Sovereign Spreads in the Euro Area: Cross Border Transmission and Macroeconomic Implications

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Abstract

I use high frequency market reactions to foreign events to identify shocks to sovereign spreads, orthogonal to the economy, during the euro crisis. I show two things. First, that an increase in sovereign spreads has a contractionary macroeconomic impact with transmission running through a deterioration in private financial conditions. Second, that market reactions to foreign events explained a meaningful share of the variation in a sovereign's cost of borrowing during the crisis. To generate these results, I build a new narrative dataset of country specific events in the crisis period that are timed to a high frequency. (JEL Codes: E44, E65, F42)

Key words: Sovereign Risk Passthrough, High Frequency Identification, Contagion.

1 Introduction

At 13:01 London time on August 28th 2012 the *Reuters* reported that the Spanish region of Catalonia would request 5 billion euros of aid from Spain's Regional Liquidity Fund.[1] The move had been signaled by the Catalan authorities but the announcement still provoked a reaction in financial markets. Just before the announcement Spanish 2 year bonds yielded 3.56%

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[1] A live summary of events during the day can be found here: http://www.telegraph.co.uk/finance/debt-crisis-live/9502734/Debt-crisis-as-it-happened-August-28-2012.html

more than the German equivalent; by 2pm this figure had risen to 3.61%. The response was matched at other maturities: the 10 year bond spread rose from 5.02% to 5.10%, a 3 standard deviation move at this frequency over the crisis period. The bailout decision was largely a domestic policy matter. It represented a transfer of liabilities from Catalonia to Madrid. It had no direct international aspect. Nonetheless, the move in yields was not confined to Spain: Italian 2-year and 10-year bond yields increased by around 5 basis points immediately after the announcement. Yields in "core" countries were stable: German and French 2 year yields had little discernible change (see Figure 1).

This example illustrates two features of sovereign borrowing costs during the recent crisis in the Euro Area. First, the prices of Euro Area sovereign bonds reacted strongly to specific events and, second, those events transmitted across borders to other crisis-hit countries.[2] This cross-border transmission is important. The Catalan authorities were not obviously acting specifically in response to a macroeconomic shock in Italy in August 2012. However, even if there was a systematic Catalan reaction to events in Italy, the high frequency market response should isolate a surprise component of the decision, orthogonal to existing information (Gurkaynak and Wright (2013)). The Italian market move therefore represents a change in sovereign borrowing costs that is arguably orthogonal to other contemporaneous shocks hitting the Italian economy.

I use this line of reasoning to consider two related questions: One, is there a portion of a sovereign's borrowing costs driven by factors unrelated to the local economy, and if so how important is it? And, two, given such orthogonal shocks exist, what are the macroeconomic consequences of shocks to a sovereign's costs of borrowing? These questions played a crucial role in policy debates during the crisis, yet the existing empirical literature struggles to provide conclusive answers.

The key challenge is simultaneity. A rise in sovereign spreads may reflect economic weakness that leads to worsening primary balances and a rise in public debt. At the same time higher borrowing costs can exacerbate fiscal distortions and tighten credit conditions, curbing

[2]To highlight a few examples from this literature documenting these facts: Gade et al. (2013) show that statements by European politicians (both nationally and at a EU level) have had meaningful impact on sovereign borrowing costs. Brutti and Saure (2015) show that during the early stages of the crisis in Greece (2009-2011), critical events related to that country passed through to CDS spreads in the remainder of the EU. Attinasi et al. (2009) and Acharya et al. (2014) conduct similar analysis for bank bailout decisions. In the broadest case, Beetsma et al. (2013) show that "news" in general (as isolated from a news summary) as opposed to specific events also move markets throughout the union.

output and weakening fiscal positions. Disentangling these mechanisms is not possible with aggregate time series alone. This paper tackles this identification problem.

I proceed in three steps, making three distinct contributions to the literature. First, I build a new narrative dataset of the crisis period allowing me to isolate bond market movements due to foreign events. I use a news summary to isolate key, country-specific, political events in economies suffering from elevated sovereign borrowing costs during the crisis period. I determine, using a news-wire, the time at which an event occurred. I then measure the impact on sovereign spreads in other crisis-hit countries by looking at the response of the relevant sovereign bond market in an immediate time window spanning an announcement. In doing so, I contribute to existing narrative/news based studies of the crisis in Europe (Beetsma et al. (2013), Brutti and Saure (2015)) by providing a narrative of precisely timed events that enable the measurement of high frequency market reactions. These reactions enable events to be signed and ranked for empirical precision and enable the isolation of a surprise component for identification.

Second, using this dataset, I show that foreign events were an important determinant of sovereign borrowing costs during the crisis period. The high frequency market reactions in narrow event windows explain a disproportionate share of the daily variation in spreads on the days when the events occur. Furthermore, the accumulated reactions to foreign events explain around 20% of the monthly variation in 2 year sovereign spreads pooled regressions covering Italy, Portugal and Spain (the contribution is much weaker for Ireland). This figure is higher still when I use these accumulated reactions as an external instrument in a VAR model, addressing attenuation through measurement error, and compute the relevant forecast error variance decomposition.

This finding has implications for the literature on the determinants of sovereign borrowing costs. In the canonical limited commitment, incomplete markets model of sovereign debt and default (Eaton and Gersovitz (1981)), the price of debt is pinned down by the state of the economy and the debt stock. My claim is that the foreign events I study are informative about neither. Instead, if one interprets foreign events as transmitting due to shifts in investor beliefs, my results are instead consistent with models where sovereign spreads are partly determined by non-fundamental factors due to multiplicity (Calvo (1988); Cole and Kehoe (2000)). Other alternative interpretations are that foreign events serve as a signal over a political willingness to repay debt, consistent with the hidden information model in Aguiar and Amador (2014, section 4) or transmit through common creditors (Arellano et al. (2017)).[3]

While my analysis is somewhat silent on the exact source of contagion between countries, I do document the following: (i) Foreign events have a persistent impact on borrowing costs, suggesting that the reactions are not a function of temporary frictions in bond markets. (ii) That the relative market reactions to events seem unrelated to trade or financial linkages between countries; suggesting the reactions are not the result of a direct external macroeconomic shock. (iii) The events do not move a simple proxy for redenomination risk (from De Santis (2015)); suggesting that the events were not transmitting through concerns about the sustainability of the single currency.

Third, I provide empirical evidence on the macroeconomic consequences of a shock to sovereign spreads. As mentioned, I use a lower frequency VAR model containing aggregate time series with the high frequency market reactions entering as an external instrument (Gertler and Karadi (2015)). The time period is restricted mostly to the crisis. The sample is short, covering 6 years. To address this, I use the cross-section of countries to enhance the precision of the estimates and estimate a partially pooled panel VAR model using Bayesian methods (Jarocinski (2010)). I find that innovations to sovereign spreads were a critical driver of unemployment dynamics in the crisis period, explaining over 20% of the variation of the unemployment rate. In terms of relative magnitudes: a shock that leads to a 100bps increase in the 2 year sovereign yield corresponds to a percentage point reduction in industrial production growth and adds 0.4 percentage points to the unemployment rate. These results confirm the contractionary effect of an increase in sovereign spreads, even only in anticipation of a potential default; a feature sometimes missing from workhorse quantitative models of sovereign debt (Aguiar and Gopinath (2006), Arellano (2008)).

I also shed light on the channels through which sovereign spreads affect the real economy. I show that higher spreads are consistent with a general tightening in the financial conditions faced by the private sector, corporate bond yields rise and equity values of both banks and non-financial fall. Less conclusively, I also find that higher sovereign spreads cause a decline in

[3] There is also a substantial empirical literature looking at the determinants of yields during the crisis. Ang and Longstaff (2013) Aizenman et al. (2013), De Grauwe and Ji (2013) and Giordano et al. (2013) all investigate how sovereign borrowing costs depend on macroeconomic conditions, either across countries or over time. This paper is complementary in that it uses high frequency reactions to foreign events to isolate movements in borrowing costs unrelated to local macroeconomic conditions.

credit volumes and capital flight. These findings corroborate the theoretical literature focusing on the macroeconomic implications of the sovereign spreads via their effect on private financial conditions. Bocola (2016) shows how sovereign risk can pass through to the credit conditions faced by firms through the balance sheets of financial intermediaries. Neumeyer and Perri (2005), Uribe and Yue (2006) argue that this pass through means that fluctuations in sovereign spreads help explain business cycles in emerging markets, while Corsetti et al. (2013) argues that this has implications for the strength of the fiscal multiplier.

In the same regard, a number of papers (Acharya et al. (2017), Bofondi et al. (2018), De Marco (forthcoming)) using micro data in a difference-in-differences set up, have documented that banks that were more exposed to risky sovereigns raised the price of their credit or reduced its supply. In turn this had contractionary effects on firms with relationships with those banks. My work is complementary in the sense that empirical analysis using aggregate time series data is able to show the macroeconomic consequences of a shock accounting for potential general equilibrium and substitution effects. Moreover, I condition on a shock to sovereign spreads that is plausibly exogenous to economic conditions.

Next, I review a couple more anecdotal examples of the international transmission of events during the crisis in the euro area, discuss the validity and strengths of my identification strategy, and then describe how I construct my dataset. Section 3 shows that intraday market reactions to foreign events explain a reasonable portion of the lower frequency variation in sovereign borrowing costs and Section 4 discusses the macroeconomic impact.

2 Market Reactions To Foreign Events in the Euro Crisis

2.1 More examples

The Greek Parliament Approves Austerity, 12th of February 2012 By early 2012 it was clear that Greece would need a second bailout. An agreement in principle was reached between Greece and the Troika of official creditors. In exchange for official sector financing, Greece was required to embark on additional austerity measures, including 150,000 public sector job losses and 3.3 billion euros of spending cuts. A deal was approved by the Greek Cabinet on February 10th but led to the resignation of 6 Cabinet members. Late in the evening on Sunday the 12th the package would be voted on in the Greek parliament. The leadership of the two main Greek political parties backed the agreement. However, political support was not unanimous. A small nationalist party withdrew its support from the governing coalition and the days leading up to the vote were marked by social unrest.[4]

A no vote could have resulted in disorderly default. Greece had a 14.5bn euro bond payment scheduled on March 20th; a payment that, in the absence of a bailout, it appeared the country would be unable to meet. Official creditors refused to sanction a bailout unless austerity measures were approved. Ultimately the legislation passed comfortably, by 199 votes in favour to 74 against, out of 300 lawmakers. The news was not all positive: Athens suffered from rioting; 40 government MPs rebelled against the measure leading to them being ousted from their parties and cutting the ruling coalition's majority. Nonetheless, financial markets reacted positively to the vote. Sovereign bond yields fell sharply in Greece when the market opened on the 13th and the response spread throughout the union. By 8:30am London time, spreads on benchmark 2 year sovereign bonds had fallen by 31.5bp in Portugal, 9.8bp in Italy and 3.0bp in Spain when compared to the close on the 10th.

Indecisive Election in Italy, 25th February 2013 On the 8th of December 2012, technocratic Italian Prime Minister Mario Monti announced that he would resign following the withdrawal of his predecessor's, Silvio Berlusconi, endorsement of his government. A general election was scheduled for the 24th-25th of February. The political situation was finely poised. Ahead of the vote, the electorate seemed split four ways between a centre-left coalition, a centre-right coalition headed by Berlusconi, a centrist coalition headed by Monti and the anti-establishment Movement 5 Star. To govern, a coalition would need to win both houses of parliament. Italian electoral law heavily favoured the coalition with most votes in assigning seats in the lower house but the upper house, the Senate, allocated seats on a regional level making the outcome unpredictable. A minority government or a hung parliament was seen as increasing economic uncertainty, reducing the likelihood of fiscal consolidation, and potentially exacerbating the ongoing crisis in the euro area.[5]

Figure 2 summarises how events on election day mapped into financial markets. A first exit poll at 14:00 London time suggested the centre-left coalition would be the largest party

[4] For an account of events and political mood surrounding the Greek vote I refer readers to Hewitt (2013) pages 238-246.

[5] See, for example, "Spectre of instability haunts Italian voters", *Financial Times* 25th February 2013 for a preview of the election and coverage of the parties.

in both houses and likely be able to govern. Markets greeted the news positively. Spreads on 2-year Italian, Spanish and Portuguese bonds all declined by 5-10bp relative to Germany.[6] At 15:10 a second poll was released showing instead that the centre-right coalition would be the largest party in the Senate; bond prices fell, giving up all their previous gains. The final result was confirmed overnight, the vote ended up split three ways between the centre-left, centre-right and Movement 5 Star; while the centrist coalition failed at the polls. Italy had a hung parliament and faced the prospect of an unstable government. This was reflected in bond yields domestically (overnight the Italian 2-year bond spread jumped by 35bp between 16:30 on the 25th and 08:30 on the 26th) but also abroad with jumps of 16bp, 73bp and 11bp in Spain, Portugal and Ireland.

2.2 Discussion of the identification strategy

For clarity, the remainder of this paper adheres to the following terminology. The "event country" refers to the country where an event takes place. The "reaction country" refers to the country for which the bond market reaction is being recorded with "local" referring to the reaction country.

These examples above and in the introduction illustrate the transmission of foreign events to the borrowing costs of other Euro Area nations. Key for this paper's identification strategy is that the market move is not a function of other shocks that are hitting the reaction country's economy at the same time. Concerns about this claim can be grouped into three broad categories.

First, one may be concerned that there is direct causality between the events and contemporaneous economic shocks in the reaction country. This is unlikely in some circumstances, for example, I argued above that Catalonia's decision was unlikely to be in response to a shock to macroeconomic conditions in Italy in August 2012. However, there could be endogenous feedback between policy choices in one country from macroeconomic conditions in others – due to common shocks for example. The idea behind high frequency identification is that systematic reactions should be anticipated by market participants. The market reaction, therefore, is a surprise component, independent of preceding shocks that market participants are aware of.

Second, the identification strategy will break down if agents in the event country have

[6] See http://www.today.it/politica/elezioni/politiche-2013/exit-poll.html (in Italian) for details how results where released throughout the day.

private information about macroeconomic conditions in the reaction country that they reveal through their actions. This is why a focus on foreign events is necessary. For example, the local bond market response to a local announcement of an austerity package would also reflect the fiscal news shock contained in the package. Thus, local policymakers can reveal information about other local economic shocks when they make their policy announcements. For the same reason, events at a pan-European level (for example, the foundation of the EFSF in 2010) cannot be considered. Pan-European policies normally have some involvement from the ECB, which implies that they have a monetary component; market reactions would be correlated with monetary shocks. In general, however, it is unlikely that policymakers in foreign countries know more about local economic conditions than market participants.

Third, it may be that the foreign event is directly informative about local economic shocks. For example, data news is correlated across euro area countries and therefore a release in one country could be informative about economic conditions in another. The market reaction to data releases cannot be used as a result. Another channel via which events can be directly informative is if the market reactions are largely due to real or financial linkages between countries and so would be correlated with external demand shocks or shocks to the domestic financial system caused by losses abroad. However, there are reasons to think this is not the case. There were strong market reactions to events in Greece and in Cyprus, small countries from the European perspective with small direct linkages to the countries in my sample. In the Appendix, I present a formal analysis showing that interlinkages have no explanatory power over the relative market reactions to events.

There is still the question of why the event provokes a market reaction. As described in the Introduction, this could be due to self-fulfilling beliefs,[7] signals about political preferences or common creditors. The transmission could also run through the nature euro area as a system. A negative foreign event could reduce the capital, either financial or political, available to back a bailout of the reaction country's creditors by other members of the single currency.[8] Alternatively, as the future of the monetary union became less certain, a redenomination risk

[7]This could be due to a lack of commitment on the part of policymakers (Calvo (1988)) or when coordination failures among creditors lead to a sudden loss of liquidity (Cole and Kehoe (2000)). This feature of sovereign debt has motivated recent theoretical literature on the the euro crisis, with the observations concerned with how self-fulfilling crises play out in a monetary union (see, for example, Aguiar et al. (2015); Bocola and Dovis (2016); Corsetti and Dedola (2016)).

[8]Corsetti et al. (2006) studies a model where the resources of an international lender of last resort alters equilibria in a self-fulilling crisis.

premia may have entered borrowing costs. In general, I do not try to disentangle these channels, although with specific regards to redenomination risk I show below that this mechanism is not obviously present in the data. Importantly, however, these are all ways that sovereign borrowing costs can move independently of the local economy.

With this identification strategy in mind, I now turn to the construction of the dataset. I proceed in three stages. First, I use a news source to build a country specific narrative events during the crisis. Then I time when these events occurred, and eliminate those that overlap with other pieces of relevant news. Last, I measure how the prices in financial markets responded to the events. In what follows, I describe these steps in turn.

2.3 Building the narrative

To isolate events I use the financial news source EuroIntelligence; this follows Beetsma et al. (2013).[9] This source compiles a daily European economic news briefing released in the morning and typically contains 10-12 paragraph long stories including a "headline" story which the editorial staff consider the main event for the day. This source has two main advantages. First, it is contemporary rather than retrospective so provides a narrative of events that were considered relevant at the time (see Romer and Romer (2017)). Second, the source is almost unique as a briefing that was primarily focused on the evolving crisis in Europe. Its pan-European nature also serves as a filter as a country specific event must be of sufficient international interest to make the briefing.

