikas, 2012, and De Santis, 2014). Consequently, in Section 4.2, we ask whether expanding our analysis to an MS model may yield results that are consistent with some earlier literature.

4.2 *Markov switching model selection and estimates*

We estimate a number of *MSVAR*(*k,p*) models. The specification search presented in Table 2 indicates that a MSIH(3,1) (i.e., featuring three regimes but time invariant *VAR* matrix) model is selected for European yields. In fact, both the HQ and the AIC criteria imply that a MSIH(3,1) model optimally trades-off in-sample fit with the promise of out-of-sample predictive accuracy. In the table, the null of a single regime is always rejected also when the issues caused by nuisance parameters are taken into account, using Davies' (1977) correction applied to a standard chi-square test.

Table 3 reports estimated parameters from the MSIH(3,1) model just selected. Interestingly, the regimes are predominantly identified by the volatility of shocks to yields. In fact, the regime-specific intercept coefficients are not always different across regimes, which indicates – in the presence of a time invariant matrix of VAR coefficients – that first moments hardly help in the definition of states. The VAR matrix is stable and this ensures stationarity of the process. Yet, the three regimes are considerably persistent. The low-volatility regime 1 has a duration of six weeks and Figure 2 shows that it characterizes a number of extended periods. However, regime 1 turns out to persistently characterize the period following the Summer of 2012, when the European sovereign crisis was eventually tackled with force by the European Central Bank (ECB).[13] Table 3 shows that the first regime is marked by below-average correlations between shocks to yields, and in particular to shocks to core and peripheral equally-weighted sovereign yields, and all shocks vs. the equity dividend yield. This means that in this regime, high- and low-quality sovereign bonds are segmented from stock markets and from each other.

Regime 2 is instead a higher volatility state, marked by considerable persistence (28 weeks on average), that on the basis of Figure 2 appears to characterize the 2007-2008 period (apart from a couple of isolated spikes in 2011), when the financial crises affected mostly the US and in continental Europe was possibly perceived as an episode of turbulent markets. In fact, it is in early 2009 that the crisis spread from the US corporate and ABS markets to fixed income and stock

[13] In May 2010 the ECB created the European Financial Stability Facility (EFSF) as a temporary facility to provide loans to euro area Member States. In addition, in June 2011 the European Stability Mechanism (ESM) was set up as a permanent crisis-handling mechanism. Constancio (2012) provides a review of these programs and other unconventional measures adopted by the ECB between 2008 and 2012.

markets on a global scale. Regime 2 presents a volatility of shocks to yields that is between 2 and 10 times larger than in regime 1. For instance, the volatility of the STOXX 600 dividend yield increases from 7 bps per week to 20 bps. In this state, shocks to core and peripheral sovereign yields become highly correlated (0.90), an indication that general, non-sovereign financial crises do move all sovereign bond yields in the same direction – presumably, down – as a flight-to-quality phenomenon occurs, as we shall see in Section 5. Anecdotal evidence (see Beirne and Fratzscher, 2013, for some pre-crisis empirical evidence) suggests that Italian and Spanish government bonds were subject to heavy purchases in 2008 as much as German bunds, effectively inflating what was to be perceived ex-post as a fixed income bubble in European government paper at large (see e.g., Oliveira et al., 2012).[14] Because this regime is marked by the Great Financial Crisis also spreading to stock markets around the globe, dividend yield shocks now appear to be positively correlated with most other series. As it is typical of situations of financial turmoil, shocks to repo rates are negatively correlated with other yields, an indication of a second layer of flight-to-quality *within* European markets, where it is plausible that investors may have unloaded positions in risky assets (long-term and junk corporate bonds) to enter in safer, short-term cash positions.

Finally, regime 3 represents a local, European sovereign crisis state characterized by high volatility in the sovereign bond markets, especially peripheral ones, the volatility of which shoots from the 19 bps per week of the tranquil period to a stunning 60 bps per week. However, the volatility of other yield series is comparable (or occasionally lower) to the one recorded for the second regime, in which market turmoil was not specifically originating from Europe. For instance, in regime 2 the spread between core and periphery bond yields is of only 1.2 bps while in regime 3 it reaches 48.6 bps (i.e., 3.5% in annualized terms). The regime is also moderately persistent (with an implied average duration over 4 weeks). On the basis of Figure 2, we appreciate that this state did start to occasionally best fit the data around mid-2009 in correspondence to rising doubts on the sustainability of the debt burden of a few peripheral European countries (at first Ireland and Greece, and later also Portugal). However, this regime effectively characterizes most of the weeks

[14] A few commentators have speculated that a market perception of an implicit bail-out guarantee, or simply ignorance of country-specific fundamentals may have been the explanation for such hardly rational dynamics., see, e.g., Philippas and Siriopoulos (2013) document a pre-crisis (2010) bubble in peripheral EU bond.

falling between early 2010 and the Spring of 2012, in correspondence to the worst bouts of the sovereign jitters. Moreover, pair-wise correlations between shocks decline relative to the high-variance regime, and in particular core and peripheral bond yield shocks now become negatively correlated (-0.10, although this coefficient is not precisely estimated), which is consistent with common expectations on sovereign market-induced disorders, when a decoupling occurs between core and GIIPS Treasuries.

The time homogeneous VAR matrix estimated in Table 3 reveals that Aaa corporate bond, repo, and peripheral sovereign rates are particularly predictable, in the sense most of the VAR(1)-type lagged coefficients are precisely estimated. Of course, as commonly found in the VAR literature applied to interest rates, all own- (partial) first-order serial correlation coefficients are estimated to be large and highly significant. Yet, capturing nonlinear dynamics through a MSVAR framework as we do in this Section, does not imply that linear predictability stops being detectable. In particular, lagged values of the peripheral government bond and dividend yields accurately forecasts subsequent movements of most series. Interestingly, lower dividend yields today, presumably deriving from higher equity valuations, forecast higher subsequent yields on fixed income securities. Therefore, equity and bond markets tend to move inversely with each other, which reflects simple and yet popular switching asset allocation strategies.

4.2. Alternative channels of contagion in Markov switching models

Figure 3 shows the IRFs resulting from a one-standard deviation positive shock to peripheral sovereign yields, to simulate the effects of a sovereign crisis. The Cholesky ordering that is adopted is in our view the natural one that puts the riskiest markets on top (i.e., Bbb corporate paper, stocks, and GIIPS sovereign rates) where most trading and news are likely to be processed and the least risky assets at the bottom of the ordering. The overall effects shown by the figure are the ones we would expect: in regime 3, which we have interpreted as a local crisis state, all "risk-on" assets are hit by a contagion from a low-credit quality sovereign shock in the crisis state. In the remaining two regimes, the responses are muted and hardly distinguishable from zero, as revealed by 90% confidence bands that generally include a zero response effect.

In general, and apart from local crisis phases, European markets appear to be largely disconnected from each other, and contagion does not represent a first-order effect or concern.

Therefore, in the following we limit our comments to IRFs that pertain to the crisis state. In particular, Bbb corporate bonds (both short- and long-term) and equities (as signaled by their implicit dividend yield) are somewhat hit by contagion, although the overall effect tends to be moderate. For instance, a shock that increases GIIPS sovereign yields by approximately 60 bps on a given week, causes an increase in short-term Bbb yields that is precisely estimated, starts out at less than 1 bp but gradually increases to 3 bps after 6 months. However, confidence bands tend to remain wide and effects as large as 7 bps cannot be ruled out. Although these effects are not as prominent as one may expect, a 3 bps per week may be (questionably) scaled up to exceed 1.5% on an annualized basis (on a portfolio that yields on average 4.6% per year). As one would expect, the effect on core sovereign yields tends to be modest, it is not precisely estimated, and it declines quickly to show negative effects (i.e., as a result of a "flight-to-quality"), consistently with Gorea and Radev (2013). Yet, the true "flight-to-quality" seems to concern Aaa corporate yields, that tend to decline as a result of a sovereign peripheral crisis.