The daily news briefings are read manually over the period July 2009 to March 2013. To be classified as an event and included in the narrative, a news story must satisfy the following criteria: (1) the story must relate to a single crisis-hit country; specifically, either Greece, Cyprus, Portugal, Ireland, Italy or Spain;[10] (2) the event must be timeable in the sense that it is possible to isolate when it occurs so as to determine the market's reaction (more on this below); (3) the story must not be an editorial comment; a report about/by private firms (excluding credit rating agencies, the ISDA or IIF) or individuals (excluding policymakers and politicians); a piece of market commentary (i.e. just describing market prices); or an economic data release (for the reason described above).

[9] Beetsma et al. (2013) also provide details on how EuroIntelligence operates.

[10] Experiments with political events in non-crisis countries revealed that bond markets do not react strongly to this form of news and as such these countries are omitted for the sake of parsimony.

One exception to the exclusion of data releases are official revisions to past and future fiscal projections which were of key importance during the early stages of the crisis in Greece. The relatively low frequency of these numbers and the lag in their release prevents the market reaction to them being related to cyclical news.

This approach provides a *wide* set of country specific events over the most intense period of the crisis. To further enhance the claim of exogeneity, I then narrow the event inclusion criteria in two further dimensions.

First, I drop events which relate to a foreign intervention in the event country.[11] These events revolve around decisions from international creditors on bailout programmes and liquidity assistance. I exclude these for two main reasons: (i) as described above, international policymakers may be internalising the entire currency union when making their decisions meaning they are more likely to either be internalising local economic shocks or have private information about those shocks that are revealed by their actions and (ii) these announcements often occur as a package of different measures affecting multiple countries simultaneously, even if the main news is about one particular country. Second, I exclude events that related to the sovereign bond market itself. These events are primarily the results of bond auctions or decisions by credit rating agencies or clearing houses. I exclude these as the announcements also often come as packages: credit rating agencies often downgrade multiple banks and sovereigns at the same time. However, it is also the case that excluding them leaves a narrative that is composed of country specific events that are of a political nature in line with examples given in the previous Section. This helps for interpretation.

This results in a *narrow* set of events that I use for my empirical analysis. The complete wide and narrow event list along with the relevant quotes and timing is available in the supplementary materials.

2.4 Event timing and event windows

Events that occur within trading hours (08:00-16:30) are timed to the minute when the first headline related to the event *occurring* appears on the *Bloomberg* newswire. The bulk of events considered in the dataset are essentially announcements, speeches or statements to the press

[11] To give a concrete example of an event like this, the Eurogroup first agreed that it was willing to provide emergency loans to Greece on the 15th March 2010. This was a decision by foreign policymakers intervening in Greece and I drop it from the narrower set of events.

from an official source; therefore, the timing is not subjective. As a caveat, for this approach to be workable, news stories as they appear in the summary often have to be broken up into discrete announcements. The Italian election above is an example of this with the event broken down into two exit polls and the final result. Events such as speeches often take place for a period of time. For events that take longer than 20 minutes, I record the end time as the last relevant headline on the newswire (otherwise the end and start time are the same). I then construct event windows from 20 minutes before the start time of the event to 20 minutes after the end time. If events occur close together so the windows overlap, I combine them into one longer window. If it is impossible to identify an initial headline in an objective fashion or the event lasts for more than 90 minutes I consider the event as untimeable. I give an example of event timing and discuss these issues in more detail in Appendix C.3.

The considerations are a little different for events that occur outside of trading hours. There is a tradeoff. Omitting the market reaction to these events altogether risks throwing out critical information. However, the long time window between close and open means that there is more chance of another piece of important information being released and distorting the market's reaction. As a compromise, and in the benchmark specification, events that occur outside trading hours are included when computing market reactions if they are the "headline" story in the following morning news briefing. This implies they should be viewed as the most important European event that occurred overnight and thus represent what the market is reacting to at the open. I then run robustness tests with these overnight events omitted. For events that happen outside of trading hours I construct event windows from 16:30 to 08:30 the next trading day.[12]

Another concern is simultaneous events. The following steps are taken to ensure markets are actually reacting to the event in question rather than other contemporaneous news. The structure of the dataset means it is straightforward to remove foreign events that overlap with a local event. The *wide* set of events comes in use in this regard. I construct equivalent time windows around all the local events in the *wide* event list and omit any foreign events that overlap with those time windows when computing market reactions.

In addition to this, I omit events that overlap with data releases. The time of local, pan-European and certain international data releases are obtained from the Bloomberg economic

[12]The period between 08:00-08:30 is noisy and subject to spikes, thus for overnight reactions I record the market position at 08:30.

calendar and events that would overlap with windows around these releases are omitted. Events that overlap with ECB decisions and press-conferences are also not included. Last, any country-specific event that overlaps with the announcement of a pan-European policy intervention is omitted. Such events are isolated using the ECB's time line of the crisis and are timed in an identical fashion to the country-specific events as described above. See Appendix C.6 for the precise data releases etc.

Despite this, it is not possible to rule out that there was another piece of news at the same time with any given foreign event. However, key for identification is that another piece of macroeconomic news is not *systematically* occurring at the same time. Once one strips out coordinated policy actions, monetary policy meetings and data releases it is difficult to see how this would be possible.

Table 1 summarises how this procedure translates into events and to time windows by number. The simple reading of the news summary provides 691 events by the wide measure. This reduces to 425 in the narrow measure, this collapses to 274 unique event windows (primarily due to the omission of overnight events that were not headline news stories). The final row summarises the final set of event windows for each reaction country once other contemporaneous events have been omitted.

2.5 Measuring market reactions

The raw intraday data is sourced from Thomson Reuters DataScope (see Appendix C for exact data definitions and a discussion of data quality). I use one-minute aggregates of the high-frequency data. I focus mainly on the change in the mid-yield on the benchmark 2-year sovereign bond in the reaction country from the start to the end of the window relative to the change in the yield on the equivalent German sovereign bond. For robustness, I also consider the reaction of the 10 year bond but, as I will describe below, market reactions measured using this maturity have a weaker relationship with overall borrowing costs. Greece and Cyprus are not included as reaction countries for the purposes of the empirical analysis due to inconsistent availability of intraday bond market data and the break in the Greek yield series associated with the debt swap. Hence, I focus only on Italy, Spain, Ireland and Portugal when measuring reactions. Greek and Cypriot *events* are included however.

Table 2, summarises these market moves for the different reaction countries. The mean reaction is near zero implying a roughly even split between positive and negative events. Approximately 60% of the sum of squared market reactions across all events is due to Greek news. This share is due to number of events that are of Greek origin – it is not the case that markets are reacting more strongly to Greek news on average, merely that there are more Greek events to react to.

The distribution of market reactions is heavy tailed: the largest reactions are several standard deviations in magnitude. As a result most of the variation arises from a relatively small number of important events. To provide more clarity over key events in the narrative, Table 3 provides a summary of 24 selected important events that generate large moves of consistent sign across the different reaction countries. Together with the examples described in Section 2.1, this list explains over half the sum of squared market reactions to foreign events in Italy and Spain, and a third in Portugal and Ireland.[13] As can be seen, similar to the examples at the start of the Section, these events are domestic in nature and largely reflect political developments.

3 Foreign Events and Sovereign Spreads

I now turn to how the reactions to foreign events translates into overall sovereign spreads, and, importantly, share of the variation in overall borrowing costs that they explain.

Figure 3 plots the cumulative reaction to foreign events against the overall 2 year bond spread relative to Germany at a daily frequency. The two series track each other well in Italy, Spain and Portugal until the summer of 2012 (less so in Ireland). The divergence after Summer 2012 coincides with interventions by the ECB (the "whatever it takes" speech by President Draghi in July and the announcement of OMTs in September). The narrative is specifically designed not to pick up such monetary interventions. If the decline in spreads after 2012 is partly a function of ECB policy (as has been argued by, for example, Acharya et al. (2017) and Krishnamurthy et al. (2018)), such a divergence is unsurprising.

Just inspecting Figure 3 suggests that foreign events are passing through to borrowing costs at a daily frequency. To confirm this formally, I consider the relationship between intraday market reactions to events and the daily changes in spreads by running the following

[13] To be clear, when calculating these shares I exclude local events and events that overlap with other news in Table 3.

regression:

$$\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}, \tag{1}$$

where $\Delta s_{c,d}$ is the change in the sovereign spread for country c trading day d, $m_{c,d}$ is the sum of country c 's market reactions to any foreign events on trading day d and $w_{c,d}$ is a residual. Column (1) of Table 4 presents results from open to close on trading days where there is some meaningful foreign news (defined as $|m_{c,d}| > 2bp$).[14] It is not possible to reject a coefficient of unity which is consistent with high frequency moves translating one for one into the daily change.[15] On the days where there is news, about 29% of the daily move is explained by the intraday reaction to foreign event.[16] Column (2) recalculates $\Delta s_{c,d}$ as the difference in spread between the close in period $d-1$ and the close in period d and adds overnight events to $m_{c,d}$, again the coefficient is near one. Column (3) considers all trading days and sets $m_{c,d} = 0$ on days when there are no foreign events. The coefficient is still near one, but as one would expect, the share of variance explained is much lower. Last, column (4) confirms that the market reaction passes through to long maturities by considering the daily change in the 10 year bond yield (see Appendix A.1 for results where $m_{c,d}$ is constructed from changes in 10 year bond yields).

How foreign events translate into different types of borrowing costs also allows for statements about what risks the market move captures. Specifically, Krishnamurthy et al. (2018) argue that the spread between bonds for the same sovereign in different denominations and jurisdictions can reflect the potential risk of redenomination embedded in the price of the Euro denominated debt of crisis hit countries. However, from a measurement perspective foreign currency sovereign bonds are relatively rare and illiquid; stale pricing and unmatched maturities can introduce error. Moreover, as argued by Corradin and Rodriguez-Moreno (2014), the discount on dollar denominated bonds could reflect that they were treated less favourably by the ECB as collateral at its refinancing operations. De Santis (2015) instead suggests using

[14] I choose this cutoff as it roughly corresponds to a one standard deviation move in Italian or Spanish yields at the frequency of the event windows. The results are robust to dropping this restriction, looking at days where $|m_{c,d}| > 0bp$ in a coefficient estimate of 1.447 and an R^2 of 22%.

[15] Taking the point estimate of unity fully at face value would suggest a degree of undershooting in the event window. This could either be because information leaks out in advance or because the effect continues to propagate after the event.

[16] Given the average length of an intraday event window, if the spread was a near random walk at daily frequency randomly sampled event times would deliver an R^2 of 0.1.

the more liquid credit default swap (CDS) contracts and argues that the spread between dollar and euro denominated CDS, known as the Quanto CDS basis, should partly be a function of redenomination risk.

In Table 5, I repeat the daily analysis measuring $\Delta s_{c,d}$ with CDS prices. Column (1) and Column (2) confirm that intraday movements in bond spreads have a near one to one impact on both Euro and dollar denominated CDS prices at a daily frequency.[17] The key takeaways though are (i) that the reaction of dollar and euro denominated CDS are not statistically distinguishable from each other, hence the insignificant reaction of the Quanto basis in Column (3); and (ii) that the negative point estimate on the Quanto CDS basis is inconsistent with the foreign event generating an increase in redenomination risk. This conclusion is largely unaffected by using 10 year CDS prices (Column (4)), starting the sample in Q3-2011 due to poorer quality CDS data in the prior period (as recommended by De Santis (2015)), Column (5)), excluding Italy from the sample as G7 countries were free to redenominate debt without triggering a credit event in pre-2014 CDS contracts (Column (6)) or using the 5 trading day cumulative change to adjust for potential slow price reactions due to illiquidity in the CDS market (Column (6)).[18]

The Quanto CDS basis is still an imperfect measure of redenomination risk. As argued by Kremens (2018), the Quanto CDS basis also picks up a second factor: the compensation for the expected depreciation (or appreciation) of the euro versus the dollar in the event of default where there is no redenomination.[19] A clean measure requires comparing CDS contracts with different treatments of redenomination events – unfortunately such contracts only came into existence in 2014. Nonetheless, the results in Table 5 suggest that the component of sovereign spreads most likely to be correlated with redenomination risk are insensitive to foreign events.

[17] The fact that EUR CDS and EUR bond yields move in statistically indistinguishable fashion means that foreign events are not associated with a shift in the bond-CDS basis. A concern during the crisis is that CDS rates did not fully reflect credit risk embedded in bond yields due to uncertainty about what actually constituted a credit event. Bonds may have therefore traded at a discount relative to CDS but this discount is not obviously being shifted by foreign events.

[18] Nothing should be read into the flip in sign in the 5 day cumulative regression. The estimates are volatile and the 4 and 6 day cumulative regressions yield negative, insignificant, coefficients.

[19] To see this, consider buying \$100 of protection against Spanish default versus Eur100 of protection, in a world where euro membership is irreversible. Imagine that the USD/EUR exchange is initially 1-1; and that Spanish default would result in a haircut of 50% and cause a Euro depreciation of 10%. In the default state, the dollar denominated protection would pay out \$50, equivalent to Eur55, and the Euro denominated protection would pay out Eur50. The dollar protection is 10% more valuable in the event of default, but the spread between the Dollar and Euro denominated CDS has nothing to do with redenomination. A similar argument applies to bonds. De Santis (2015) argues that subtracting the German Quanto CDS basis adjusts for this but it is not clear why this would be the case nor does that transformation affect the results in Table 5.

In Appendix A.1, I show this is also true for the bonds used by Krishnamurthy et al. (2018). Taken at face value, this suggests that the source of contagion of foreign events is through other channels.

If the source of contagion stems from imperfections in bond markets (e.g, due to losses for current investors and a slow reallocation of capital from new investors) the reactions may only have a temporary impact on borrowing costs. For there to be a macroeconomic consequence, the impact of foreign events on sovereign spreads must persist beyond the trading day and explain meaningful share of variation in sovereign borrowing costs at a lower frequency. To confirm persistence, I consider a dynamic version of the specification in Column (3) of Table A1 and run the following pooled local projection across all trading days (where h denotes regression horizon):

$$s_{c,d+h} - s_{c,d-1} = \alpha + \beta^h m_{c,d} + w_{c,d+h}$$

Figure 4 presents the estimated impulse response. The impact of foreign events persists unambiguously for around 10 trading days. However, the point estimates are less stable at longer horizons with the impulse response first decaying after around 12 trading days but then recovering back to unity from around 15-25 trading days. Therefore, there is evidence that the reaction to foreign events has a lasting impact on sovereign spreads, that persists for a calendar month, rather than generating a temporary high frequency movement in borrowing costs. In Appendix A.1, I show this also applies for country specific models.

Having established persistence, I now turn to the importance of foreign events for lower frequency moves in sovereign spreads (from here on the t subscript denotes months). Figure 5 plots the aggregation of market reactions to foreign events in the month ($m_{c,t}$) against the monthly change in sovereign spreads ($\Delta s_{c,t}$). Reflecting Figure 3, there is a tight correlation between changes in the month on month change in spreads in Italy, Spain and Portugal; as before the relationship for Ireland is weaker. Table 6 presents regressions of $\Delta s_{c,t}$ on $m_{c,t}$. From a simple pooled regression (Column (1)) one can see that the $m_{c,t}$ explains around 6% of the pooled monthly variation in bond yields. Appendix A.2 shows that the estimates in Table 6 are unaffected by excluding events that occur overnight from $m_{c,t}$, but that the results are weaker when $m_{c,t}$ is measured using the reactions of ten year bonds; this motivates the use of the 2 year maturity for measuring market reactions.

Pooling masks heterogeneity among countries and Columns (2)-(5) present country-specific models. In Spain and Italy, 20-30% of the monthly move in spreads is explained by the aggregated market reactions to foreign events; the figure falls to 12% in Portugal. The lower pooled figure is due to the poor explanatory power of $m_{c,t}$ for Irish bond spreads. While I can not offer a conclusive reason for this, there are a couple of somewhat mechanical explanations. First, as discussed in Appendix C.2, the Irish bond market is less illiquid and prices are more volatile; this reduces the information content of $m_{c,t}$. Second, the movements in the Irish yield in July-August 2011 that explain a large share of the variation in the overall series coincided with a series of European decisions about Greece's bailout which are excluded from $m_{c,t}$. However, it could simply be the case that Ireland's borrowing costs were less sensitive to foreign events than in other crisis countries. Excluding Ireland, the pooled share of variance in sovereign spreads explained by foreign events rises to 16%.

This share of variance represents a lower bound. As emphasised by Gurkaynak et al. (2005) direct regressions of changes in assets prices on intraday market reactions to events, as presented here, have an errors in variables interpretation. The high frequency market reactions are an imperfect measure of the influence of external factors on bond spreads. Mismeasurement can occur if news about foreign events leaks in advance or continues to propagate after the time window. Foreign factors may also be influencing spreads in periods outside of the limited set of event windows identified in the narrative. Classical measurement error implies the regression coefficients, and correspondingly the R^2, are biased towards zero. Attempting to adjust for this error requires putting more structure on the data than the simple OLS regressions above. In the next section, I use the aggregation of reactions to foreign events as an external instrument in a monthly Bayesian VAR model estimated with macro data. This allows for a variance decomposition undistorted by attenuation bias (Mertens and Ravn (2014)), as well as an estimate of the macroeconomic effects of higher sovereign spreads identified through foreign events.

Encouragingly, the F-statistic in Column (1) of Table 6 is 10.9 suggesting that, on a pooled basis, $m_{c,t}$ is above the commonly used rule of thumb for a relevant instrument (although concerns about weak instruments are less pertinent when Bayesian methods are used; see Caldara and Herbst (Forthcoming)). It is also the case that exploiting the cross-section of countries is necessary given the short sample: inspecting Columns (2)-(4), $m_{c,t}$ only satisfies the relevance criteria in Spain when the regressions are run on a country by country basis.

Beyond instrument relevance, which has been the main focus of this Section, for instrument validity $m_{c,t}$ must also be exogenous to other macroeconomic shocks that hit the reaction country at time t. As already outlined above, there are good reasons to think this is the case. And, while exogeneity is not directly testable, I have also carried out a battery of statistical checks see if the $m_{c,t}$ has properties consistent with an exogenous shock. For the sake of brevity, I refer readers to the Appendix B for exact details regarding these analyses and the associated tables and figures. Here I discuss the main conclusions.

The aggregated reactions to foreign events cannot be predicted by its past observations or macroeconomic data. I show this by running predictive regressions on the local market reaction to foreign events aggregated at weekly, biweekly and monthly frequencies. I find that these reactions are not explained by market reactions to past events, either local or foreign, or macroeconomic data. The reason for using high frequency data is that existing public information is already reflected in market prices; hence any systematic reaction of foreign agents to local shocks is orthogonal to the market reaction about an event. If the proxy were predictable then this claim would be questionable.

The time series $m_{c,t}$ is uncorrelated with market reactions to other local economic, monetary and fiscal news. I show this by calculating the local market reactions to local economic and fiscal data releases and to ECB announcements. I then calculate the correlation between $m_{c,t}$ and monthly aggregations of different types of data and monetary news. This can be thought of as a test of whether the proxy is correlated with the local macroeconomic shocks that are being captured by a data surprise or a central bank announcement.[20] The result of this analysis is mostly negative: $m_{c,t}$ is uncorrelated with market reactions to data surprises and ECB meetings. Similarly: *aggregations of market reactions to foreign events have a low correlation with reactions to local events and such correlations are inconsistent in sign across countries.* Such correlations being near zero suggest there is not a systematic transmission from local events, which could be a function of local economic shocks, to foreign events.

Real and financial linkages do not explain an economically meaningful share of the relative market reactions to events. If direct linkages explained market reactions to foreign events then the $m_{c,t}$ could be correlated with shocks to external demand or to the balance sheets of local

[20] This is an imperfect test as the causality could run in the other direction; for example, events that raise yields may lower confidence and cause negative survey releases or provoke an ECB reaction.

financial institutions and the exogeneity assumption would be violated. To check whether this is the case, I exploit the richness of the dataset and conduct a regression analysis, using a variety of specifications, attempting to explain the relative market move in response to an event across reaction countries using the size of trade and financial linkages with the event country. The message from this analysis is that trade and financial linkages are unimportant: the effect is insignificant and inconsistent in sign.

4 Macroeconomic Impact

To trace through the impact of a shock to spreads to the macroeconomy, I use $m_{c,t}$ as an external instrument (see Stock and Watson (2012); Mertens and Ravn (2013)) in partially pooled panel VAR model estimated using Bayesian methods (see Jarocinski (2010)). As the F-statistics in Table 6 illustrated, the short sample means that pooling estimates across countries is important for statistical precision. Yet there are differences across countries which implies that imposing homogeneity is potentially an overly strong ex-ante assumption. As the name suggests, the partially pooled model is a compromise between these two extremes: it exploits the fact that the countries share similarities, such that the cross-section can be used to enhance the precision of the estimates but does not impose homogeneity. The next section offers a brief sketch of the model structure; Appendix D provides a full discussion along with the posterior sampler.

4.1 Methodology

For each country the reduced form VAR is of the form:

$$y_{c,t} = \sum_{l=1}^{L} B_{c,l} y_{c,t-l} + \Gamma_c z_t + u_{c,t}, \tag{2}$$

where c and t are as above, the index $l = 1, ..., L$ denote VAR lags, $y_{c,t}$ is a $N \times 1$ vector of endogenous country variables, $B_{c,l}$ is the matrix of country-specific coefficients on lag l of the endogenous variables, z_t are deterministic variables with corresponding coefficient Γ_c and $u_{c,t}$ is the vector of i.i.d VAR innovations with distribution $u_{c,t} \sim N(0, \Sigma_{c,u})$, where $\Sigma_{c,u}$ is a covariance matrix to be estimated. Let $x_{c,t} = [y'_{c,t-1}, ..., y'_{c,t-L}]$; stacking the T observations on $y'_{c,t}$, $x_{c,t}$ and z'_t vertically to create data matrices allows the model to be expressed as: $Y_c = X_c B_c + Z_c \Gamma_c + U_c$.

Where $B_c = [B_{c,1}, ..., B_{c,L}]$'. Lastly, I define the vectorised data and parameter terms as $y_c = vec(Y_c)$, $\beta_c = vec(B_c)$ and $\gamma_c = vec(\Gamma_c)$. The likelihood for the model corresponding to country c is given by

$$p(y_c|\beta_c, \gamma_c, \Sigma_{c,u}) = N((I_N \otimes X_c)\beta_c + (I_N \otimes Z_c)\gamma_c, (\Sigma_{c,u} \otimes I_T)). \tag{3}$$

The country slope coefficients β_c are assumed to have an exchangeable Gaussian prior with common mean $\bar{\beta}$: $\beta_c|\bar{\beta}, \Lambda_{c,\beta} \sim N(\bar{\beta}, \lambda_\beta L_{c,\beta})$. The parameter vector $\bar{\beta}$ is the slope coefficients in a cross-country average model. The covariance matrix is decomposed into a country-specific positive definite matrix ($L_{c,\beta}$) and a common scale parameter contained in the set of positive real numbers (λ_β). The matrix $L_{c,\beta}$ is deterministic and is constructed from the ratios of the variances of the residuals from univariate autoregressive estimates of endogenous country variables as described in Jarocinski (2010). What matters for the tightness of the parameter estimates about the common mean is λ_β which acts as a scale parameter for the overall variance of the slope parameters across countries. An estimate of $\lambda_\beta = 0$ is equivalent to a homogeneous slope panel VAR. Conversely, $\lambda_\beta \to \infty$ implies estimates of β_c are almost equivalent to those as if each country has been estimated separately. The parameter λ_β, therefore, determines how close the model is to either of the two extremes of country-specific and homogeneous slopes. It is desirable to let the data determine how similar the countries are; hence, I use a non-informative inverse-Gamma prior: $\lambda_\beta \sim IG_2(0, -1)$.[21] I depart from Jarocinski (2010) by imposing the prior that the covariance matrix of the residuals is also drawn from a cross country distribution

$$\Sigma_{c,u}|\bar{S} \sim iW(\bar{S}, N + 2). \tag{4}$$

The purpose of equation 4 is to formalise the existence of a cross-country average covariance matrix, alongside $\bar{\beta}$, for use in calculating the impulse responses of the cross-country average model. This prior implies that the posterior of $\bar{S}$ can be used to estimate a cross-country

[21]This prior is specified, as recommended Gelman (2006), such the standard deviations for the individual coefficients in the VAR have a uniformly distributed prior over the positive portion of the real line, i.e. $p(\lambda_\beta) \propto \lambda_\beta^{-1/2}$. An alternative is to set the shape and scale parameters to an arbitrary small number - i.e. to approximate $p(\lambda_\beta) \propto 1$. This means the variance rather than the standard deviation approaches the uniform prior. However, Gelman (2006) shows that this can have an unforeseen impact on the posterior as the prior density has a fat right tail which places less weight on cases where the models are very similar (and λ_β is small).

covariance matrix centered around the harmonic mean of the individual country estimates.[22]

To identify a shock with an external instrument, I make the standard SVAR assumption that there exists a full rank matrix $\mathcal{A}_c$ such that $\mathcal{A}_c u_{c,t} = \zeta_{c,t}$, where ζ_{ct} is an $N \times 1$ vector of uncorrelated structural shocks of unit variance. The vector of structural shocks can be partitioned into the shock to the sovereign spread and the other structural shocks $\zeta_{c,t} = (\varepsilon_{c,t}, \tilde{\varepsilon}'_{c,t})'$. The critical assumption for identification is that $m_{c,t}$ is a linear function of the spread shock and no other: $m_{c,t} = \phi_c \varepsilon_{c,t} + \omega_{c,t}$, where $\omega_{c,t}$ is measurement error uncorrelated with any shock such that $\mathbb{E}(m_{c,t}\tilde{\varepsilon}_{c,t}) = 0$ and ϕ_c is an arbitrary scalar. The $\mathcal{A}_c$ matrix can also be partitioned such that $\mathcal{A}_c = [a'_{c,1}, a'_{c,2}]'$ with $\varepsilon_{c,t} = a_{c,1}u_{c,t}$, such that

$$m_{c,t} = \phi_c a_{c,1} u_{c,t} + \omega_{c,t} = \Upsilon'_c u_{c,t} + \omega_{c,t}. \quad (5)$$

The parameters in equation (5) have a similar prior structure as the parameters in the reduced form VAR: the country slope coefficients Υ_c have a Guassian prior with a common mean, $\bar{\Upsilon}$, and variance, $\lambda_\Upsilon L_{c,\Upsilon}$: $\Upsilon_c|\bar{\Upsilon}, \Lambda_{c,\Upsilon} \sim N(\bar{\Upsilon}, \lambda_\Upsilon L_{c,\Upsilon})$, with $L_{c,\Upsilon}$ set along the same lines as $L_{c,\beta}$ and the parameter λ_Υ playing the same role, with the same prior as λ_β, above. As is well known, Υ_c identifies $a_{c,1}$ and $\varepsilon_{c,t}$ up to sign and scale. Last, I Studentise the errors in equation (5) to account for the heavy tailed, infrequently observed nature of the event variable: $m_{c,t}|u_{c,t}, \Upsilon_c, \sigma^2_{c,w}; \nu \sim t(\Upsilon'_c u_{c,t}, \sigma^2_{c,\omega}; \nu)$ with ν calibrated to match properties of the market reactions to events (see Appendix D). All other parameters have diffuse priors.[23] Note, one implication of using Bayesian methods is that the reduced form model and equation (5) are estimated jointly; there is no separate first and second stage regression and the estimates of β_c depend on Υ_c.

Last, how should one think about $\varepsilon_{c,t}$ in the context of the theoretical literature on the macroeconomic impact of changes in sovereign spreads? There are a few potential interpretations. First, sovereign default is a political process, $\varepsilon_{c,t}$ could be modeled as a shock to creditors' perceptions of the parameters governing a politically determined fiscal limit (Bi (2012),Corsetti et al. (2013)). Second, $\varepsilon_{c,t}$ could be a non-fundamental shock, picking up shifts between belief driven equilibria (Bocola and Dovis (2016)). Third, and least concretely, a branch of the litera-

[22] I use a value of $N+2$ for the degrees of freedom parameter (as suggested in Giannone et al. (2012)) as this value imposes minimum shrinkage conditional on the existence of a prior mean for $\Sigma_{c,u}$. Allowing the residuals to be correlated across countries is attractive from an efficiency perspective but is computationally intensive. Re-estimating the benchmark specification allowing for such correlations does not meaningfully alter the results. Nor are the estimated correlations between residuals across countries large. Thus for computational convenience I restrict the residuals to be uncorrelated.

[23] Specifically: $p(\bar{\beta}) \propto 1$, $p(\bar{S}) \propto |\bar{S}|^{-0.5(N+1)}$, $p(\bar{\Upsilon}) \propto 1$ and $p(\gamma_c) \propto 1$.

ture has treated the sovereign spread itself as a time series process with an exogenous shock (Neumeyer and Perri (2005); Uribe and Yue (2006); Bocola (2016)).

4.2 Specification

I retain a focus on the same four crisis-hit Euro Area countries: Italy, Spain, Portugal and Ireland.[24] In terms of the included variables: output is proxied on a monthly basis using the unemployment rate and a broad index of industrial production, including the manufacturing, energy, utilities and construction sector. Prices are taken as the core HICP reading. Given the context of this paper, it is natural to add the borrowing cost of local sovereign: consistent with the previous Section I use the monthly average yield of the benchmark 2-year bond in each country. To capture the impact of elevated sovereign borrowing costs on private financial conditions in a concise manner, a composite measure of the private cost of finance is used. This is calculated as the weighted average in the cost of equity, debt securities and bank credit for non-financial corporations and households in each country. The series is computed internally by the ECB by weighting yields on the various sources of finance in accordance with flows of new lending. I return to the issue of the pass through of sovereign spreads to private financial costs below. As a measure of the fiscal stance, the monthly general government primary balance is included as an annualised percentage of nominal GDP. Comparable fiscal data (across the countries) is available only at a quarterly frequency from Eurostat's flow of funds dataset. However, all four countries publish monthly fiscal data using a variety of definitions. To generate a monthly fiscal series that has the same definition across countries I use the regression based interpolation methodology of Mitchell et al. (2005) on the quarterly series using the country specific monthly fiscal balances as interpolands. A deficit is a negative reading. Last, I include the German 2 year bond yield as a measure of risk free rates and to enable consistency with $m_{c,t}$ which is computed as a spread over Germany. All details of data definitions and construction are in Appendix C.

This sets $N = 7$. The set of deterministic variables, Z, is set to include country specific constants. The trended series (the CPI and Industrial Production) enter the VAR as annual

[24] That $m_{c,t}$ has weak explanatory power for Ireland is not necessarily problematic. The advantage of the prior structure is that information from the Italian, Spanish and Portuguese data is used to discipline the estimates for Ireland. Furthermore, as discussed, weak instrument considerations are less relevant for inference when Bayesian methods are used.

log differences; other series are included in levels. The lag length L is set to 2.[25] To enhance precision of the estimates I run the VAR over the period from January 2007 to March 2013 with $m_{c,t}$ taking a value of zero prior to July 2009. This assumption can be justified as (i) sovereign spreads were very small prior to the crisis so it makes sense to assume that shocks to spreads were very small also and (ii) foreign events had minimal transmission prior to the crisis period, this is clearly seen by inspecting Figure 3 where the cumulative market reactions do not differ meaningfully from zero until 2010. However, it is possible to estimate the identification equation over a different sample period from the reduced form equation. Restricting the observations on $m_{c,t}$ to post July 2009 yields similar point estimates and the key result of a statistically significant contractionary effect on the economy is maintained but it results in posteriors with much fatter tails (see Appendix A.3).

4.3 Benchmark impulse responses and variance decompositions

Figure 6 presents the impulse responses to a sovereign spread shock scaled to be consistent with a 100bps increase in the 2-year bond yield on impact using the mean country model (constructed from the estimates of $\bar{\beta}$, $\bar{\Sigma}$ and $\bar{\Upsilon}$). Several features are apparent. The impact of the shock on borrowing costs is relatively short-lived declining steadily such that after 18 months the impact has dissipated. Part of the explanation for this correction may lie in the soothing impact of policy: easing of risk free rates follows the shock, albeit with a lag, with a peak response of a 25bp decline in the German 2 year rate after 4 months. The unemployment response is statistically insignificant on impact but the shock propagates and leads to a peak response of 0.4ppt after 20 months. The response is also persistent. The response of industrial production is imprecisely estimated on impact, but the median shows the growth rate has a negative response of 0.8ppt after 12 months.

Figure 7 presents the results of the country-specific models. These are less precisely estimated than the mean country model but, in general, what stands out is the similarity of the responses. The data is returning models which are close to the mean country estimates. This is evidence that the countries did behave similarly in response to innovations to sovereign spreads during the crisis period.[26] In the Appendix A.3, I present both the pooled and un-

[25] Due to the short sample period I use a parsimonious lag selection procedure. I do this by testing up: starting by setting $L = 1$ and adding more lags until the median estimated residuals display no serial correlation. This lag-selection matches the Schwarz-Bayesian criterion assessed on the homogeneous parameter version of the VAR.

[26] One notable difference is that the Italian unemployment response is less significant than in other countries.

pooled models to show the affect of partial pooling. The Appendix also shows that the general pattern of impulses in Figure 6 is robust, to differing degrees of statistical significance, to: (i) excluding overnight events from $m_{c,t}$; (ii) working with 10 year bond yields; (iii) altering the treatment of trended variables and (iv) perturbing the lag order.

Figure 8 presents the forecast error variance decomposition. First, note that the decomposition suggests that 25% and 15% of the forecast error variance of unemployment and industrial production growth is explained by the sovereign spread shocks at a horizon of greater than 18 months. Along with the impulse responses, this suggests there are more persistent consequences of the shock and that sovereign spreads contributed heavily to the variation in activity over the crisis period. Second, the decomposition reveals that on impact around 50% of the variation in the bond yield is explained by the spread shock. This implies that a substantial portion of the variation in borrowing costs appears to be explained by shocks to spreads orthogonal to economic conditions and identified through foreign factors. This share of variance is larger than the explanatory power of $m_{c,t}$ in a direct regression on $\Delta s_{c,t}$ as presented in Table 6. This is for two main reasons. First, from a technical perspective, by construction a pooled least squares estimator up weights the countries with volatile spreads (Ireland and Portugal). The prior in the partially pooled model instead tilts the posterior of the mean country parameters towards the countries that can be estimated with more precision (Italy). Second, the variance decomposition in Figure 8 asks what share of the variation in borrowing costs is explained by the identified *shock*, $\varepsilon_{c,t}$, not what share is explained by $m_{c,t}$. This generates a larger figure as $\varepsilon_{c,t}$ omits the measurement error $\omega_{c,t}$ which is potentially biasing down the estimates in the previous section. However, adjusting for $\omega_{c,t}$ requires an estimate $u_{c,t}$. The reduced form residuals are a model dependent output from the VAR and hence require placing a particular structure on the data; this is a disadvantage compared to the direct regression.