Next, we proceed to identify and measure the alternative channels of financial contagion. The evidence in favor of a flight-to-liquidity channel (i.e., sales of illiquid assets to buy liquid ones) is positive but also weak. If we take short-term Aaa corporate bonds and equities (because here we are dealing with the constituents of the STOXX 600) as instances of liquid assets, and Bbb corporate bonds as examples of illiquid securities (see Bolognesi et al., 2014), then we find evidence of a liquidity channel as the differential between the yields of liquid and illiquid assets. Indeed, over a long horizon, we find a spread of about 4-5 bps if we compare short-term Aaa corporate bond yields with short-term Bbb bond rates, and in excess of 5 bps between short-term Aaa bonds and long-term Bbb bonds. However, figure 3 shows that, as already pointed out by De Santis (2014) and others, the alleged European contagion was mostly driven by a flight–to–quality channel, by which, following a shock to a few low credit-quality government bond markets, investors attempt to sell risky assets and purchase safer assets. Consequently, the price of the former declines (yield climbs up), while the price of the latter increases (yield falls down). Indeed, as already noticed, while the yields of investment grade, Aaa corporate bonds and (at least eventually) core European government bonds decline, the yields of "junk", Bbb corporate bonds and the equity dividend yield increase. Although the effects of the flight-to-quality channel may partially overlap with the ones produced under the flight-to-liquidity channel, we find evidence of the former in the fact that also the spread between long-term (and thus generally less liquid) Aaa corporate bond yields and Bbb corporate bond rates increases as a consequence of the crisis shock. Indeed, if we compare the IRFs of long-term Aaa and Bbb bond yields we get an estimated difference that ranges between 2 and 4 bps per week, hence to be cumulated over time.

We also test the presence of a correlated information channel by measuring the non–linear and immediate effect captured by the *MSVAR* framework, vs. the short-term IRF estimated under a single–state *VAR* framework in Figure 1. A comparison of the two sets of IRFs shows that during the European GIIPS debt crisis, the correlated information channel was certainly at work, and that it explains an important portion of yield responses in the case of short-term Bbb corporate bonds and equities, in the sense that their medium-term IRFs increase in the non-linear case by 1-2 bps, which appears to represent between 20 and 40 per cent of the overall contagion effect that we have reported early on.[15] Interestingly, this effect also extends to the repo rate increase. This implies that a GIIPS shock would contain information useful to support an upward revision of short-term, virtually riskless repo rates.

Finally, under the risk premium channel, contagion occurs because shocks to one market lead to an increase in the overall risk aversion of financial market participants. This triggers an upward adjustment of the risk premia on all the risky assets in the economy. To assess the strength of this channel, we repeat the analysis that we performed for the yields for European yield spreads, obtained as the difference between the nominal yields and the overnight repo rate and hence compute the regime specific IRFs also for this model.[16] Figure 4 shows that in general we obtain weak evidence of a European risk premium channel, even in the crisis regime: the spreads of most series are predicted to decline as a result of the GIIPS sovereign yield shock. Rising yields (as implied by a flight-to-quality channel) and declining spreads are compatible when the underlying, baseline riskless rate climbs up, as it turns out to be case for the repo rate on German bonds, (see

[15] The correlated information is measured over a horizon of 6-12 weeks from the original shock because, by construction, our Cholesky ordering implies that a peripheral yield shock must imply very limited effects on other risky yields in the very short run.

[16] MSIH(3,1) estimates for the spread exercise are not reported to save space, but are available upon request from the Authors.

also Arghyrou and Kontonikas (2012) with reference to long-term Bunds). The only limited exception is represented by non-investment grade corporate bonds in the crisis regime, for which even though the immediate impact is a reduction of the spreads, over time we estimate an increase in spreads up to 2 and 3 bps (hence modest) for long- and short-term bonds, respectively. All in all, we conclude that the mild evidence of contagion from a positive (crisis) shock to peripheral European yields is mostly explained by a flight-to-quality channel being active during 2010 and 2011, when – whilst the yields of investment grade, Aaa corporate bonds and (at least eventually) core European government bonds declined – the yields of "junk", Bbb corporate bonds and on equities increased. Moreover, there is some evidence of flight–to–liquidity and correlated information channels although these account for at most 50% of the (already modest) size of the contagion effects.

Interestingly, the chances that the 2010-2011 European crisis was characterized by timid contagion outside the low quality, peripheral government bond market has received at best scant attention by the literature. Even though we may be tempted to interpret this evidence as an indication of success of the policies implemented by the ECB (see the discussion in Constancio, 2012), the possibility remains that other, more structural features of the European financial markets (for instance, a superior degree of segmentation) may have prevented more widespread and damaging contagion effects.

5. *Cross-Country, Cross-Market Shocks: Did the Subprime Crisis Spill Over to Europe?*

Our final empirical exercise further extends the eight-series estimation exercises of Section 4 to a richer data set composed of 10 series: the same eight series of Sections 3-4, augmented with two US ABS yield series (also published by BAML). The first series concerns Aaa–rated ABS and the second collects data on lower grade ABS that belong to the rating bracket AA–Bbb. We use this richer data set to investigate whether and how subprime shocks that have occurred between 2007 and 2008 may have spilled over to European markets. Our conjecture is that the type of shock may matter in determining how and whether any contagion occurs. Moreover, the question of whether US subprime-originated shocks may have caused a contagion affecting European markets is interesting in itself (see e.g., Acharya et al., 2014, and Ureche-Rangau and Burietz, 2014).

Similarly to the exercise performed for the European series only, we firstly estimate a VAR(2) model – selected by a specification similar to the one performed in Section 4.1 and available upon request from the Authors – that serves as a benchmark especially for the assessment of the correlated information channel. This selection also helps in performing comparisons with the results obtained in above. Of course, the inclusion of two additional yield series in the single-state VAR model raises the R-squares (even in adjusted terms) to levels that are even higher than those commented earlier. This occurs because lags of the two US ABS yield series have good forecasting power for most other yields under examination, in particular, European Bbb short-term corporate yields and the dividend yield. Moreover, the US ABS series are both rather predictable using the past history of European yields.[17]

An Appendix reports the IRFs obtained from this augmented model when low-grade, US Bbb-Aa ABS yield are hit by a one-standard deviation shock that raises yields. Therefore, these IRFs potentially track how a subprime-type shock in the US may have spilled over to European fixed income and equity markets. Also in this case the effects of international spillover to European markets are pervasively small in economic terms. Apparently, investors replace non-Aaa US ABS with European sovereign bonds, especially peripheral, high-yield ones, and European corporate paper. However, the effects are small and generally not statistically significant with the exception of GIIPS sovereign bonds, the yield of which declines over time by almost 20 bps with a response that turns statistically significant after 4 weeks. Clearly, this spillover effect represents the opposite of contagion: it seems that the bubble bursting in the US travels over to Europe fueling an increase in sovereign bond prices, especially peripheral ones, as if low grade US assets have been substituted in investors' portfolios with European peripheral sovereign bonds. Even though this is consistent with the empirical evidence of the years 2007-2009 included in our sample, the magnitude of the effect – a decline of 10-15 bps in peripheral sovereigns and a spread compression by roughly 10 bps – remains rather modest.

Second, we estimate a *MSIH(3,1)* model (with three regimes, one VAR lag, and a regime-invariant coefficient matrix) similar to the one presented in Section 4. An Appendix table with structure similar to Table 2 shows that two information criteria out of three, namely the H-QC and

[17] To save space, we omit a table, available upon request, with detailed coefficient estimates.

the SC, selects this model. Furthermore, this choice enhances comparability with the exercise performed in Section 4. Table 4 shows the estimates for the model. A vast majority of the estimated coefficients appears to be statistically significant, at least in tests with size of 10% or lower. Also in this case, one of the regimes may be branded as a low-volatility, low (average) correlations regime that characterizes periods of quiet stock and bond markets and of low yields (equity dividend yields included). Plotting the implied smoothed probabilities (see the Appendix) confirms this impression, in the sense that this regime characterizes almost without interruptions the sub-sample 2007-October 2009, essentially before the financial crisis takes hold of sovereign yields and spreads in Europe.

The two other regimes are a high volatility state with relatively low correlations and intermediate yields, and a crisis state with high volatility (a factor of 1.5 to 3 of the low volatility estimates), high correlations, and altered, high yields (especially US ABS, low quality corporate bonds, and PIIGS sovereign yields). These two regimes are less persistent and do communicate with each other, in the sense that it is relatively easy to "cycle" back and forth between regimes 2 and 3. However, and this is fully consistent with the historical record, the crisis regime mostly characterizes (almost 70% of the sub-sample can be classified as such) the period November 2009 – March 2012, with spikes in correspondence to the Spring/Summer of 2010 and then to the Fall of 2011, marking well-known bouts of European sovereign crisis.