To corroborate the findings from the variance decomposition, Figure 9 presents an historical decomposition asking what the counterfactual path of bond yields would have been if the identified shocks had not occurred.[27] On the top panel is the actual data versus the median counterfactual time series for the 2-year yield; the bottom panel has the difference between the

However, in the Appendix I confirm that an increase in sovereign spreads is contractionary in Italy in an unpooled model.

[27]Specifically, for each draw from the posterior, a time series of sovereign spreads shocks is extracted; the draws of the model parameters can then be used to remove the impact of these shocks from the data. This is equivalent to a counterfactual dataset where no sovereign risk shocks occurred over the course of the sample.

two series with accompanying credible intervals.

Upon the intensification of the crisis in 2011-2012, statistically significant differences are apparent between the actual and counterfactual bond yield.[28] The peak in the median estimated difference is around 100bp for Spain, 150bp for Italy, 400bp in Ireland, and around 700bp for Portugal. Between 40-60% of the trough-to-peak move in yields across the four countries can be explained by sovereign spread shocks. The pattern varies across countries: Italy suffers from two periods of deviations between counterfactual and actual yields, first over the autumn of 2011 and then in the spring of 2012. Both periods are contemporaneous with political instability in Greece, with the fall of the country's government followed by an indeterminate election. Spanish yields also peak around the Greek election and in November 2011. Portugal and Ireland suffer an extended run where yields deviate from the counterfactual, peaking around the summer of 2011.

Given that $\varepsilon_{c,t}$ captures a shock to spreads independent of economic conditions, an interpretation of the results in Figure 9 is that, at certain points in the crisis, sovereign bond yields divorced themselves from a level purely justified by macroeconomic conditions. In that sense, the results compare somewhat favourably with the estimates in Bocola and Dovis (2016). They estimate using a structural model, embedding multiple equilibria, that shocks unrelated to fundamentals, in their case rollover risk, added 150bp to Italian borrowing costs in 2012. This aligns with the estimate here.

4.4 Channels of pass through

The result that sovereign spreads were an important driver of unemployment and output dynamics over the crisis period raises the question: what is the evidence on the channels via which shocks to sovereign spreads pass through to the economy? There is no obvious direct channel through fiscal tightening. On impact the median impulses show that a sovereign spread shock reduces the fiscal balance in all countries (although this effect is not statistically distinguishable from zero). Since the data series is defined as the primary balance this response is not an automatic reaction to a higher interest burden. Instead, it is likely a reflection of lower revenues due to a weakening economy. Note also that the response of the primary balance is near zero at the same point where the unemployment response peaks. So if one de-

[28]Note that only 68% intervals are presented. At the 90% level, there are statistically significant differences in Portugal and Italy but not in Ireland or Spain.

fines fiscal tightening as an adjustment in the *cyclically adjusted* primary balance, then there is a potential contraction.

The macroeconomic literature has emphasised that a rise in sovereign spreads can act as a negative financial shock, tightening firms' and households' access to credit and thereby disrupting economic activity (see Neumeyer and Perri (2005), Uribe and Yue (2006), Corsetti et al. (2013)). There can be a direct effect on private sector borrowing costs if there is a sovereign ceiling whereby no firm can borrow at a cheaper rate from their sovereign due to the risk of expropriation (see Durbin and Ng (2005)) or because anticipated fiscal tightening reduces the net worth of firms causing financial constraints to bind. Alternatively, a change in spreads could alter the credit supplied by financial intermediaries. An increase in sovereign spreads may erode the value of subsidies from sovereign guarantees (Acharya et al. (2014)) or generate a direct shock to the net worth of banks through a decline in the value of bond holdings on their balance sheets, tightening funding constraints and reducing bank risk appetite (Bocola (2016)). Tighter private financial conditions then feedback to the sovereign by dampening activity, lowering tax revenues and potentially raising sovereign spreads further. Capturing this general equilibrium feedback loop is one of the advantages of using a VAR.

The contractionary affect on private financial conditions is already apparent with the increase in the composite measure in Figure 6, 100bp increase in sovereign borrowing costs raises the weighted average private sector costs of funds by 30bps (an effect that is stable across countries). To explore further the pass through onto private financial conditions, I follow Gertler and Karadi (2015), and re-estimate the baseline model including additional series individually, I then present the mean-country impulse responses to the additional series in Figure 10 (I show shorter horizons as the responses are less persistent). This averts the problems of overparameterisation and multicollinearity. I have verified that the additional series make little difference to the response of other variables.

I focus on the following variables country specific series: (i) the yield on corporate bonds;[29] (ii) the average interest rate on loans to the non-financial private sector; (these first two series are both components of the composite financial conditions indicator used in the baseline specification) (iii) the growth in the volume of bank credit to the non-financial private sector; (iv) the

[29] The Irish corporate bond yield has a severe spike to over 89% in early 2009. To prevent this distorting the results I have excluded Ireland from the model when estimating the impact of sovereign spreads on corporate bond yields. Including Ireland only strengthens the results with both incredibly tight error bands and an impact response of 250bps.

level of the country's Target2 balance (as a % GDP) to proxy capital flight; and the log monthly change in the Datastream equity return index for (v) banks and (vi) non-financials. Exact data sources are in Appendix C.

Inspecting the responses in Figure 10, scaled to be consistent with a 100bp increase in the sovereign yield on impact, it is clear that the increase in sovereign borrowing costs causes a sharp, 150bp increase in corporate bond yields. This is consistent with the sovereign ceiling and indicates that private borrowing costs also rise in response to a rise in sovereign borrowing costs. The other series are less conclusive, with the 90% credible intervals on the impulse responses all intersecting zero. However, taking the point estimates and the 68% confidence intervals at face value, the results suggest that loan rates are more sticky than bond yields, rising by 12bp, but loan growth slows by 0.5 percentage points at the trough of the response suggesting an adjustment through quantities. There is a persistent fall in the Target2 balance by around 2% GDP, consistent with a fall in net financing by foreigners. Moreover, this is evidence that the loss of access to external financing associated with sovereign *defaults*, which plays a key role in Mendoza and Yue (2012), is also apparent when there is an increase in sovereign spreads. Last, the final column in Figure 10, shows a cumulative -13% return on bank equities and a -4.8% return on non-financials in the 6 month following the shock consistent with a decline in the net worth of both financial intermediaries and non-financial firms. This finding is consistent with shocks to sovereign spreads lowering the market value of the net worth of firms and financial intermediaries, potentially tightening financial constraints.

As the equity return indices are daily, the less precise monthly VAR estimates can be corroborated using higher frequency analysis. Specifically, in Table 10, I re-estimate equation (1) replacing $\Delta s_{c,d}$ with the log change in the equity return index on the trading day. On a pooled basis, a foreign event that raises 2 year bond spreads by 100bp results in a -4.0% daily return on banks and -3.7% for non-financial firms. There is some heterogeneity across countries, with much larger coefficients in Spain (for banks) and Italy (for non-financials), which is partly a function of outliers. Nonetheless, this confirms that a rise in sovereign spreads does cause a reduction in equity values.

5 Conclusion

This paper uses a high frequency, narrative identification strategy relying upon market reactions to foreign events during the Euro crisis to obtain variation in sovereign spreads that are plausibly orthogonal to innovations to local economic conditions. This addresses an identification problem in the literature: discriminating between a change in the riskiness of the sovereign that is itself a function of macroeconomic conditions and the macroeconomic implications of fluctuations in sovereign risk. I show that foreign events were an important driver of sovereign spreads during the crisis suggesting spreads were not just pinned down by local economic conditions. The identified shocks to sovereign spreads also had macroeconomic implications in crisis-hit Euro Area countries. The identified shocks explain 20% of the forecast error variance in the unemployment rate. The main channel of pass through appears to be a deterioration in private financial conditions.

How generalisable these results are needs to be assessed. First, the euro area presents a rather specific circumstance: It is a monetary union where local sovereigns issue debt in a currency that they do not control. Euro Area sovereigns may be more vulnerable to belief-driven crises than those with a monetary backstop. The loss of access to the printing press may also limit the inflationary consequences of a crisis due to controls over monetary financing. Second, the results are from a short sample estimated during a crisis period. While the pooling across countries helps with inference, the estimates are still from a specific point in time with recessionary economic conditions. Therefore, the results refer to the macroeconomic effects of shocks to sovereign spreads during a crisis. However, since a crisis is when sovereign spreads become relevant this is not necessarily a severe limitation.

The dataset I provide has further applications for future research. Although this paper is largely silent regarding the explanation behind observed market reactions, the narrative dataset can also be used to explore the determinants of the transmission of sovereign risk between countries. For example, one could investigate if events certain countries were of importance at different times or which sorts of events markets were sensitive to. Another area for future research is whether these results can inform the micro-foundations of channels by which sovereign risk feeds through into the real economy.

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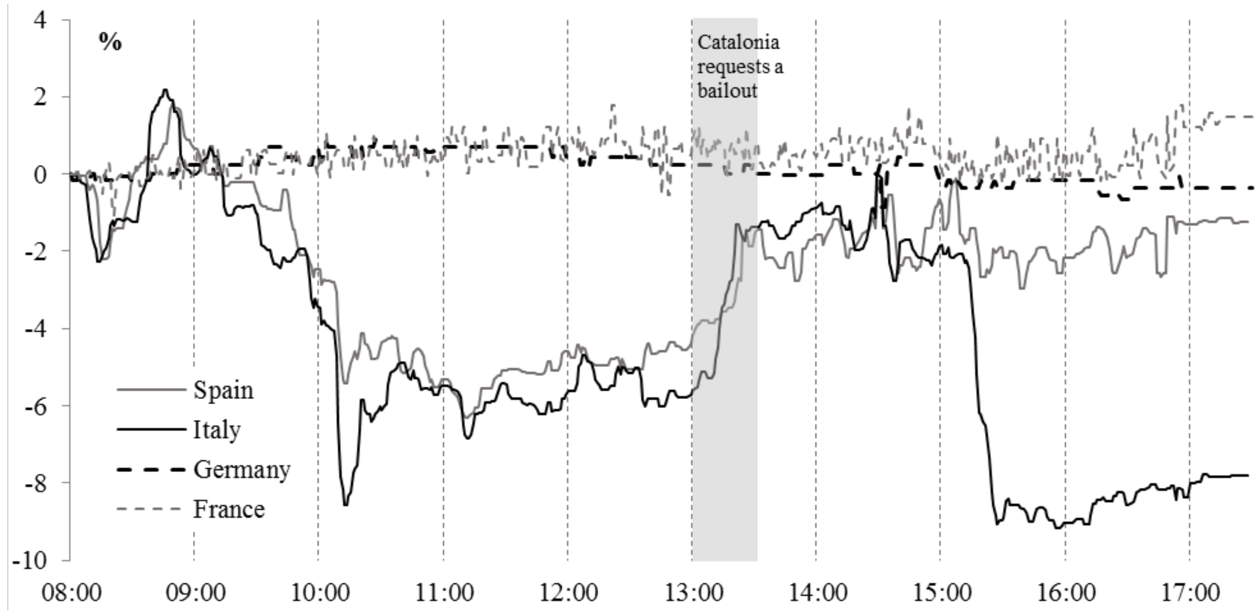
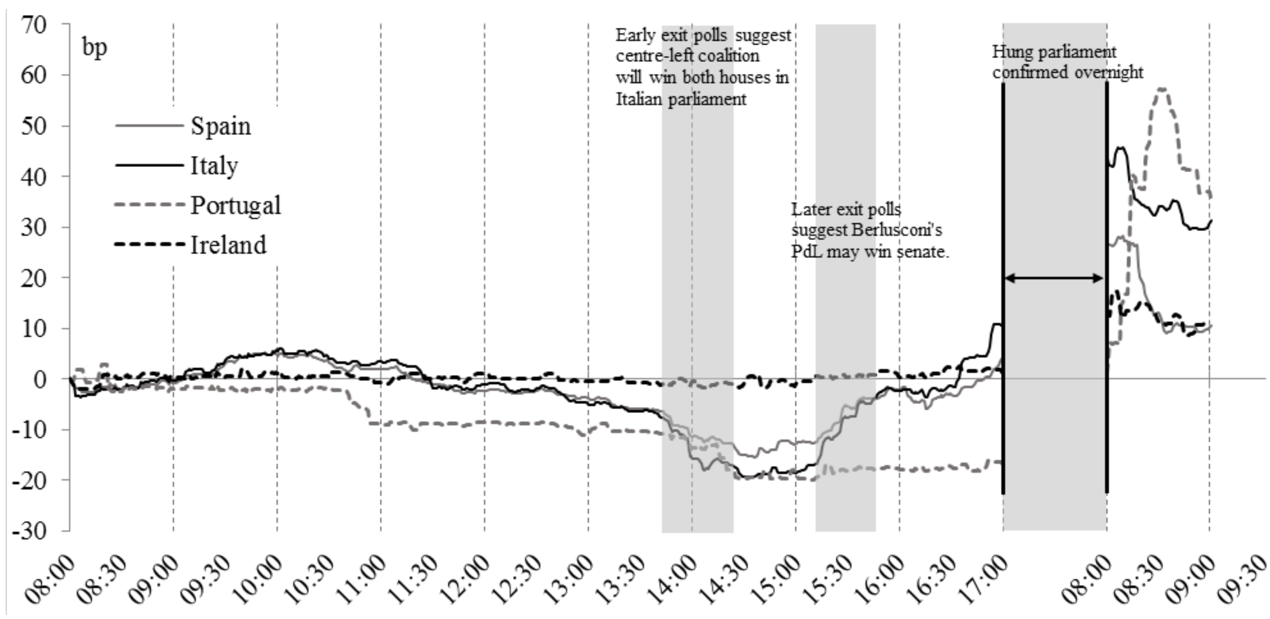
Tables and Figures

Figure 1: Catalonia Requests a Bailout, 28th August 2012: High Frequency Market Reaction



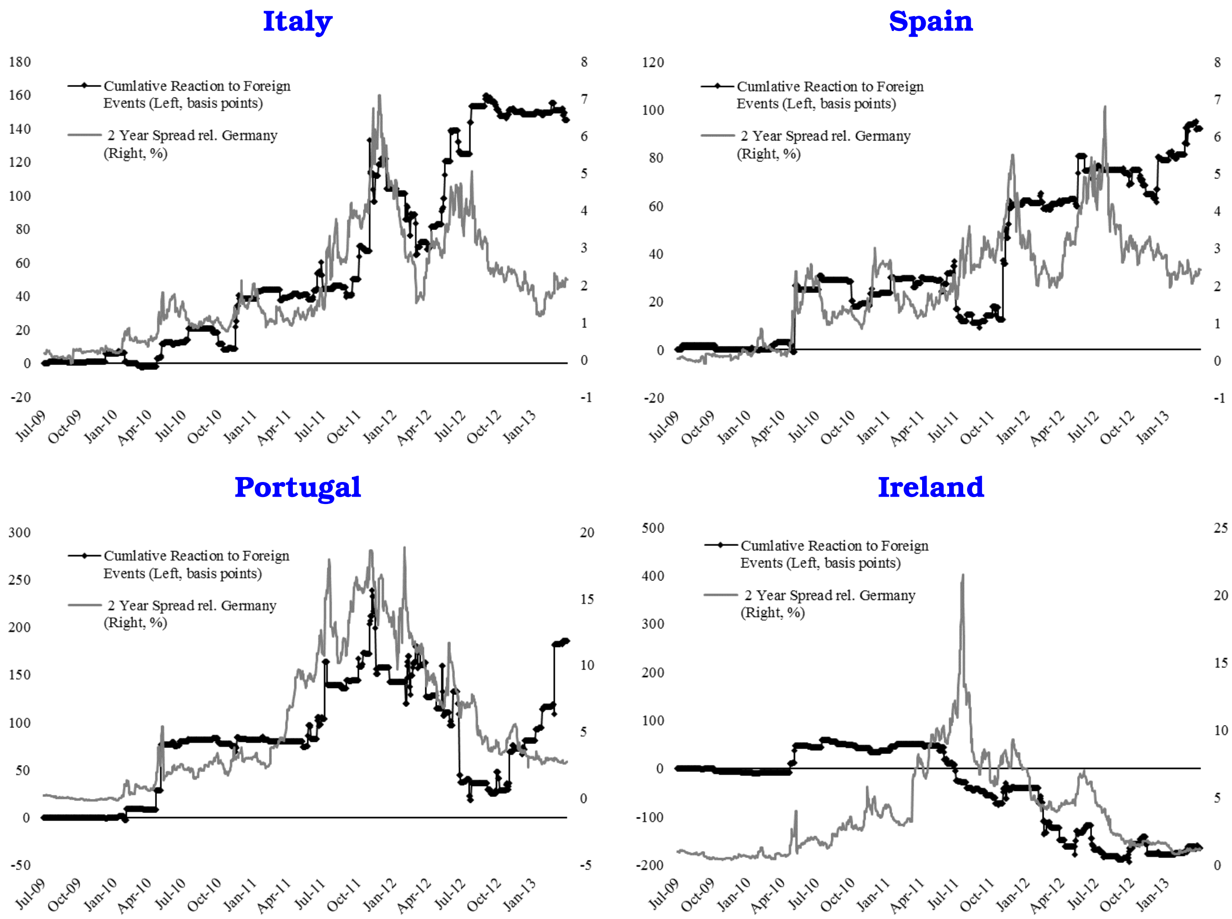
Notes: Intraday bond market moves in Spain Italy, Germany and France around Catalonia's decision to request a bailout from the Spanish regional liquidity mechanism (28th of August 2012). The series are centred 5 minute medians of the mid yield relative to 8:00am (Italy and Spain are presented as a spread relative to Germany). The x-axis refers to London time.
To give some context to moves throughout the day: Italian and Spanish yields first fell sharply, corresponding to successful short-term debt auctions in both countries. This was followed by an upward movement at 13:00 when Catalonia made its announcement. Italian yields then declined again at around 15:15, there was no obvious event that caused this (the move was not matched in Spain or with the Italian 10 year bond whose yield was stable). The French and German yields were stable throughout the day.