Figure 5 investigates the existence and strength of contagion from US risky markets – specifically from low-grade ABS yields – to European markets in the aftermath of one standard deviation shock, used as a stylized way to capture the onset of the 2007 subprime crisis. Similarly to Figure 3, each panel in Figure 5 showcases response effects in each of the markets in the three different regimes. However, it is immediately clear that propagation effects in the first two regimes—i.e., low- and high- volatility, non-crisis regimes—are modest and almost never statistically significant. On the opposite, a more interesting story emerges with reference to the third, crisis state. In fact, in this state there is evidence of spillover effects from a US ABS shock to European markets, but *no evidence of contagion in an economic sense*. Indeed, similarly to what we find in the case of a single-state model, most European markets do react to US crisis shocks (and to a large scale, often to a statistically significant extent), but moving in an opposite direction, as if during a crisis regime, money would regularly flow out of "risk-on" US markets to a few European markets, and in particular the equity, the lower merit of credit corporate, and the government bond markets, including peripheral ones. Some have speculated that a market perception of an implicit European-wide bail-out guarantee, or simply ignorance among financial market participants of country-specific fundamentals, may have been the main explanations for this co-movement between core and GIIPS yields in 2007-2009 (see e.g., Beirne and Fratzscher, 2013) in the face of the US sub-prime crisis. In particular, while there are weaker spillovers of US ABS shocks to European Aaa corporate bonds and short-term cash (repo) investments, effects are large and precisely estimated in the case of other assets. Differently from Figure 3, now the effects are rather strong. For instance, in the case of GIIPS sovereign yields, these react slowly over time (up to 8-9 weeks the effect is not statistically significant), but the effect gradually ramps up over time reaching almost 250 bps within 6 months. Further evidence indicates that the effect levels off and starts being re-absorbed only over horizons that exceed the year. There is also an effect on core Europe sovereigns that is however weaker, in the order of 120-130 bps and significant only after 2-3 months from the shock. Of course, these effects may be perceived as rather large, especially when compared to those in Figure 3. However, here one has to consider that under the crisis regime in Table 4, a one-standard deviation shock to lower grade US ABS yields amounts to a hefty 33 bps per month, i.e., almost 4% in annualized terms. Of course, during the 2007-2009 subprime crisis, such a 4% annual increase in yields did occur and in hindsight almost seems negligible when compared to the shocks that actually took place.[18]

Interestingly, and reinforcing the effects already noted when a single-state VAR model had been estimated, the time value of essentially riskless overnight cash investments is significantly lowered by up to 40 bps. These represent clear flight-to-quality and flight-to-liquidity effects that occurs across different regions of the world, besides occurring across markets. Once more, these effects are the opposite of the standard notion of contagion, although these represent a case of spillover in a quantitative sense. It appears that one bubble bursting in the US may travel over to

[18] We have also performed the analysis assuming a more parsimonious (its saturation ratio exceeded 17), two-state MSIH(2,1) model finding results that are qualitatively similar but that were quantitatively smaller. This derives from the fact that a two-state model essentially groups the high-volatility and crisis states into a single regime that therefore is characterized by much smaller shocks to US ABS yield series.

Europe fueling an increase in sovereign bond prices, especially peripheral ones. In the case of a MSVAR framework, not only the sign of the effect helps us to make sense of widespread anecdotal evidence from the years 2007-2009, but delivers a reaction that is of a non-negligible magnitude: a decline of 20 bps in peripheral sovereign bonds.

6. Conclusions

In this paper we have studied whether and how European financial markets considered in their aggregate (i.e., without distinguishing among national markets) are subject to contagion effects from shocks originating in low-quality (e.g., peripheral sovereign bonds), low-liquidity (e.g., US Bbb ABS) markets. All in all, we find evidence that European financial markets would be more insulated from shocks – both of internal (Section 4) and of external origins (Section 5) – than US markets are, as reported by Longstaff (2010) and Guo et al., (2011). Such a result appears to be a new empirical finding so far unexplored in the literature and may have important policy implications. While it has been emphasized that financial markets have overreacted across the board during the crisis and that European sovereign risk was mispriced because of widespread contagion, especially for the GIIPS countries, our empirical evidence casts a few doubts on these often-heard claims. It seems that the real issue at stake may not been primarily contagion or mispricing but the structural, long-lasting imbalances that have characterized the Euro area instead and that most literature agrees upon in terms of representing the key drivers of the crises.

Of course, our analysis just allows us to measure the size of the phenomenon and to assess its statistical significance but is mute on the causes of this higher degree of insulation among different markets, whether these may be related to European policy-making or the very structure and segmentation features of markets. Future research will have to shed light on this phenomenon. It would be interesting to investigate whether these dissimilarities are caused by differences with the timing and way in which unconventional monetary policies were applied by the ECB vs. the Federal Reserve in the US (see, e.g., Borio and Disyatat, 2010). There is also a growing literature on contagion between government debt markets and banks (see, e.g., Acharya et al., 2014, Alter and Beyer, 2014, Alter and Schüler, 2012, Banerjee et al., 2016, Gorea and Radev, 2013, Mink and de Haan 2013). In July 2011 sovereign tensions spread not only to Italy and Spain, but also to banks exposed to the sovereign debt of these countries. The sovereign crisis has clearly affected funding availability and funding costs for individual banks in the euro area. It would be also interesting to extend our vector of variables under investigation to include either bank bond yields or bank CDS spreads to assess whether contagion did (also) concern these important elements of the cost of capital of the European banking systems.

Finally, the analysis in Section 5 has simply incorporated data on US asset-backed security yields. Even though there may be issues with over-parameterization, it would have been interesting to extend the analysis to include yield data from the European ABS market. Unfortunately, the European ABS market has basically shrunk to non-existence just after Lehman's demise in the Fall of 2008. Moreover, also the liquidity of the secondary market has plummeted: before the crisis, almost 70 percent of new issuance was placed on the market, and the remainder retained by originators; after 2008, the share of new issuance placed on the market dropped to below 10 percent, signaling virtual market refusal of these securities. In this environment, it may be hard to find European ABS yield series able to retain representativeness of ABS markets between 2008 and 2014, i.e., for roughly half of our sample.

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Table 1

Summary statistics for European bond and stock yields

Key summary statistics for weekly yield series over the sample period March 23, 2007 - December 19, 2014. The data are expressed in terms of annualized nominal yields. For instance, 1.00 stands for 1.00%. Jarque-Bera is a test statistic used to assess whether a series is normally distributed; asterisks denote statistical significance at conventional levels. EV stands for "equally weighted", NIG stands for "Non-Investment Grade", ST for "short term", and LT for "long term".

Panel A Summary statistics

	Mean	**Max.**	**Min.**	**Std. Dev.**	**Skewness**	**Kurtosis**	**Jarque-Bera**
Inv. Grade Corp. ST	2.647***	5.823	0.661	1.370	0.425*	1.972	30.014***
Inv. Grade Corp- LT	3.797***	6.280	1.569	1.106	-0.133	1.940	20.173***
NIG Corp. ST	4.639***	9.554	1.502	1.900	0.413	2.843	11.913***
NIG Corp. LT	5.392***	8.498	2.792	1.382	0.258	2.681	6.221**
Dividend Yield	3.769***	6.848	2.835	0.775	1.814***	6.183**	393.1***
Repo Rate (Bunds)	1.104**	4.362	-0.163	1.544	1.216**	2.721	101.2***
EV Core Sovereign	3.013***	4.834	0.765	1.024	-0.166	1.905	22.092***
EV PIIGS Sovereign	6.370***	15.254	2.915	2.958	1.416***	4.084*	155.2***
US ABS Aaa	1.964**	1.251	8.555	0.084	1.776***	1.981	38.000***
US ABS Aaa-Bbb	5.958***	3.930	20.98	2.206	4.929***	1.637*	1207.08***

*** = significant at a size of 1% or less; ** = significant at a size btw. 1 and 5%; * = significant at a size btw. 5 and 10%.

Panel B Correlations

	Inv. Grade ST	**Inv. Grade LT**	**NIG Corp. ST**	**NIG Corp. LT**	**Div. Yield**	**Repo Rate (Bunds)**	**EV Core Sovereign**	**EV PIIGS Sovereign**
Inv. Grade ST	1.000							
Inv. Grade LT	0.940***	1.000						
NIG Corp. ST	0.796***	0.853***	1.000					
NIG Corp. LT	0.791***	0.860***	0.989***	1.000				
Dividend Yield	0.490***	0.492***	0.769***	0.744***	1.000			
Repo Rate (Bunds)	0.883***	0.712***	0.526***	0.518***	0.320***	1.000		
EV Core Sovereign	0.953***	0.949***	0.789***	0.804***	0.412***	0.782***	1.000	
EV PIIGS Sovereign	-0.266***	-0.233***	0.053	0.063	0.005	-0.316***	-0.178**	1.000

*** = significant at a size of 1% or less; ** = significant at a size btw. 1 and 5%; * = significant at a size btw. 5 and 10%.

Table 2

Model selection results for Markov Switching models

This table reports the statistics used to select multivariate MSVAR models of the form:

$$y_t = \mu_{S_t} + \sum_{j=1}^{p} A_{j,S_t} y_{t-j} + \Omega_{S_t}^{1/2} e_t \qquad e_t \sim IID\ N(0, I_N).$$

The specification search is applied to weekly yield series over the sample period March 23, 2007 - December 19, 2014.

Model (k,p)	N. of parameters	Saturation Ratio	Log-likelihood	LR test for linearity	Akaike Criterion	Hannan-Quinn Criterion	Schwarz Criterion
Baseline model: Two-state, Markov Switching							
MSI(2,0)	54	60.000	-2245.21	866.9594 (0.000)	11.354	11.565	11.888
MSIA(2,1)	182	17.758	2486.39	397.0678 (0.000)	-11.408	-10.694	-9.605
MSIA(2,2)	310	10.400	2583.25	403.1568 (0.000)	-11.282	-10.064	-8.206
MSIH(2,0)	90	36.000	-1201.57	2954.237 (0.000)	6.378	6.730	7.268
MSIH(2,1)	154	20.987	2917.79	1259.881 (0.000)	-13.682	-13.078	**-12.157***
MSIH(2,2)	218	14.789	3005.37	1247.401 (0.000)	-13.833	-12.977	-11.670
MSIAH(2,1)	218	14.826	2908.25	1240.797 (0.000)	-13.318	-12.463	-11.159
MSIAH(2,2)	346	9.318	2923.37	1083.402 (0.000)	-12.791	-11.432	-9.358
Baseline model: Three-state, Markov Switching							
MSI(3,0)	66	49.091	-1700.06	1957.256 (0.000)	8.721	8.980	9.374
MSIA(3,1)	258	12.527	2662.21	748.7161 0.000	-11.902	-10.891	-9.347
MSIA(3,2)	450	7.164	2870.21	977.0681 (0.000)	-12.011	-10.243	-7.546
MSIH(3,0)	138	23.478	17.090	5391.557 (0.000)	0.597	1.137	1.961
MSIH(3,1)	202	16.000	3123.96	1512.223 (0.000)	**-14.269***	**-13.277***	**-12.068**
MSIH(3,2)	266	12.120	3136.38	1509.421 (0.000)	**-14.245**	**-13.200**	-11.606
MSIAH(3,1)	330	9.794	3101.77	1627.844 (0.000)	-13.722	-12.428	-10.453
MSIAH(3,2)	522	6.176	3333.46	1903.576 (0.000)	-13.953	-11.902	-8.773

* Model selected by the criterion stated in the header of the corresponding column.

Table 3

Estimates of a MSIH(3,1) model

This table reports the ML estimates of a VAR(2) model of the form:

$$y_t = \mu_{S_t} + \sum_{j=1}^{p} A_{j,S_t} y_{t-j} + \Omega_{S_t}^{1/2} e_t \qquad e_t \sim IID\ N(0, I_N) \quad S = 1, 2, 3.$$

Estimation is performed with reference to weekly bond and stock yield series over the sample period March 23, 2007 - December 19, 2014.

	Aaa Corp. Short	Aaa Corp. Long	Bbb Corp. Short	Bbb Corp. Long	STOXX 600 Div. Yield	Repo Rate	EV Core Country	EV PIIGS Country
1. Intercept terms								
Regime 1	**0.171****	**0.197****	**0.241*****	**0.279*****	**0.363*****	0.010	0.044	0.041
(Low volatility)	(0.012)	(0.017)	(0.002)	(0.000)	(0.000)	(0.749)	(0.540)	(0.713)
Regime 2	**0.151****	**0.203****	**0.385*****	**0.374*****	**0.494*****	0.058	0.039	0.083
(High volatility)	(0.086)	(0.042)	(0.000)	(0.000)	(0.000)	(0.248)	(0.665)	(0.534)
Regime 3	**0.179*****	**0.231*****	**0.251*****	**0.289*****	**0.370*****	-0.015	0.034	0.091
(Crisis)	(0.007)	(0.004)	(0.001)	(0.000)	(0.000)	(0.650)	(0.630)	(0.463)
2. VAR (1) Matrix								
AAA Corporate Short (t-1)	**0.890*****	-0.023	0.033	0.022	0.002	**0.036*****	**0.070****	0.025
	(0.000)	(0.479)	(0.299)	(0.472)	(0.950)	(0.002)	(0.016)	(0.525)
AAA Corporate Long (t-1)	0.000	**0.916*****	0.003	-0.005	-0.034	0.007	0.024	**0.075****
	(0.983)	(0.000)	(0.854)	(0.817)	(0.106)	(0.183)	(0.215)	(0.029)
BBB Corporate Short (t-1)	**0.034***	(0.021)	**1.022*****	**0.064*****	0.032	**-0.016*****	-0.016	0.001
	(0.059)	0.380	(0.000)	(0.004)	(0.143)	(0.007)	(0.449)	(0.974)
BBB Corporate Long (t-1)	-0.016	(0.001)	**-0.072*****	**0.885*****	-0.015	**0.016****	0.041	0.007
	(0.507)	0.980	(0.009)	(0.000)	(0.618)	(0.047)	(0.128)	(0.869)
STOXX 600 Dividend Yield (t-1)	**-0.044****	**-0.031***	**-0.034***	-0.023	**0.904*****	**-0.021****	**-0.039****	**-0.051*****
	(0.018)	(0.085)	(0.100)	(0.194)	(0.000)	(0.031)	(0.014)	(0.006)
Repo Rate (German Bunds) (t-1)	**0.059*****	0.022	-0.016	-0.012	-0.004	**0.969*****	-0.011	-0.018
	(0.000)	(0.200)	(0.439)	(0.507)	(0.837)	(0.000)	(0.471)	(0.356)
EV Core Yields (t-1)	**0.056****	**0.069****	-0.001	0.010	0.000	**-0.023****	**0.893*****	**-0.075****
	(0.021)	(0.012)	(0.970)	(0.709)	(0.991)	(0.019)	(0.000)	(0.025)
EV PIIGS Yields (t-1)	**-0.004**	**-0.008**	**0.007****	**0.005***	0.003	**0.004*****	0.001	**0.996*****
	(0.079)	(0.005)	(0.016)	(0.059)	(0.310)	(0.000)	(0.697)	(0.000)
3. Unconditional mean	2.647	3.797	4.639	5.392	3.769	1.104	3.013	6.370

*** significant at 1% level, ** significant at 5% level, * significant at 10% level.

Table 3 (continued)

Estimates of a MSIH(3,1) model

	Aaa Corp. Short	**Aaa Corp. Long**	**Bbb Corp. Short**	**Bbb Corp. Long**	**STOXX 600 Div. Yield**	**Repo Rate**	**EV Core Country**	**EV PIIGS Country**
4. Correlations/ Volatilities								
Regime 1								
AAA Corp. Short	**0.064*****							
AAA Corp. Long	**0.417*****	**0.099*****						
BBB Corp. Short	**0.549*****	**0.487*****	**0.063*****					
BBB Corp. Long	**0.559*****	**0.615*****	**0.748*****	**0.071*****				
STOXX Div. Yield	0.006	0.084	0.070	**0.118***	**0.073*****			
Repo Rate	**0.264****	0.012	**0.169****	**0.154***	-0.057	**0.015*****		
EV Core Yields	**0.269****	**0.325*****	**0.257****	**0.326*****	**-0.171***	0.066	**0.075*****	
EV PIIGS Yields	0.100	0.096	**0.203****	**0.184****	**0.196****	**0.139***	-0.014	**0.194*****
Regime 2								
AAA Corp. Short	**0.183*****							
AAA Corp. Long	**0.774*****	**0.154***						
BBB Corp. Short	**0.612*****	**0.635*****	**0.182*****					
BBB Corp. Long	**0.514*****	**0.654*****	**0.812*****	**0.142*****				
STOXX Div. Yield	**0.309****	**0.338*****	**0.544*****	**0.411*****	**0.203*****			
Repo Rate	**-0.228****	**-0.358*****	**-0.255*****	**-0.289*****	**-0.178****	**0.164*****		
EV Core Yields	**0.278*****	**0.183****	-0.033	0.057	**-0.372*****	0.097	**0.121*****	
EV PIIGS Yields	**0.358*****	**0.308*****	**0.131***	**0.175***	**0.142***	-0.037	**0.898*****	**0.133*****
Regime 3								
AAA Corp. Short	**0.079*****							
AAA Corp. Long	**0.853*****	**0.104*****						
BBB Corp. Short	**0.368*****	**0.250****	**0.153*****					
BBB Corp. Long	**0.561*****	**0.589*****	**0.632*****	**0.130*****				
STOXX Div. Yield	**0.132***	**0.142***	**0.281****	**0.258****	**0.107*****			
Repo Rate	0.067	-0.012	-0.030	-0.031	**-0.236****	**0.177*****		
EV Core Yields	**0.376*****	**0.475*****	0.060	**0.210****	0.071	0.070	**0.117*****	
EV PIIGS Yields	-0.039	-0.013	**0.215****	**0.156***	**0.172****	0.062	**-0.104***	**0.603*****