Figure 2: Indecisive Italian Election, 25th February 2013: High Frequency Market Reaction



Notes: Intraday bond market moves in Spain, Italy, Portugal and Ireland around Italy's indecisive election (25th of February 2013). The series are centred 5 minute medians of the mid yield of the 2-year bond relative to 8:00am on the 25th (the data are presented as a spread relative to Germany). The x-axis refers to London time.

Figure 3: Cumulative Reaction to Foreign Events and Sovereign Spreads



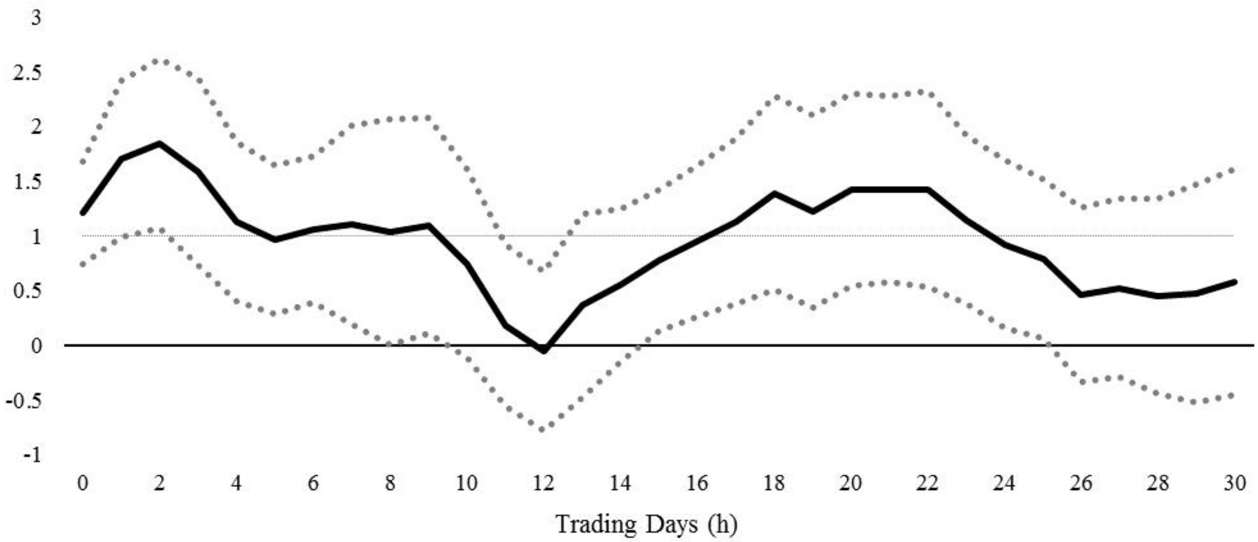
Notes: Daily time series plots of the daily 2 year sovereign spread relative to Germany (grey line) against the accumulated market reactions to foreign events.

Table 1: Numbers of events

Event Country:	Cyprus	Greece	Ireland	Italy	Portugal	Spain	Total
Number of identified events (wide measure)	25	226	88	115	108	129	691
Narrow Measure	13	133	49	94	62	74	425
Unique event windows (ex. non-headline overnight)	9	92	32	60	37	44	274
Included event window by reaction country:							
Ireland	8	77	–	49	29	40	203
Italy	7	79	28	–	33	35	182
Portugal	8	82	27	53	–	38	208
Spain	6	69	27	45	28	–	175

Notes: Breakdown of events and event windows by event country. Narrow measure excludes foreign interventions and market news. Unique event windows combines events into single windows if they occur at a similar time; drops events that overlap with something else and drops overnight events that are not the headline story in the news summary.

Figure 4: Local Projection Estimates of the Persistence of Foreign Events.



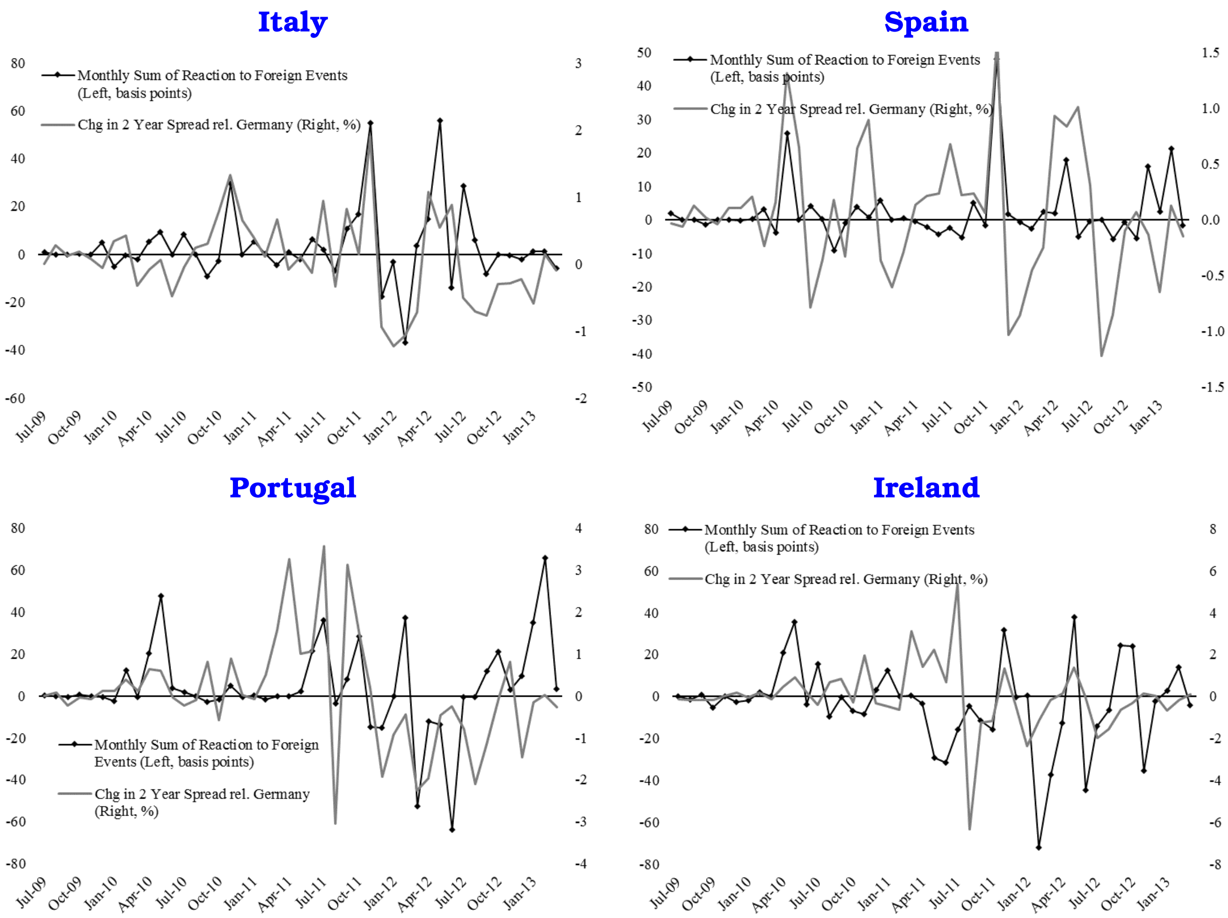
Notes: Impulse response to the following local projection: $s_{c,d+h} - s_{c,d-1} = \alpha + \beta^h m_{c,d} + w_{c,d+h}$, where: (i) $s_{c,d}$ is the 2 year sovereign bond spread at 16:30 London time on day d for country c; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day t (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. The y-axis is the estimates of β^h and can be interpreted as a cumulative percent change in the spread following a foreign event that raises borrowing costs by 100bps on day 0; x-axis is the horizon h. Grey dashed lines are 90% confidence intervals constructed from Driscoll-Kray standard errors with a 30 day lag window.

Table 2: Descriptive Statistics for the Market Reactions to Foreign Events

	Italy	Spain	Portugal	Ireland
Total Number of Event Windows	182	175	208	203
Share outside trading hours (%)	24.2	16.6	27.9	28.1
Mean Market Reaction (bp)	0.9	0.6	0.9	-0.8
Std. Dev. Market Reaction (bp)	6.2	4.2	15.1	8.5
Max Market Reaction (bp)	47.2	27.7	72.6	37.6
Min Market Reaction (bp)	-31.1	-22.1	-63.4	-34.8
Percentage of sum of squared reactions due to:				
Greece (%)	69.4	64.9	35.7	58.6
Italy (%)	0.0	22.3	42.1	10.5
Portugal (%)	3.5	6.3	0.0	5.3
Spain (%)	16.0	0.0	21.1	24.9
Ireland (%)	9.9	5.5	1.1	0.0
Cyprus (%)	1.1	1.1	0.1	0.7

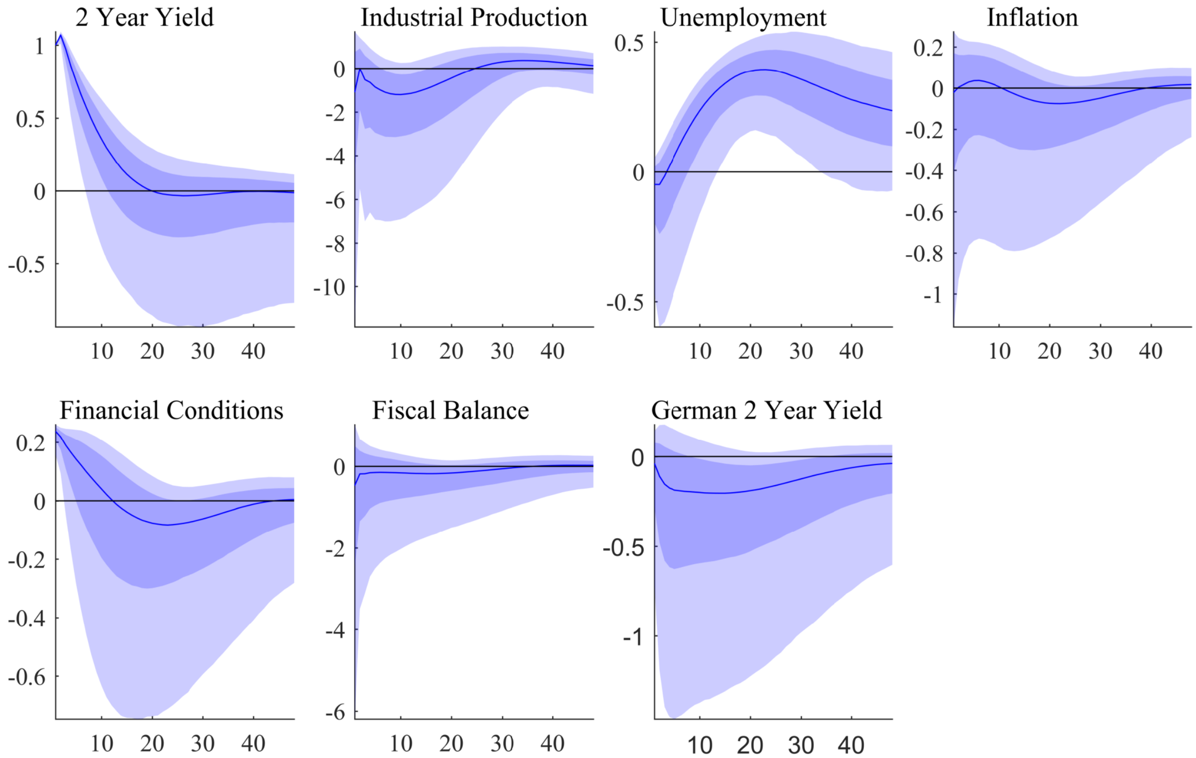
Notes: Market reactions in events windows included in the proxy variable satisfying the criteria in Section 2.3. Data period is July 2009 - March 2013. Market reactions refer to change in local 2 year bond yield relative to the change in Germany. The percentage shares refer to the share of the sum of the squared market reactions across all included event windows that can be attributed to events in a particular country. Percentages may not sum due to rounding.

Figure 5: Monthly Changes in Sovereign Spreads and Foreign Event Reactions



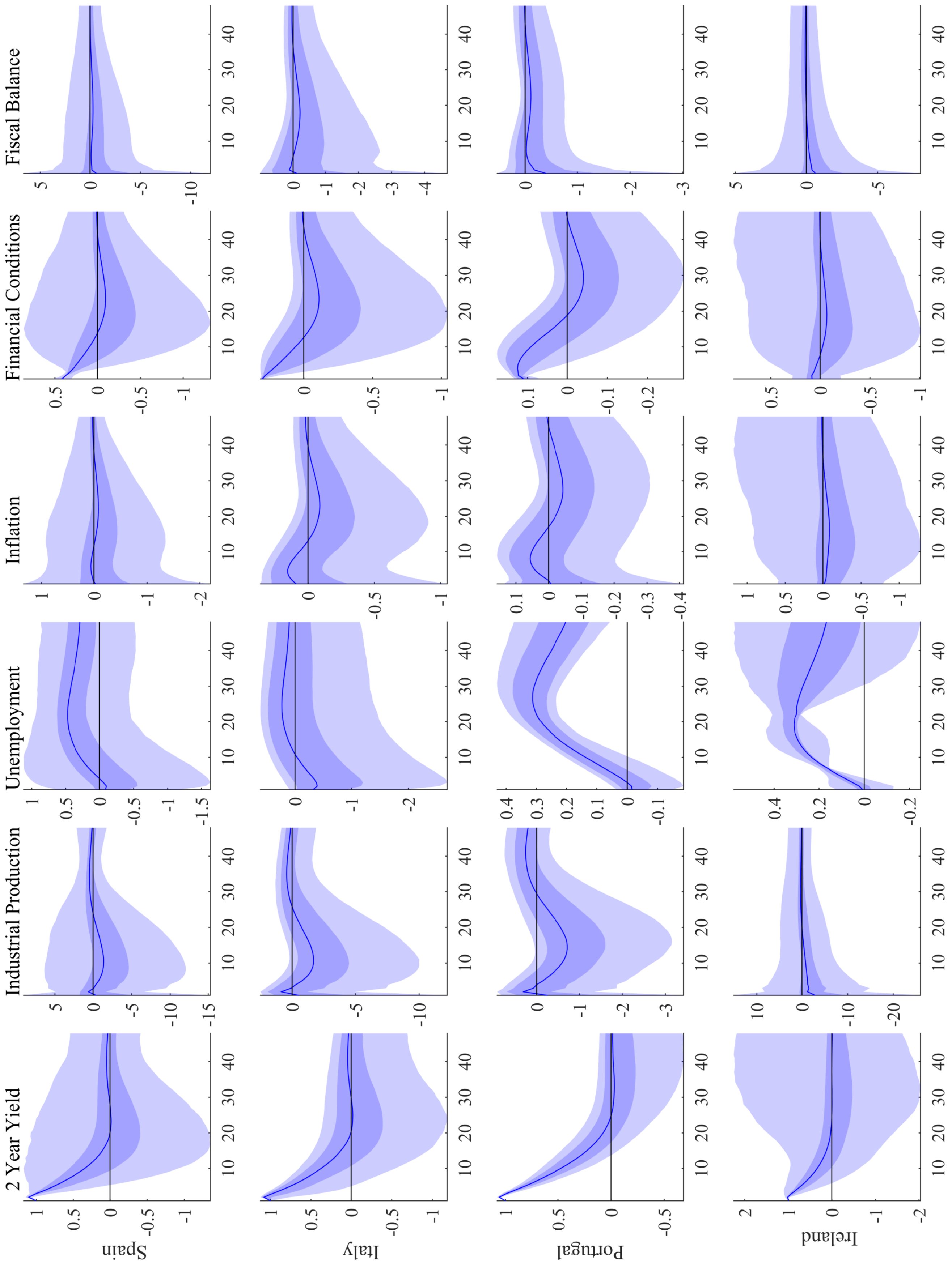
Notes: Cross-country comparison between the change in average monthly 2-year sovereign spreads (relative to Germany) and the monthly aggregate market reactions to foreign events. Black, dotted line is the monthly sum of market reactions to foreign events (left hand axis, in percentage points). Grey line is the change in the average sovereign spread (right hand axis, percentage points).

Figure 6: Mean Country Impulse Responses to a Sovereign Spread Shock



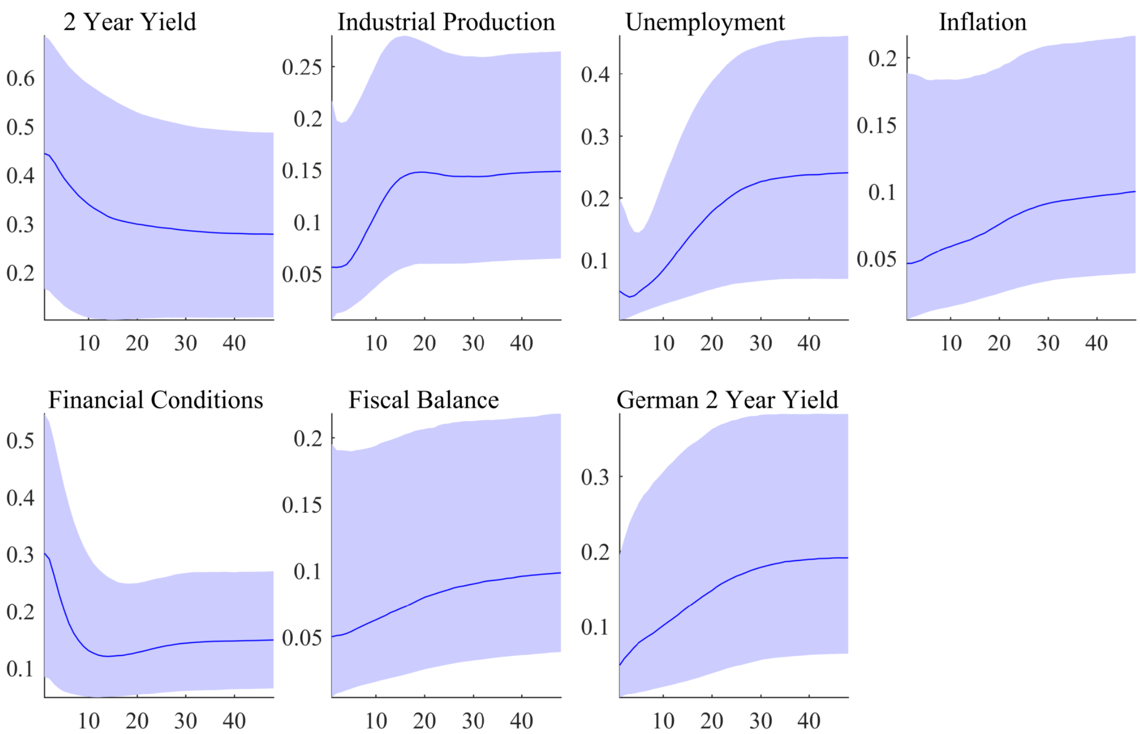
Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; for exact data definitions see the data appendix. Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{S}$. Centre line is the median of 10000 non-sequential draws from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

Figure 7: Country Specific Impulse Responses to a Sovereign Spread Shock



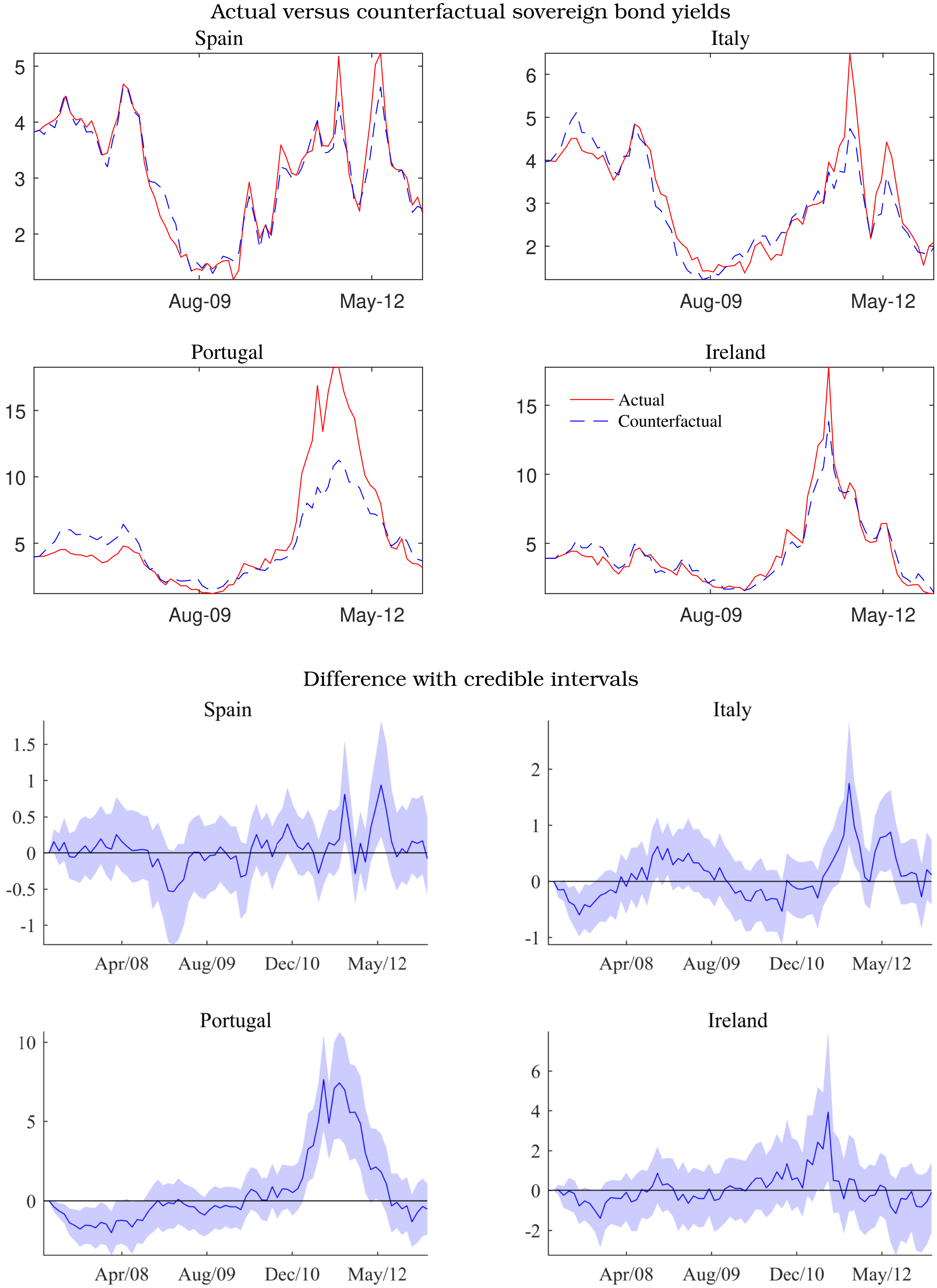
Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; for exact data definitions see the data appendix. Centre line is the median of 10000 non-sequential draws from the simulated posterior. Error bands are 68% and 95% Bayesian credible intervals. Response of German yield excluded for compactness and due to similarity with figure 6.

Figure 8: Forecast Error Variance Decomposition (Contribution of Sovereign Spread Shock)



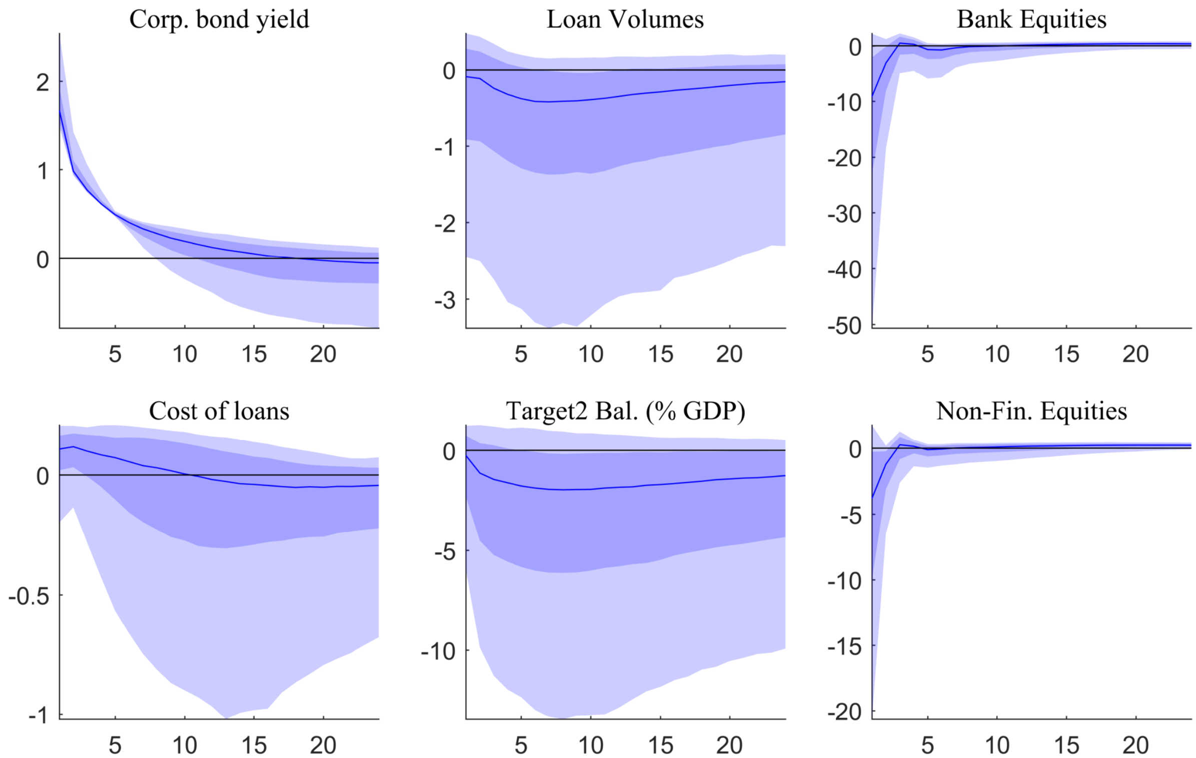
Notes: Share of forecast error variance explained by sovereign spread shock. X-axis is months. Blue line is the median of 10000 non-sequential draws from the simulated posterior. Decomposition computed from mean country model (using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{S}$). Error bands are 68% Bayesian credible intervals

Figure 9: Historical Decomposition: Bond Yields



Notes: Counterfactuals are constructed by zeroing the sovereign spread shocks and recreating the yield. Y-axis is percentage points. Centre line is the median of 10000 non-sequential draws from the simulated posterior. Difference (lower pane) is Actual−Counterfactuals. Error bands are 68% confidence intervals.

Figure 10: Passthrough of Sovereign Spreads to Financial Conditions



Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; for exact data definitions see the data appendix. The impulse responses are computed using mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{S}$. Centre line is the median, error bands are 68% and 90% Bayesian credible intervals.

Table 3: Selected Examples of Important Events

Date	Event Country	Event Description	Reaction (bp)			
			Italy	Spain	Portugal	Ireland
05/05/10	Greece	Three people are killed when a bank is set on fire during a national strike in Greece which led to violent protests. [Timing taken from first press report of the fire on the news wire as start time, with police confirmation as end time].	7.125	27.675	33.725	10.775
14/07/10	Spain	Prime Minister Zapatero delivered his speech in Spain's state of the nation debate in parliament and warned that more budget cuts are required to restore confidence in Spain.	6.725	5.850*	2.475	15.575
30/09/10	Ireland	Details of the Irish Bank Bailout programme are released. The Irish Finance Minister announces that as a result of support for the banking system, the general government deficit will be around 32% of GDP in 2010, with only modest corresponding participation of private bondholders. He adds that further fiscal consolidation is required.	-6.350	-8.400	-4.225	-8.350*
22/05/11	Spain	The ruling Spanish socialist party suffered a rout in Spanish regional elections, falling 10 points behind the opposition in weekend polls.	18.025	20.3*	11.75	4.175
30/05/11	Italy	Italian Prime Minister Silvio Berlusconi 's ruling PdL party suffered a severe loss, including the mayoralty in the conservative stronghold of Milan, in the 2011 Italian regional elections.	-5.750*	-6.325	-10.85	-4.025
15/06/11	Greece	Greek Prime Minister George Papandreou announces that talks with the opposition to form a new unity government to pass an austerity bill had failed and that instead he will reshuffle his cabinet and hold a parliamentary confidence vote in the government.	8.925	11.550	18.925	37.575
23/06/11	Greece	Greek opposition leader Antonis Samaras says he continues to oppose the austerity measured required by EU and that the programme needs "corrective measures".	3.075	5.175	5.500	3.850
27/06/11	Greece	Over the course of Sunday, four more PASOK MPs announce they will consider opposing the Greek government's critical austerity bill. Antonis Samaras announces his continued opposition to the austerity bill and a new poll shows 75% of Greeks do too. Greek Deputy PM Pangalos comments that he doubts that the Greek government has the votes to pass the bill.	7.175	2.125	6.975	1.8
28/06/11	Portugal	Portugal's new centre right government released its programme for the next four years, vowing it will "be more ambitious" and go beyond terms set out under the Eur78bn bailout. The programme confirmed the tax hikes agreed with the EU and the IMF as part of Portugal's bailout.	5.575	4.425	9.65*	-3.375
30/06/11	Greece	In parliament, the Greek government gains sufficient votes to approve the first part of an austerity bill critical to securing the next loan tranche in its bailout.	-8.125	-9.05	-1.6	-20.2
30/09/11	Greece	The Greek cabinet released the draft budget for 2012, with a deficit projected at 8.5% compared to a target of 7.6% of GDP set by the Troika.	13.225	12.225	23.025	18.200
01/11/11	Greece	The Greek ruling party PASOK's majority falls to 151 out of 300 as two MPs resign in quick succession. Within an half an hour, Six PASOK party members announce that they are jointly calling for Greek PM Papandreou to resign.	35.675	13.55	36.750	23.600
04-06/11/11	Greece	Greek PM Papandreou survives a no confidence vote on the evening of Friday the 4th, after which, at a press conference, he says a new government of national unity will be formed (perhaps without him in charge). The new Greek government of national unity is formed over the weekend and the hunt for the new PM begins. Talks on new PM fail to find a resolution but in statements to the press, following the meetings, between PM Papandreou and opposition leaders suggest progress was made.	47.225	13.7	26.550	10.250
10/11/11	Greece	Lucas Papedemos is announced as the new Prime Minister of Greece.	-7.175	-3.775	-8.00	-0.4
17/11/11	Italy	The newly appointed, technocratic, Italian Prime Minister, Mario Monti, presents his new government's policy programme to the Italian parliament.	-30.925*	-3.500	-43.325	-1.700
02/02/12	Spain	Spanish Finance Minister de Guindos gave a press conference to announce a new programme to provide Eur50bn of new capital for the Spanish banking system.	-5.100	-5.650*	-38.35	-34.75

Notes: Reaction refers to the change in the 2 year sovereign spread (relative to Germany) for each country. The description is a paraphrasing of the story by the author. Please see the supplementary materials for exact quotes. An asterisk denotes a reaction to a local event, these figures are given for comparison and are not used in the empirical analysis.

Table 3: Selected Examples of Important Events (Continued)

Date	Event Country	Event Description	Reaction (bp) Italy	Spain	Portugal	Ireland
09/02/12	Greece	Greek coalition leaders announce they have reached a compromise deal regarding the additional austerity measures required by the Troika.	-4.225	-2.200	-4.400	-6.025
17/06/12	Greece	Antonis Samaras' New Democracy emerges as the largest party in the 2012 Greek parliamentary elections; however coalition partners were required to form a government.	-10.275	-4.400	-7.525	-29.25
29/05/12	Spain	Spanish Central Bank governor, Miguel Angel Fernandez Ordóñez, announced his resignation in protest of the Spanish government's handling of the crisis at the bank Bankia.	18.250	15.550*	3.700	6.225
01/06/12	Ireland	The first count in the Irish referendum on the European Fiscal Compact is released revealing that the treaty will be approved comfortably by voters.	-0.800	-5.925	-4.225	-13.85*
25/06/12	Spain	The Spanish government announced it has formally submited a request to the European Financial Stability Facility to obtain funds to help recapitalise its banks.	-7.100	-2.500*	-7.875	-5.400
25/06/2012	Portugal	In a speech to parliament, and following weaker fiscal data, Prime minster Passos Coelho did not exclude further austerity but said "it is too early" to think about new measures. He vowed to meet targets laid out in the country's bailout agreement and rejected a no-confidence motion in his government.	5.575	4.425	9.650*	-3.375
06/12/2012	Italy	Silvio Berlusconi announced that he asked PdL MPs to abstain from supporting Italian Prime Minister Monti in a critical vote on the government's growth plan in the senate.	6.775*	3.850	5.225	5.800
08/12/2012	Italy	At a press conference, Italian PM Mario Monti announced that he plans to resign after the passing of the 2013 Italian budget.	25.925*	15.025	8.450	5.175

Notes: Reaction refers to the change in the 2 year sovereign spread (relative to Germany) for each country. The description is a paraphrasing of the story by the author. Please see the supplementary materials for exact quotes. An asterisk denotes a reaction to a local event, these figures are given for comparison and are not used in the empirical analysis.