5. Transition Matrix	Regime 1	Regime 2	Regime 3
Regime 1 (Low volatility)	**0.833*****	0.000	0.167
Regime 2 (High volatility)	0.000	**0.964*****	0.036
Regime 3 (Crisis)	0.208	0.026	**0.766*****

*** significant at 1% level, * significant at 5% level, * significant at 10% level.

Table 4

Estimates of a MSIH(3,1) model that includes US ABS series

This table reports the ML estimates of a VAR(2) model of the form:

$$y_t = \mu_{S_t} + \sum_{j=1}^{p} A_{j,S_t} y_{t-j} + \Omega_{S_t}^{1/2} e_t \qquad e_t \sim IID\ N(0, I_N) \quad S = 1, 2, 3.$$

Estimation is performed with reference to weekly bond and stock yield series over the sample period March 23, 2007 - December 19, 2014.

	Aaa Corp. Short	Aaa Corp. Long	Bbb Corp. Short	Bbb Corp. Long	STOXX 600 DY	Repo Rate	EV Core	EV PIIGS	US AAA ABS	US BBB ABS
1. Intercept terms										
Regime 1	**0.180*****	0.114	**0.163*****	**0.239*****	**0.330*****	**-0.036****	-0.0620	-0.083	0.041	**-0.245*****
(Low volatility)	(0.004)	(0.251)	(0.009)	(0.001)	(0.000)	(0.085)	(0.370)	(0.665)	(0.215)	(0.000)
Regime 2	**0.184****	0.139	0.174	**0.245***	**0.334*****	-0.049	-0.076	-0.019	0.051	**-0.246*****
(High volatility)	(0.041)	(0.201)	(0.270)	(0.070)	(0.002)	(0.772)	(0.534)	(0.976)	(0.361)	(0.003)
Regime 3	0.200	0.072	0.123	**0.252***	**0.303***	-0.010	-0.015	-0.205	0.048	**-0.545***
(Crisis)	(0.230)	(0.616)	(0.467)	(0.059)	(0.097)	(0.955)	(0.896)	(0.118)	(0.815)	(0.095)
2. VAR (1) Matrix										
AAA Corporate Short (t-1)	**0.853*****	0.009	0.075	-0.043	-0.020	**0.089*****	**0.068****	-0.014	**-0.043*****	**0.001****
	(0.000)	(0.479)	(0.299)	(0.472)	(0.950)	(0.002)	(0.016)	(0.525)	(0.005)	(0.025)
AAA Corporate Long (t-1)	-0.029	**0.911*****	0.028	0.034	-0.010	0.053	0.053	**-0.018****	**-0.031***	0.000
	(0.983)	(0.000)	(0.854)	(0.817)	(0.106)	(0.183)	(0.215)	(0.029)	(0.088)	(0.857)
BBB Corporate Short (t-1)	**-0.010***	0.010	**1.040*****	**-0.060*****	-0.013	**0.038*****	-0.009	-0.001	-0.001	0.001
	(0.059)	(0.380)	(0.000)	(0.004)	(0.143)	(0.007)	(0.449)	(0.974)	(0.820)	(0.478)
BBB Corporate Long (t-1)	-0.007	0.008	**0.084*****	**0.883*****	-0.007	**0.020****	-0.001	0.000	-0.007	-0.000
	(0.507)	(0.980)	(0.009)	(0.000)	(0.618)	(0.047)	(0.128)	(0.869)	(0.234)	(0.954)
STOXX 600 Dividend Yield (t-1)	**0.003****	**-0.035***	**0.024***	-0.002	**0.903*****	**0.003****	**-0.017****	**0.004*****	**0.039****	**-0.001***
	(0.018)	(0.085)	(0.100)	(0.194)	(0.000)	(0.031)	(0.014)	(0.006)	(0.011)	(0.083)
Repo Rate (German Bunds) (t-1)	**0.024*****	0.008	-0.002	0.008	0.004	**0.983*****	-0.013	-0.001	-0.005	**-0.008***
	(0.000)	(0.200)	(0.439)	(0.507)	(0.837)	(0.000)	(0.471)	(0.356)	(0.248)	(0.093)
EV Core Yields (t-1)	**0.043****	**0.031****	0.022	0.063	-0.003	**0.007****	**0.883*****	**-0.014****	**-0.028****	**-0.018*****
	(0.021)	(0.012)	(0.970)	(0.709)	(0.991)	(0.019)	(0.000)	(0.025)	(0.011)	(0.008)
EV PIIGS Yields (t-1)	**-0.021****	**0.061****	0.081	0.052	-0.012	**0.015****	**-0.040*****	**0.948****	**-0.025*****	**-0.030*****
	(0.021)	(0.012)	(0.970)	(0.709)	(0.991)	(0.019)	(0.000)	(0.025)	(0.008)	(0.004)
US AAA ABS Yields (t-1)	**0.018****	**-0.003****	0.005	0.013	-0.002	**0.016****	**-0.032*****	**-0.004****	**0.987*****	-0.006
	(0.021)	(0.012)	(0.970)	(0.709)	(0.991)	(0.019)	(0.000)	(0.025)	(0.000)	(0.143)
US BBB ABS Yields (t-1)	**0.087***	**0.012*****	**-0.031****	**0.055***	0.058	**0.012*****	-0.121	**-0.003*****	**0.119*****	**0.962*****
	(0.079)	(0.005)	(0.016)	(0.059)	(0.310)	(0.000)	(0.697)	(0.000)	(0.000)	(0.000)
3. Unconditional mean	2.749	3.829	4.603	5.370	3.588	1.195	3.168	7.991	1.837	4.421

*** significant at 1% level, ** significant at 5% level, * significant at 10% level.

Table 4 (continued)

Estimates of a MSIH(3,1) model that includes US ABS series

	Aaa Corp. Short	Aaa Corp. Long	Bbb Corp. Short	Bbb Corp. Long	STOXX 600 DY	Repo Rate	EV Core	EV PIIGS	US AAA ABS	US BBB ABS
4.Correlations / Volatilities										
Regime 1										
AAA Corp. Sh	0.063***									
AAA Corp. Lng	0.401***	0.099***								
BBB Corp. Sh	0.560***	0.502***	0.063***							
BBB Corp. Lng	0.573***	0.617***	0.770***	0.071***						
STOXX DY	-0.008	0.099*	0.020	0.161*	0.073***					
Repo Rate	0.130*	-0.050	0.049	0.086*	0.039	0.021***				
EV Core Yields	0.266**	0.256**	0.289**	0.298***	-0.171*	0.075	0.069***			
EV PIIGS Yields	0.530***	0.048	0.135*	0.135*	0.275**	0.132*	-0.100*	0.191***		
US AAA ABS	0.107*	0.283**	0.124*	0.248**	-0.111*	-0.101*	0.460***	-0.081	0.033***	
US BBBABS	0.088*	0.282**	0.168*	0.265**	-0.017	-0.064	0.377***	-0.083*	0.890***	0.052***
Regime 2										
AAA Corp. Sh	0.090***									
AAA Corp. Lng	0.804***	0.108***								
BBB Corp. Sh	0.219**	0.214**	0.158***							
BBB Corp. Lng	0.433***	0.565***	0.662***	0.135***						
STOXX DY	0.015	0.067	0.296**	0.268**	0.110***					
Repo Rate	0.039	0.026	-0.041	-0.029	-0.235**	0.167***				
EV Core Yields	0.382***	0.475***	0.007	0.155*	0.043	0.059	0.123***			
EV PIIGS Yields	-0.080	-0.046	0.183*	0.140*	0.155*	0.045	-0.161*	0.620***		
US AAA ABS	0.306***	0.254**	0.258**	0.094*	0.054	-0.057	0.345***	-0.010	0.055***	
US BBBABS	0.310***	0.246***	0.170*	0.159*	0.008	0.016	0.184*	0.008	0.625***	0.082***
Regime 3										
AAA Corp. Sh	0.167***									
AAA Corp. Lng	0.787***	0.144***								
BBB Corp. Sh	0.648***	0.594***	0.170***							
BBB Corp. Lng	0.538***	0.639***	0.781***	0.133***						
STOXX DY	0.364***	0.356**	0.534***	0.394***	0.182***					
Repo Rate	-0.192*	-0.314**	-0.123	-0.183*	-0.117	0.169***				
EV Core Yields	0.276**	0.248**	0.034	0.143*	0.029	0.081	0.115***			
EV PIIGS Yields	0.383***	0.349**	0.193*	0.231**	0.203*	-0.050	0.887***	0.131***		
US AAA ABS	0.260**	0.200*	0.459***	0.458***	0.180*	0.041	0.120*	0.116	0.205***	
US BBBABS	-0.038	-0.002	0.130	0.027	0.117	0.025	-0.032	-0.056	0.335**	0.327***

5. Transition Matrix	Regime 1	Regime 2	Regime 3
Regime 1	0.993***	0.017	0.000
Regime 2	0.010	0.770***	0.210***
Regime 3	0.000	0.174**	0.993***

*** significant at 1% level, * significant at 5% level, * significant at 10% level.

Figure 1

VAR- Impulse response functions to a shock to peripheral (GIIPS) sovereign yields

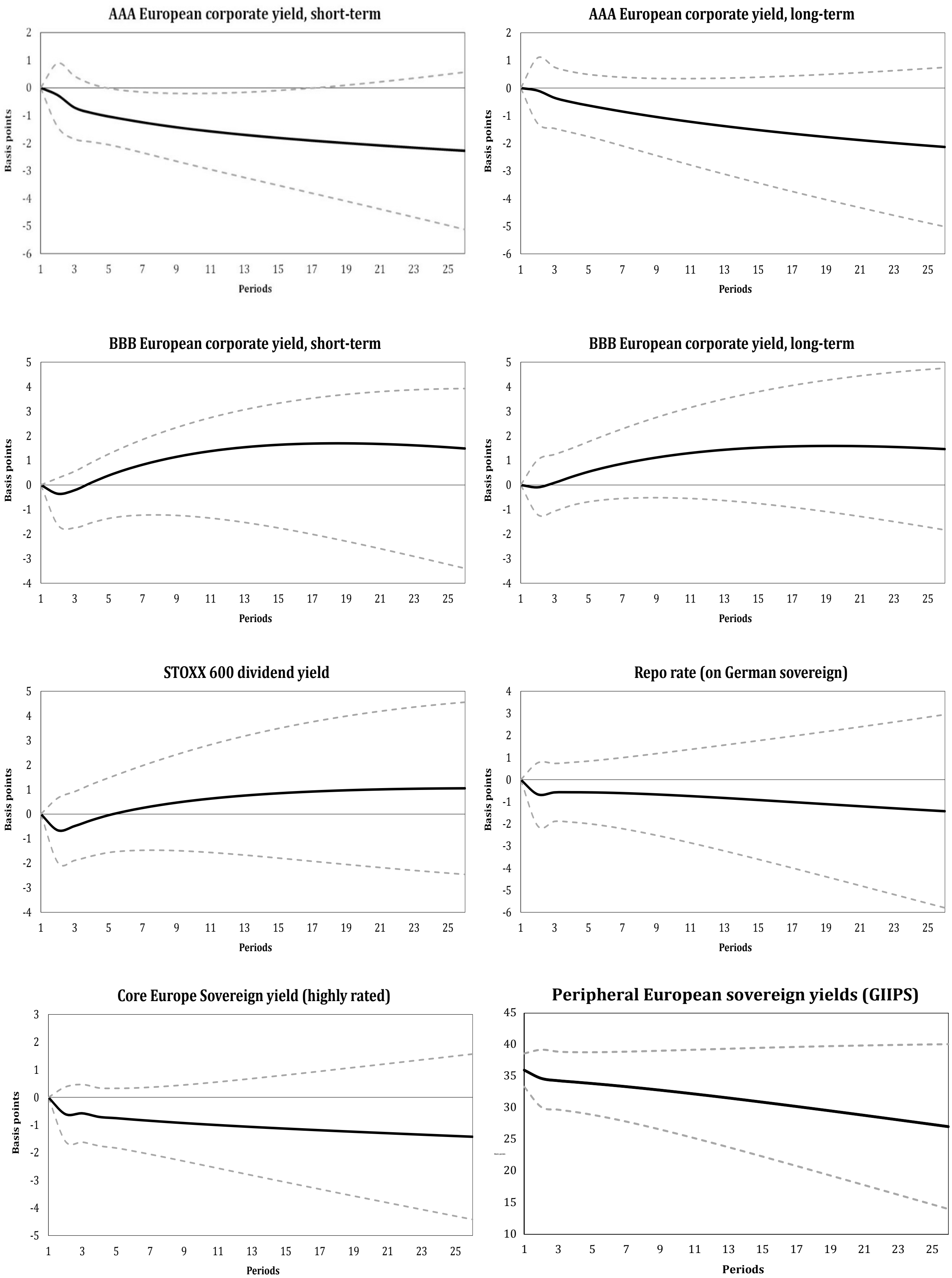

Figure 2

Smoothed probabilities estimated from a MSIH(3,1) model

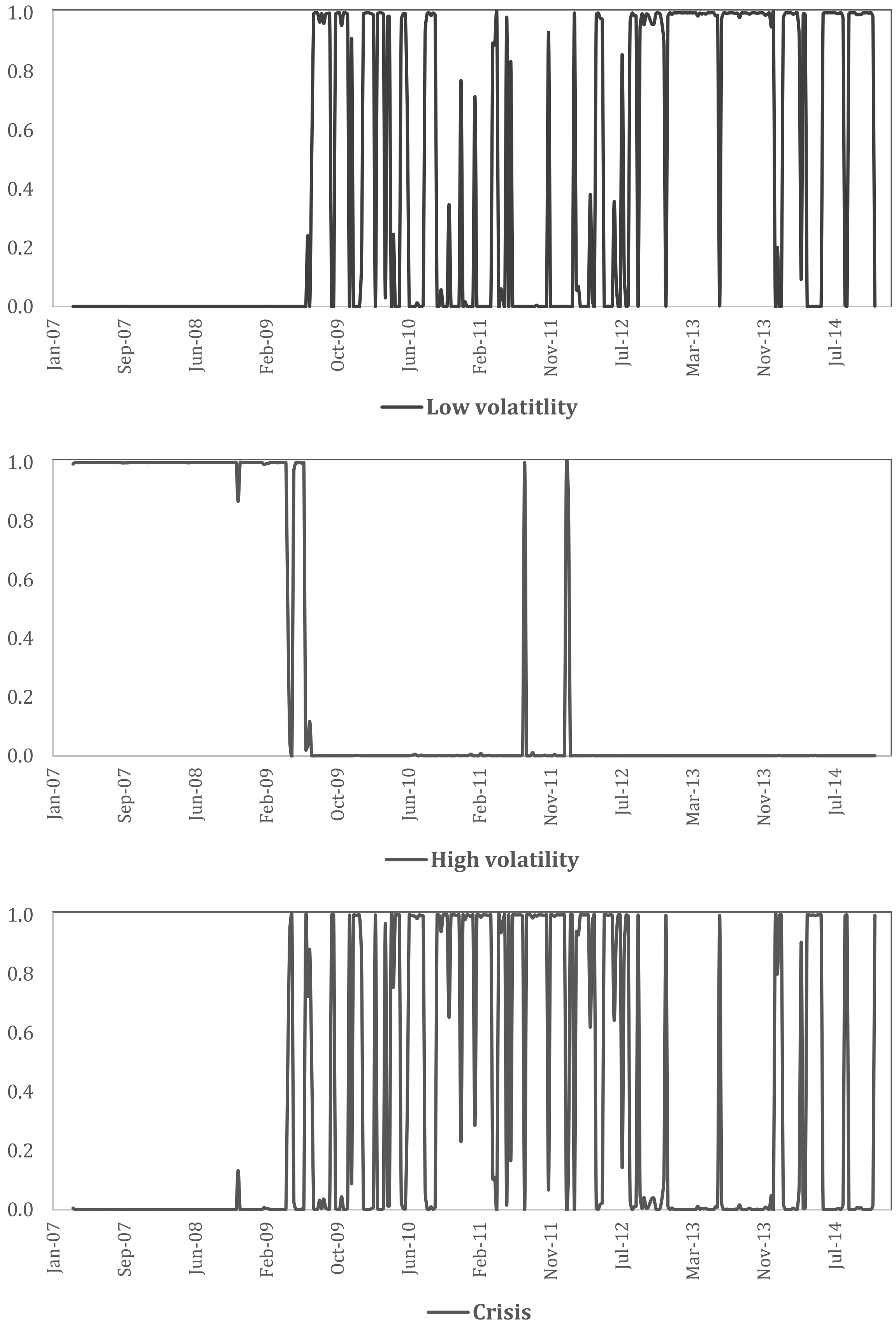

Figure 3

MSVAR- Impulse response functions to a shock to peripheral (GIIPS) sovereign yields