Table 4: Contribution of Foreign Events to Daily Changes in Sovereign Spreads

	(1) Open-Close, event days	(2) Close-Close, event days	(3) All Days	(4) Ten Year
$m_{c,d}$	1.474***	1.205***	1.195***	0.636***
	(0.36)	(0.267)	(0.267)	(0.36)
N	219	319	3908	219
R^2	0.29	0.25	0.03	0.21
Robust F-stat	16.8	20.3	20.0	11.6

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}$, where: (i) $\Delta s_{c,d}$ is the change in the 2 year sovereign bond yield between 08:30 and 16:30 London time on trading day d for country c relative to the change in the German 2 year yield; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. Column (1): Pooled regression on days where $|m_{c,d}| > 2bp$ such that a meaningful quantity of foreign news is observed. Column (2): includes the overnight time period such that $\Delta s_{c,d}$ is calculated between 16:30 $d-1$ and 16:30 on trading day d; similarly $m_{c,d}$ is computed including events that happen overnight. Column (3): includes all trading days with $m_{c,d}$ set as zero when no news is observed. Column (4): equivalent to Column (1) with $\Delta s_{c,d}$ measured as the change in the 10 year sovereign spread.

Table 5: Foreign Events and the Quanto CDS Basis

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\Delta s_{c,d}$:	USD CDS	EUR CDS	Basis	Basis	Basis	Basis	Basis
	2yr	2yr	2yr	10yr	2yr, 2011Q3-	2yr, ex IT	2yr, 5 day cum.
$m_{c,d}$	0.9126***	1.0666***	-0.1540	-0.0766	-0.1760	-0.1635	0.1501
	(0.182)	(0.221)	(0.109)	(0.049)	(0.126)	(0.126)	(0.468)
N	319	319	319	319	230	250	319
R^2	0.19	0.16	0.02	0.01	0.02	0.01	0.00
Robust F-stat	25.22	23.23	2.00	2.45	1.95	1.68	0.10

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}$, where: (i) $\Delta s_{c,d}$ is the change in the relevant CDS spread on trading day d (close to close) for country c; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. Column (1)-(2): Pooled regression on the using 2 year Euro and USD CDS on days where $|m_{c,d}| > 2bp$ such that a meaningful quantity of foreign news is observed. Column (3): $\Delta s_{c,d}$ is the change in the Quanto CDS basis, the difference between the USD denominated and EUR denominated CDS prices (a figure higher is consistent with more redenomination risk). Column (4): as Column (3) but Quanto CDS basis computed with 10 year CDS contracts. Column (5): as Column (3) but sample starts in 1st July 2011. Column (6): as Column (3) but Italy dropped from pooled sample. Column (7): as Column (3) but left hand side is given by $s_{c,d+4} - s_{c,d-1}$,where $s_{c,d}$ is the 2yr Quanto CDS basis at close on trading day d.

Table 6: Contribution of Foreign Events to Monthly Changes in Sovereign Spreads

	(1)	(2)	(3)	(4)	(5)	(6)
	Pooled	Spain	Ireland	Italy	Portugal	Pooled ex-Ireland
$m_{c,t}$	1.6374***	2.7601***	0.4448	2.1451***	2.3064**	2.3071***
	(0.497)	(0.701)	(0.768)	(0.751)	(1.013)	(0.631)
N	180	45	45	45	45	135
R^2	0.06	0.20	0.00	0.29	0.12	0.16
Robust F-stat	10.9	15.5	0.3	8.2	5.2	13.4

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,t} = \alpha + \beta m_{c,t} + w_{c,t}$. Where: (i) $\Delta s_{c,t}$ is the change in the average 2 year sovereign bond spread relative to Germany in month t; (ii) $m_{c,t}$ is the accumulation of the country c market reactions to foreign events in month t (relative to Germany and excluding events that overlap with other news). Sample period July 2009 to March 2013.

Table 7: Foreign Events and Equity Returns

	(1) Pooled	(2) Ireland	(3) Spain	(4) Italy	(5) Portugal
Bank equity returns					
$m_{c,d}$	-0.0403***	-0.0322	-0.1982***	-0.0437***	-0.0298***
	(0.010)	(0.028)	(0.073)	(0.013)	(0.011)
N	319	96	56	69	98
R^2	0.05	0.02	0.17	0.13	0.07
Robust F-stat	17.07	1.30	7.40	12.07	7.38
Non-financial equity returns					
$m_{c,d}$	-0.0367***	-0.0195*	-0.0765**	-0.1040***	-0.0209***
	(0.010)	(0.010)	(0.033)	(0.032)	(0.006)
N	319	96	56	69	98
R^2	0.05	0.04	0.12	0.09	0.09
Robust F-stat	14.80	3.53	5.53	10.73	11.03

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$,*** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}$, where: (i) $\Delta s_{c,d}$ is the log change, close to close, in the equity return index (banks, non-financials) on trading day d for country c; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. Column (1): Pooled regression on days where $|m_{c,d}| > 2bp$ such that a meaningful quantity of foreign news is observed. Column (2-5): country specific estimates.

Supplementary Appendix - For Online Publication

Appendix A provides robustness tests on the empirical results presented in in Section 3 (Appendix A.1-A.2) and Section 4 (Appendix A.3). Appendix B provides the evidence to support the claim that $m_{c,t}$ satisfies the exclusion as described in the second half of Section 3. Appendix C covers sources and manipulations of the data series used; in addition it provides an assesment of liquidity in the soveriegn debt market. Appendix D provides details of the posterior sampler used to estimate the VAR.

A Additional Results

A.1 Daily Regressions

Table A1 shows country specific estimates of Column (1) of Table 4. Spain, Portugal and Ireland all have estimates similar in magnitude to the pooled estimates. Italy is an exception, with less than 4% of the daily variation in spreads explained by foreign events. The explanation partly lies in outliers. On the 9th and 10th of November 2011 the Italian sovereign spread went through a substantial gyration, first increasing by 87bp as Silvio Berlusconi's government collapsed and then falling 91bp after President Napolitano appointed Mario Monti as a life senator paving the way for a technocratic government. Respectively, these sequential moves were both the largest daily increase and decrease in the Italian spread in the sample period and are 6.5 standard deviation daily changes in magnitude. On the same days but in non-overlapping time periods, news about a new technocratic government in Greece also emerged (with Prime Minister Papandreou first stepping down and announcing a new government on November 9th followed by former ECB official Lucas Papademos being appointed on November 10th). The Greek news still was associated with a decline in Italian sovereign spreads (-10bp on the 9th and -7bp on the 10th) but this was a small shift compared to the overall daily move. This example serves to validate a couple of points: (i) domestic events still drive a substantial portion of the variation in sovereign borrowing costs and (ii) changes in sovereign spreads have heavy tails at a daily or intradaily frequency, a relatively small number of key dates can drive the results. Removing the 9-10th November 2011 from the sample raises the coefficient estimate to a statistically significant 0.73 with an associated R^2 of 14%.

Table A2 shows the equivalent of Table 4 with $m_{c,d}$ (and $s_{c,d}$ in Columns (1)-(3)) calculated using the change in the 10 year sovereign spread. The conclusions are similar but foreign events explain a smaller share of the daily move when using the 10 year maturity.

Figure A1 shows the country specific versions of the local projection estimates presented in Figure A1. In Ireland, Italy and Spain the average event has a persistent, statistically significant, effect on sovereign spreads three weeks (15 trading days) after the event occurs.

Table A3 reconsiders the analysis in Table 5 but uses the bonds of different denominations selected by Krishnamurthy et al. (2018).[30] Specifically, redenomination risk is proxied using the spread between yields on bonds from the same issuer at similar maturities but under different denominations (USD vs EUR) and juridictions (the exact bonds are detailed in Appendix C.4). Column (1) shows the response of the difference in spreads between Italian euro denominated sovereign bonds and dollar denominated bonds, a rise in redenomination risk should raise the spread on euro vs dollar denominated debt. On impact, the point estimate suggests this is the case but the effect is statistically insignificant. Column (2) presents similar results using bonds issued by a large Italian corporate (ENI); again there is no significant effect. Columns (3)-(6) present equivalent results for Spanish and Portuguese bonds (Ireland is excluded from Krishnamurthy et al. (2018)'s sample).

In Italy and Portugal, on impact the coefficients are neither economically or statistically significant. The Spanish coefficients are very large but of a counter intuitive sign and not statistically significant. The large coefficients may reflect temporary measurement issues: looking at 5 day horizons, the Spanish coefficients have decayed to near zero.

A.2 Monthly Regressions

Table A4 presents the equivalent to the Table 6 with two alternative definitions of $m_{c,t}$. Columns (1)-(2) exclude events that occur outside the market open. As can be seen this does very little to alter the predictive power of foreign events for monthly spreads. Columns (3)-(4) show that measuring market reactions using 10 year spreads rather than 2 year; as one would expect the explanatory power is weaker as $\Delta s_{c,t}$ is still defined in using the two year maturity. In Table A5, I instead define $\Delta s_{c,t}$ as the 10 year spread. In general, foreign events explain less variation

[30] Krishnamurthy et al. (2018) also consider the spread between certain bonds and CDS prices. However, since I have already inspected Quanto CDS bases in Table 5 I have not included these in my analysis.

at a 10 year maturity and it is the case that when $m_{c,t}$ is defined using 2 year bond spreads it explains just as much of the overall monthly variation of 10 year bond spreads as when $m_{c,t}$ is calculated using 10 year bond spreads. Hence, I focus on the shorter maturity.

A.3 VAR Estimates

Figure A2 presents results from a fully pooled model when all parameters, except Γ_c, are assumed to be homogeneous across countries. The estimates are similar to the benchmark model in the main text. Instead, Figure A3 presents results from unpooled models where all four countries are estimated simultaneously. The Italian and Portuguese models align quite closely with the baseline model in the main text and have the same qualitative conclusions. The Spanish model is imprecisely estimated and it is impossible to draw conclusions. The Irish model has a pattern of responses consistent with an expansionary demand shock. However, The Irish instrument is weak and the estimates look to be biased towards the average combination of shocks that move the interest rate.

Figures A4-A7 consider the robustness of the baseline specification. Figure A4 and Figure A5 show that the main results are mostly robust to the lag order, although adding a third lag reduces the significance of the estimates. In the baseline specification, the trended series are presented as annual growth rates, this produces more precise estimates but is not ideal as it potentially introduces an MA structure into the error. Instead, in Figure A6 I present results where the trended series are converted to log first differences; this still generates a contractionary response of unemployment (and industrial production) and a tightening in private financial conditions. Last, Figure A7 shows results when $m_{c,t}$ is treated as missing prior to July 2009 rather than zero. The point estimates are identical to the baseline estimates, and there is statistically significant response of unemployment and private financial conditions. However, the tails of the posterior have widened substantially pushing out the 90% confidence intervals.

Figure A8 shows the model when I use 10 year maturity bonds in the VAR and to measure market reactions. It delivers similar impulses to the baseline. Note that the impulse responses are somewhat greater in absolute magnitude (e.g. unemployment increases by 0.7%) this picks up that a shock that raises the 10 year bond by 100bp translates to a larger shock at 2 year maturity. A9 presents results excluding overnight events. The point estimates (and 68% confidence intervals) are in line with the baseline but the tails of the posterior are fatter.

B Statistical tests on the market reactions

The critical identifying assumption is that the $m_{c,t}$ is only correlated with the structural shock of interest at time t. While it is not possible to conclusively rule out endogeneity from a statistical perspective, this section sets out to provide some evidence backing this assumption.

B.1 Are the reactions predictable?

The identification strategy used here relies on the market reaction to a foreign event reflecting a "surprise" component. Any systematic reaction to local macroeconomic shocks by foreign agents should already be anticipated by market participants at the time of the announcement. The corollary of that assumption is that the proxy variable should not be predictable and therefore should not be caused by the market reaction to past events, both domestically and in other crisis countries, or the past realisation of local macroeconomic aggregates. This can be verified empirically.

To do this, I set up a suite of univariate predictive regression models. As dependent variables I use the aggregation of market reactions to foreign events at weekly, biweekly and monthly frequencies; the weekly series is the sum of all events contained within the proxy that occur within a particular week etc. Exploring higher frequencies than monthly is necessary as market reactions within the month should also be unpredictable for the identification strategy to be justified. The predictive variables are lags of the dependent variable, lags of aggregated market reactions to local events[31] at the same frequency and in the case of the monthly model, lags of the macroeconomic time series included in the VAR.[32]

Table A6 presents F statistics of the regression and the adjusted R^2 as the outputs of interest from this analysis. In general, the message is as one would expect with rational and efficient markets: historical market moves and macroeconomic data have little predictive power over the market's reaction to current news. There is some evidence of predictive power in Portugal and Ireland at a monthly frequency but this is only in one specification and the statistical significance is only at the 10% level. The overall result supports the identifying assumptions.

[31] Local events that overlap with data releases, ECB meetings or pan-European policy interventions are omitted.

[32] With the model with the VAR data I use two lags. Otherwise, the lag orders for the various models are determined automatically by selecting the order that minimises the Bayesian Information Criterion up to a maximum of order of three months or equivalent.

B.2 Are the the reactions to events a function of other macroeconomic news?

A second empirical test is to gauge the extent to which the reactions to foreign events are related to changes in local economic, monetary or fiscal conditions by testing to see if $m_{c,t}$ is correlated with the market reaction to local economic and fiscal data releases. This could be thought of as a test of whether foreign events are uncorrelated to local macroeconomic shocks that are being captured by a data surprise. One can also consider the correlation between the proxy and the market reaction to ECB announcements - a potential proxy for monetary shocks. This is still an imperfect test as the causality could run in the other direction; for example, events that raise yields may lower confidence and cause negative survey releases or provoke an ECB reaction.

For ECB meetings the market reaction is considered from 12:30pm-14:50pm on the day of the meeting to capture both the interest rate announcement and the press conference. The timing of the release of local economic data in each country is obtained from the *Bloomberg* economic calendar; the market reaction in terms of the local spread is considered in a twenty minute window about the release to be consistent with the main analysis. For comparative purposes data releases are grouped into three categories to distinguish between their content: *(1)* Output releases correspond to industrial production, various confidence surveys and unemployment data; *(2)* Inflation releases are the consumer and producer price releases; *(3)* Fiscal releases: correspond to monthly data on government finances from a cash accounting basis.

Data releases and ECB meetings are at a monthly frequency so I do not consider weekly and biweekly aggregations in this case. Furthermore, the use of a monthly measure of data reactions means that releases that occur on a quarterly basis cannot be considered for consistency reasons. Releases are grouped by the month they are released rather than the month they refer to. The market reactions are then aggregated for the month and the correlation with the foreign events. See section C.6 for a detailed discussion of the included data releases.

There is no general pattern of market reactions to local data being correlated with $m_{c,t}$; this is true as well for ECB announcements. This suggests little systematic correlation between market reactions and that participants are internalising information when forming expectations as one would hope given the identification strategy.[33] There is a weak correlation be-

[33] One may also be concerned that the market reaction to events are not just related to economic conditions locally or in other crisis countries but also to economic shocks in creditor countries. Economic conditions in Germany could determine how willing the country is to lend to crisis countries and this could determine behaviour of agents in those countries. However, in an additional study, not presented, I also show that the proxy is uncorrelated with

tween the reactions to ECB meetings and foreign events in Italy. The sample is short so this correlation could be spurious. More importantly, the correlation is positive yet the shock identified in the VAR associates a tightening in spreads with lower risk free rates and an economic contraction; this is inconsistent with a positive monetary policy shock.

B.3 Are the market reactions to local and foreign events correlated?

I also consider the contemporaneous correlation between market reactions to local and foreign events. Again, given the identifying assumptions one should not see any regular patterns between these reactions when aggregated. At a monthly frequency the correlations are: -0.07 in Italy, -0.13 in Spain, -0.19 in Portugal and 0.13 in Ireland. None of these figures are statistically distinguishable from zero.

B.4 To what extent are market reactions due to real or financial linkages?

The last empirical test considered is the extent to which the bond market reactions to events are explained by real or financial linkages between countries. If it was the case that the market reactions were largely explained by such linkages then the concern would be that $m_{c,t}$ would be correlated either with external demand shocks or shocks to the domestic financial system caused by losses abroad. To explore this issue formally, I consider a regression model that attempts to explain the relative market reaction to the identified events using trade and financial linkages.

Data on nominal trade flows (in euros) between countries is taken on a seasonally adjusted basis from Eurostat's monthly trade statistics and are summed into a quarterly series. Data on financial linkages are sourced from the Bank of International Settlements international banking statistics on a consolidated, immediate borrower basis.[34] I consider the stock of claims of domestically owned banks on counterparties in other countries at the end of the quarter. The raw data is in dollars, I convert it to Euros using the end of quarter exchange rate as quoted by the BIS.

To describe the regression formally: let the index e denote events, the index c be over reaction countries and the index k be over event countries, q denotes the time period in which

German data surprises as well, which argues against this effect.