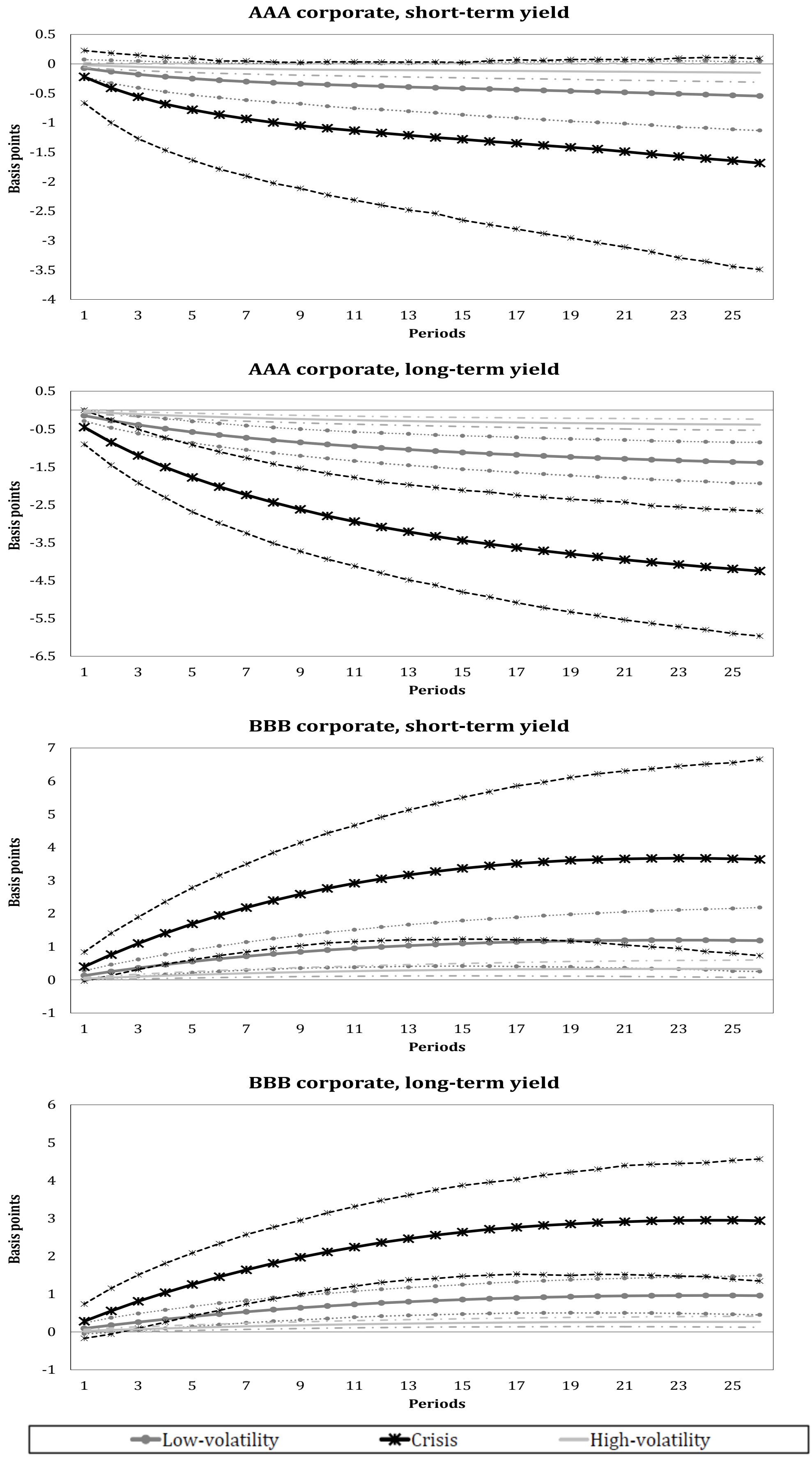

Figure 3 (continued)

MSVAR- Impulse response functions to a shock to peripheral (GIIPS) sovereign yields

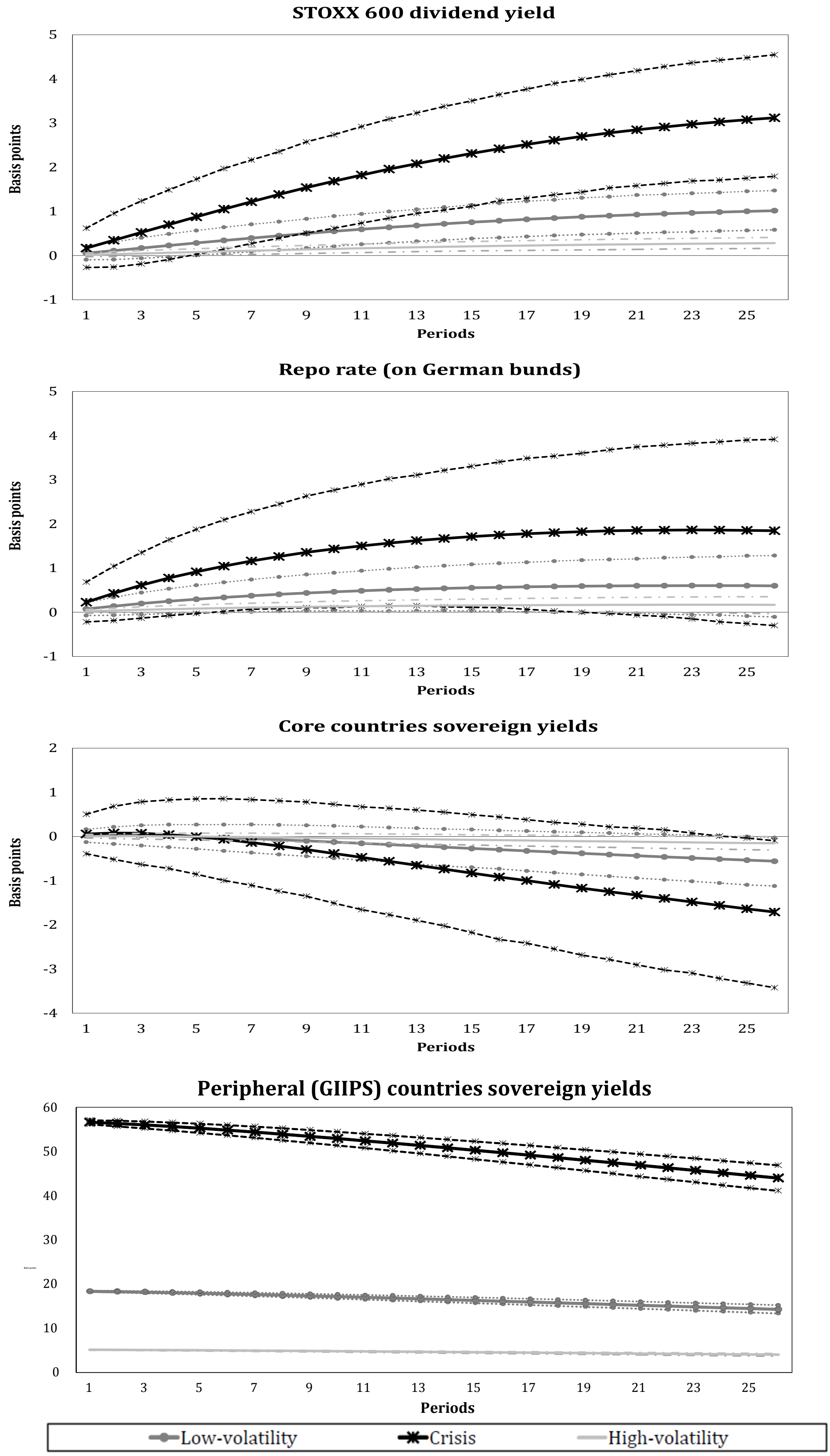

Figure 4

***MSVAR* - Cumulative impulse response functions of spreads to a shock to peripheral (GIIPS) sovereign yield spreads**

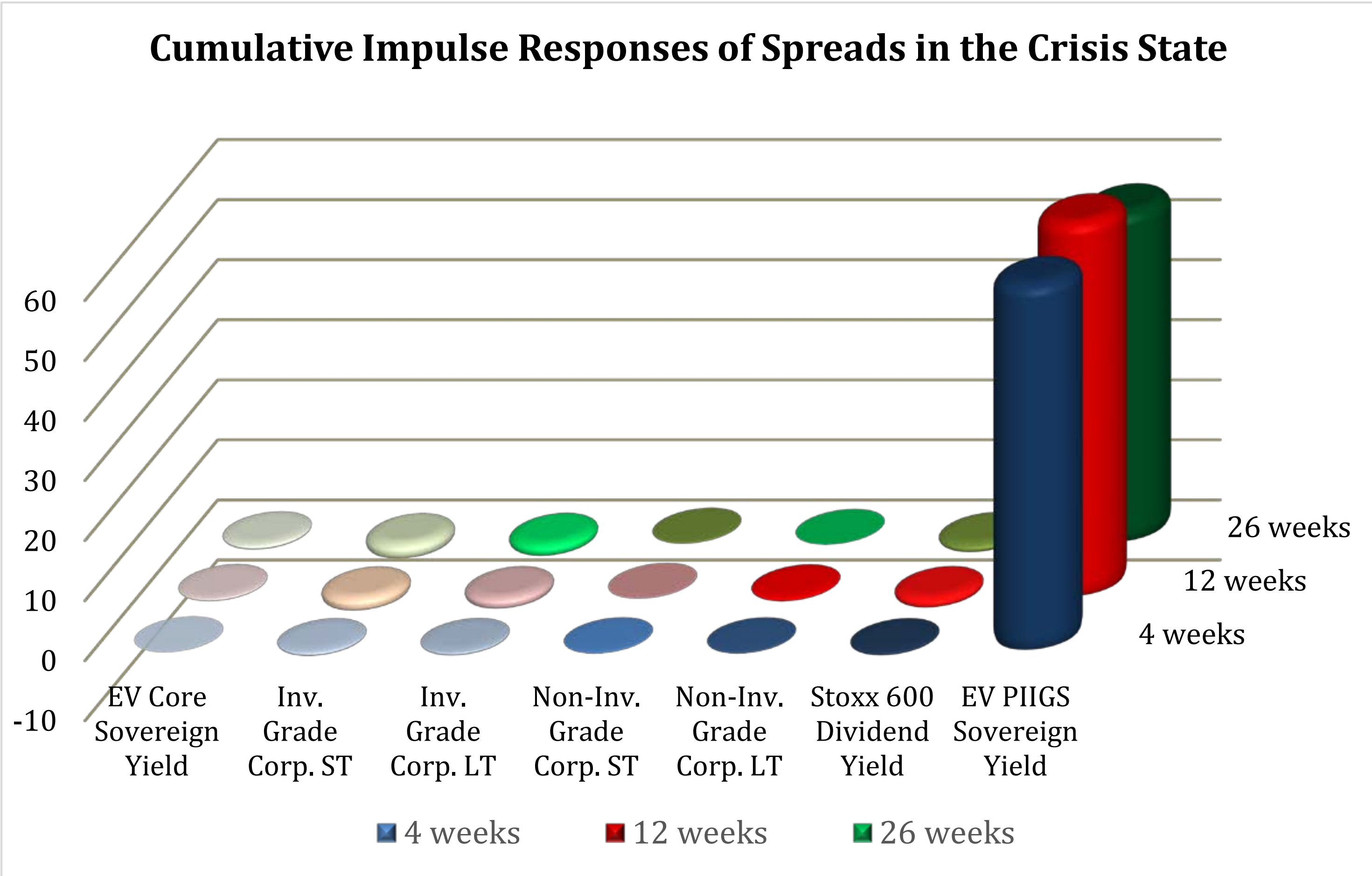

Figure 5

MSVAR- Impulse response functions to a shock to US low-credit quality ABS yields

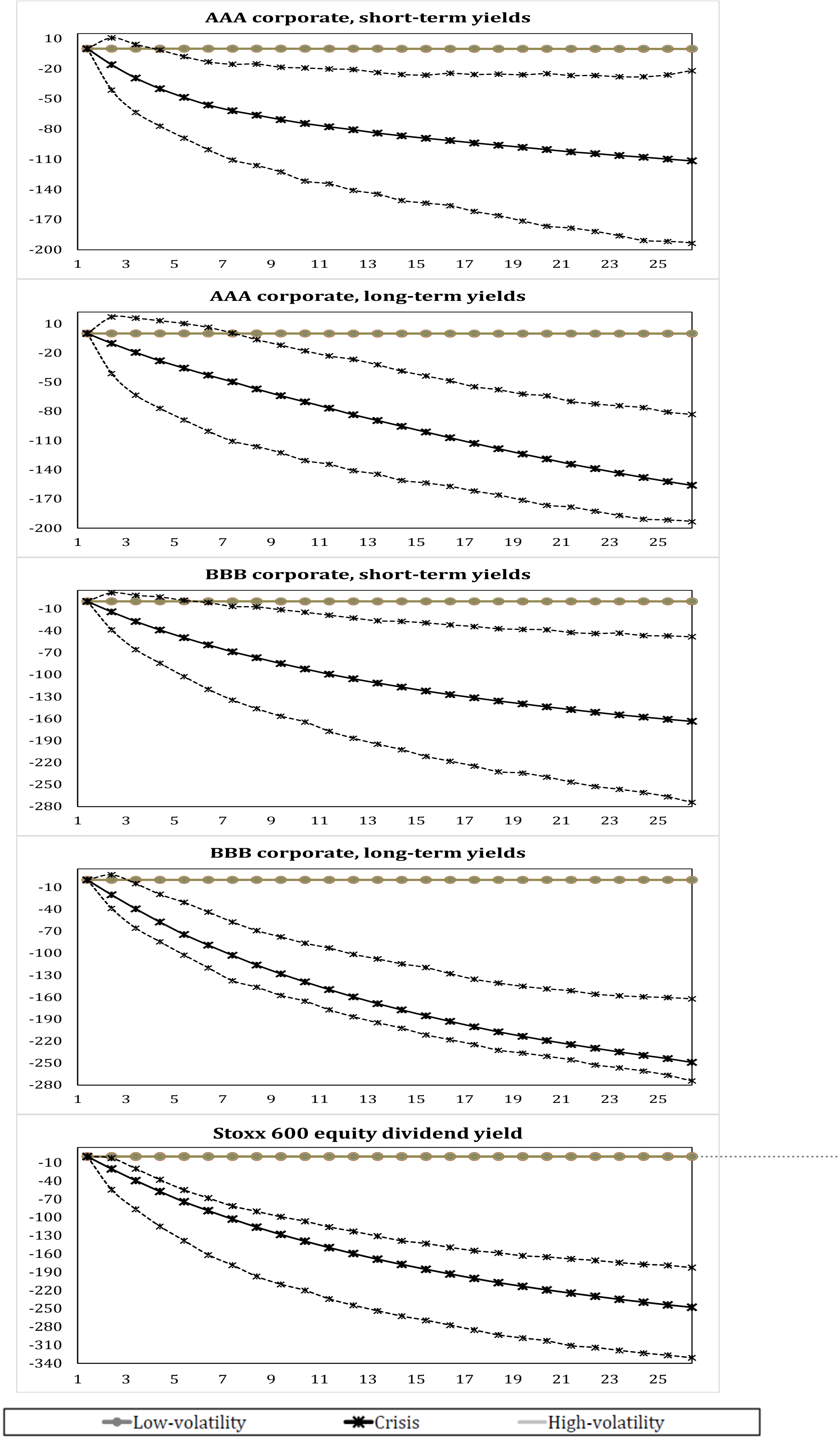

Figure 5 (continued)

MSVAR- Impulse response functions to a shock to US low-credit quality ABS yields