[34] The ideal would be to use data on an ultimate risk basis. Unfortunately, for a number of country pairs in this sample the data on this basis is not available.

the event occurs - quarters, in this context. Given these definitions, let $|m^e_{c,k,q}|$ denote the absolute market reaction to country *c*'s bond yield to event e in country k that occurred in quarter q. Define the term $Trade_{c,k,q}$ as exports from country c to country k in quarter q as a percentage of country c GDP. And the term $Financial_{c,k,q}$ as the stock assets of banks in country c hold that have a counterparty in country k in quarter q as a percentage of country c GDP. The regression model can then be written as:

$$|m^e_{c,k,q}| = \alpha^1_e + \alpha^2_{c,k} + \beta_1 Trade_{c,k,q} + \beta_2 Financial_{c,k,q} + u^e_{c,k,q} \tag{A1}$$

Where α^1_e is the fixed effect for an individual event invariant across reaction countries and $\alpha^2_{c,k}$ is a fixed effect for an event-country reaction-country pair and is invariant across events. The remaining term, $u^e_{c,k,q}$, is a stochastic disturbance term. Given this specification, the *relative* within country-pair variation in trade and financial linkages is being used to estimate β_1 and β_2.

The initial benchmark specification is defined as follows: The reaction countries are the four considered in the main model. The set of event countries is necessarily limited to the 6 included in the proxy data. I then consider all observations where the reaction country does not have an overlapping local event, pan-European event or a data release. As with the benchmark proxy, "non-headline" overnight events are not considered and nor are local reactions to local events (i.e. the case where $c = k$). Market reactions are considered in relation to a 20 minute window on either side of the time the event occurred.

Alongside this benchmark I consider a number of different variations of the specification expressed in equation (A1). The second specification expands the set of reaction countries to include other Eurozone countries that did not suffer from elevated sovereign yields during the crisis (specifically: Austria, Belgium, Germany, France, and the Netherlands).[35] The third specification lags real and financial linkages by one quarter - i.e. the financial linkage is measured as claims at the start of the quarter. The goal here is to avert potential reverse causality between events and linkages: it could be that a particularly severe event causes linkages between countries to diminish, biasing down the estimates of β_1 and β_2. By looking at the lagged linkage we can think of the values as predetermined as well as better reflecting

[35]My dataset does not time local events in these additional countries; hence, I can only omit observations when the event overlaps with a pan-European event or a data release.

the information set of market participants. The fourth specification expresses the linkages (as a share of GDP) in logarithmic terms.

The results for these regressions are presented in Table A8. The estimates suggest that neither trade nor financial linkages have a statistically meaningful relationship with the market reaction to events. Furthermore, the size of the point estimates suggests little economic significance. For example, consider the relationship between financial linkages and bond market reactions. In the benchmark specification, an additional 1% of GDP worth of financial linkages increases the absolute market reaction by 0.14bp: less than a twentieth of the average market reaction to an event of 4.8bp. This coefficient seems quite robust to using lagged linkages, although increasing the set of reaction countries diminishes it. Note also that 1% of GDP is a large shift in terms of relative financial exposures. For example, in Q1 2013, the range between the most exposed Eurozone country to Greece (Germany) and the least exposed (Italy) was 0.7% of GDP. In absolute terms, on average over the sample, Italian, Irish, Portuguese and Spanish exposures to Greece averaged 0.14%, 1.17%, 1.00% and 0.07% of GDP respectively.

The findings here are at odds with Brutti and Saure (2015) who find that financial linkages explained a substantial portion of the market reactions to Greek events in 2010. The methodology in their paper are similar to those employed here, the main difference being that Brutti and Saure (2015) interact their measures of linkages with a shock identified via a proxy SVAR where as this paper uses the high frequency market reaction directly. However, the larger difference is in the sample, this paper has a broader set of events over a longer time frame covering more event countries. This may explain the discrepancy.

C Data Appendix

C.1 Sources and construction of the intraday data

The intraday data is sourced from Thomson Reuters Datascope. The Thomson Reuters tickers (RICs) for the benchmark two year sovereign bond for the four countries in the sample are *ES2YT=RR*, *IE2YT=RR*, *IT2YT=RR* and *PT2YT=RR*. The ticker for Germany is DE2YT=RR; other countries' data can be obtained by altering the first two letters. The tickers for the benchmark 10 year bond are equivalent with 2 replaced with 10 in the code. I download the data at minutely intervals and calculate the mid-yield in the minute as the average of the open ask, open bid, close ask and close bid, where the open and close refer to the first and last observation recorded by Thomson Reuters in the interval. To smooth out the impact of any unusual spikes I then take centered 5 medians of the data for the purposes of calculating market reactions.

C.2 Features of the intraday data

To get some sense of the liquidity in the markets that provide the intraday day data, the upper panel Figure A10, presents the average number of quotes per minute recorded in Thomson Reuters Datascope for the benchmark 2 year sovereign bond of Ireland, Italy, Portugal and Spain between 07:30 and 17:30 London time in trading days over the period 2009-2013. Market open hours between 08:00 and 16:30 are visible. The Italian and Spanish bonds are clearly more liquid than Portuguese or Irish bonds on this basis. The liquidity also evolves differently during the crisis period. The lower panel in Figure A10 presents the average number of quotes per minute at 14:00 across the years in the sample. While the recorded number of quotes went up in Italy and Spain (which, I conjecture, likely represents an improvement in data collection), activity drops off in Portugal and Ireland as the crisis becomes more intense in 2011-2012 before recovering again in 2013.

These trends are reflected in Bid-Ask spreads (see Figure A11). The benchmark 2-year Italian and Spanish bonds averaged substantially lower bid-ask spreads (at around 8bp and 13bp) than the Irish or Portuguese or equivalents (at around 65bp and 75bp). This disguises substantially heterogeneity through time during the sample period. Irish bid-ask spreads were actually lower on average than on the Italian bond in 2009, but peaked at 140bp at the most intense phase of the crisis in Ireland in 2011. A similar pattern exists in Portugal. While Italian and Spanish bonds also experience higher bid-ask spreads in 2011-2012. The higher bid-ask spreads in Ireland and Portugal also reflect that mid-yields were much higher. For example, Portuguese mid-yields averaged 12.5% in 2011 so the bid-ask spread was just over a tenth of the mid-yield. The Italian mid-yield averaged 3.8% in 2011 so the bid-ask spread was about a twentieth the size. Relative to overall borrowing costs, there is less difference between bid-ask spreads across countries.

Nonetheless the less liquid Portuguese and Irish bonds do have have more volatile intraday yields. This is shown in Figure A12 where the average absolute change in yield on a minute-by-minute basis over trading days is presented. This pattern is not picked up in the time series dimension however (lower panel in Figure A12). Despite liquidity worsening for Portugal and Ireland, realised intraday volatility was on a downward trend over the crisis period. This is promising as it suggests that worsening liquidity was not introducing dramatically more noise into the high frequency market reactions.

A couple of further points are in order. First, note that the period 08:00-08:30 has somewhat higher realised volatility and bid-ask spreads than the rest of the day, this is what motivates extending the window to 08:30 when measuring the market's reaction to overnight events. Second, by way of external validation: note that the intraday data also coincides with more commonly used daily data. For instance, the 2-year mid bond yield at close (16:30) from the intraday data is closely correlated in all four countries with the daily yield extracted from Datastream (to some extent this is unsurprising since the source is also Reuters).

C.3 An example of event timing

To understand the technical steps involving an event, here I present a brief example of the Portuguese parliament's rejection of the government's proposed austerity package on the 23rd of March 2011. This is illustrative as it is an event that potentially could last for an extended period and it was known in advance that the vote would take place so there are preview news stories to eliminate. Figure A13 presents an unformatted screenshot for the Bloomberg news output for Portuguese news in the immediate vicinity of the time of the vote. As can be seen there are multiple news reports listed. The question is how to determine the first relevant headline and when the event ended. I am only interested in Bloomberg newswire headlines not articles or news reports from other sources. Such headlines can be identified as they begin with an asterisk and are in all capitals, with the source listed as "BN". Key headlines are marked in red. As can be seen in figure A13 there are five newswire headlines displayed - the remainder of the news stories can be ignored for the moment. The headlines descend in reverse chronological order, i.e. the most recent headline is listed first. Headline number 11) is a newswire headline that refers to the vote but as it is only a preview statement from a policymaker this can be disregarded. Headline number 6) signals the start of the vote and has a time stamp (not shown) at 9:51pm London time; headline 4) confirms the result of the vote and was printed at 9:55pm London time. Headlines 2) and 3) repeat the information in 4) and can be disregarded. I therefore conclude that the vote finished at 9:55pm London time and lasted less than 20 minutes so that time is taken as the event announcement time. If the vote had lasted until 10:11pm I would then consider the event start time at 9:51pm with the end time 20 minutes later. However, in this context the exact event timing is moot since it occurs outside the market hours and the reaction would then be considered from the market open on the 24th.

What would have needed to have occurred for this event to be considered untimeable? The simplest case is if there are no Bloomberg newswire headlines associated with the event on the country news page. A second case would be if the set of Bloomberg headlines associated with the event lasted for more than 90 minutes (e.g if 6) and 4) where 90 minutes apart) and the time stamp on the headlines was during the market open. A third case, which happens rarely in the dataset, is when an article relating to the result of the event has been published in a non-newswire news report prior to the newswire headline about the event. For example, news story 1) in figure A13 is an article about the details of the vote published after the vote had concluded. However, if such an article was published with a timestamp that was earlier than that of article 4) then there is an inconsistency and I would not consider the event timeable.

C.4 Daily Financial Market Data:

Sovereign CDS Prices These are sourced from Datastream. For Euro CDS the relevant codes is *??GXEAC* where *??* is the two digit country code (Germany is *DE* etc) and *X* is the maturity in yield ("*A*" refers to ten years). The equivalent codes for dollar CDS are *??GX$AC* with equivalent *X* and *??*. These prices are for full restructed senior credit derivative swaps.

Benchmark Sovereign Bonds For the longer VAR sample the sovereign bond price is constructed using the the average of daily data from Datastream with yields expressed as spreads to Germany. The relevant tickers are: *Italy: TRIT2YT; Spain: TRES2YT; Ireland: TRIE2YT; Portugal: TRPT2YT, Germany: DEPT2YT.* These series are all very tightly correlated (~0.99) with the equivalent series from Datascope at close. Different maturies are obtained by altering the number that is the fifth character in the ticker: e.g *TRIE3YT* is the 3 year.

Ireland lacks a benchmark 2 year sovereign bond between 23/10/2006 and 04/03/2009 (note that this is irrelevant for the construction of $m_{c,t}$, which starts in July 2009). To address this I approximate the yield using the daily change in the following synthetic bond yields: (i) the yield on the 3 year benchmark Irish bond less the spread between yields on the German 3 and 2 year benchmark bond (25/10/07-24/10/06). (ii) the yield on the 5 year benchmark Irish bond less the spread between yields on the German 5 and 2 year benchmark bond (25/10/07-26/10/07). (iii) the yield on the 5 year benchmark Irish bond less the spread between yields on the German 5 and 2 year benchmark bond less the spread between 5 and 2 year Irish USD senior CDS (05/11/08-03/03/08). (iv) the yield on the 3 year benchmark Irish bond less the spread between yields on the German 3 and 2 year benchmark bond less the spread between 3 and 2 year Irish USD senior CDS (03/03/09-06/11/08).

Dollar and Euro Sovereign Bonds from Krishnamurthy et al. (2018): For the exercise in Appendix A.1, I use the following bonds and CDS, all data is sourced from Datastream unless otherwise stated:

- *Italy:* Euro sovereign bond is BTP ITALY 2005 3 3/4% 01/08/15 (ticker: 498629); Dollar sovereign bond is ITALY 2005 4 1/2% 21/01/15 (yield sourced from Bloomberg); USD corporate bond is ENI FINANCE INTL. 2008 6 1/8% 17/12/18 (ticker: 3642Q0). Corporate EUR CDS is ENI S.P.A. SNR CR14 ?Y (ticker: ENI?EAC) where ? refers to the maturity.

- *Spain:* Euro sovereign bond is OBLIGACION ESTADO 2003 4.2% 30/07/13 (ticker: 24427H); Dollar sovereign bond is SPAIN 2008 3 5/8% 17/06/13 (ticker: 2159N0); USD corporate bond is TELF.EMISIONES SAU 2006 6.421% 20/06/16 (ticker: 72675K). Corporate EUR CDS is TELEFONICA SA SNR CR14 ?Y (ticker: TEF?EAC) where ? refers to the maturity.

- *Portugal:* Euro sovereign bond is PORTUGAL 2005 3.35% 15/10/15 (ticker: 56319R); Dollar sovereign bond is PORTUGAL 2010 3 1/2% 25/03/15 (ticker: 5595T3); GBP corporate bond is EDP-ENERGIAS DE PT SNR CR14 1Y (ticker: 22146U). Corporate CDS is EDP-ENERGIAS DE PT SNR CR14 ?Y (ticker: EDP?EAC) where ? refers to the maturity.

I match the bonds to the swap rate of the nearest maturity when calculating spreads.

Equity Return Indices I use the broad Datastream equity return indices for Banks (ticker: BANKS??(RI)) and non-financials (ticker: TOTLI??(RI)) where ?? refers to a two digit country code (Portugal: PT, Ireland: IR, Italy: IT, Spain: ES). When including the indices in the VAR, I take averages over the trading days in the month and then calculate log changes.

C.5 Other VAR data sources:

Unemployment: The harmonised unemployment rates are sourced from Eurostat and expressed as a percent of the labour force (*Eurostat code: une_rt_m*).

Industrial Production: The industrial production index is sourced from Eurostat. The broadest index possible is used, including the manufacturing, energy and construction sectors (*Eurostat code: sts_inpr_m*). For Ireland, the measure excludes construction.

Consumer Prices: The harmonised index of consumer prices (HICP) is sourced from Eurostat. The core index is used - all items excluding food and energy (*Eurostat code: prc_hicp_midx*).

Private Sector Cost of Finance: This is computed internally by the Capital Markets/Financial Structure division of the ECB for each country in the Euro Area. It is the amalgamation of the cost of loans to the non-financial private sector, the cost of corporate bonds and the cost of equity (the latter two apply to non-financial corporations only). The cost of the three sources of finance are weighted using flows of new liability acquisition by non-financial private sector. This creates an average cost of finance faced by the private sector analogous to an overall interest rate on financial liabilities. The cost can be decomposed into its constituent components as in Figure 10. The cost equity is not available consistently throughout the sample so equity prices are used instead.

Primary Fiscal Balances: This is the most complex input into the VAR. As no official monthly data for fiscal balances exists on an accruals basis, one is constructed using interpolation methods. Since fiscal numbers are available on a cash accounting basis at monthly frequency, these series serve as natural interpolands. The quarterly primary fiscal balance is defined as the net lending/borrowing of the general government sector plus interest payments. This is sourced from the Eurostat flow of funds database; the fiscal balance is created using the non-financial accounts (*Eurostat code: nasq_nf_tr*). Flow of funds data are in millions of nominal euros and are not seasonally adjusted. The unadjusted balance as a percentage of GDP is calculated by dividing through by quarterly, nominal GDP from Eurostat in millions of Euros (*Eurostat code: namq_gdp_c*). The adjusted quarterly balance is created by placing this data through an X.12 filter. Monthly nominal GDP is constructed by linearly interpolating the quarterly series in each country; since monthly GDP is the relatively stable denominator in the monthly fiscal series this choice of interpolation technique is of little importance. The interpolation procedure for the fiscal balance is conducted in percentage of GDP terms using the regression based procedure in Mitchell et al. (2005). The interpolation is regression estimated using maximum likelihood; it is assumed the underlying fiscal balance is an ARX(1,1) on a monthly basis restricted such that the sum of the monthly balances equal the quarterly figure. Experiments with alternative lag structures revealed little sensitivity to alternative specifications. The differences across countries in the availability of monthly fiscal data across countries mean that the interpolands and sample periods are country specific:

- **Italy:** The first interpoland is monthly the central government balance less central government interest payments (both millions of Euros, calculated on a cash accounting basis and non-seasonally adjusted). The second interpoland is the change in general government debt (millions of Euros, non-seasonally adjusted). Both interpolands are divided through by monthly nominal GDP and seasonally adjusted using an X.12 procedure. Both series are sourced from the Italian Finance Ministry. The sample period for the estimation is January 2000 to March 2013. The model is extended beyond the sample for the VAR to improve the quality of the fit.

- **Spain:** The first interpoland is monthly the central primary government balance (in millions of Euros, calculated on a accruals basis and non-seasonally adjusted). The second interpoland is the monthly change in central government gross debt outstanding (millions of Euros, non-seasonally adjusted). Both interpolands are divided through by monthly nominal GDP and seasonally adjusted using an X.12 procedure. Both series are sourced from the Spanish Finance Ministry. The sample period for the estimation is January 1999 to March 2013.

- **Portugal:** The first interpoland is monthly the central government balance (in millions of Euros, calculated on a cash accounting basis and non-seasonally adjusted). The second interpoland is the change in general government debt (millions of Euros, non-seasonally adjusted). Both interpolands are divided through by monthly nominal GDP and seasonally adjusted using an X.12 procedure. Both series are sourced from the Portuguese Finance Ministry. The sample period for the estimation is January 2000 to March 2013.

- **Ireland:** There is a single interpoland which is monthly: the Exchequer surplus, equivalent to the central government balance, (in millions of Euros, calculated on a cash accounting basis and non-seasonally adjusted). The interpolands are divided through by monthly nominal GDP and seasonally adjusted using an X.12 procedure. The series is sourced from the Irish Finance Ministry. The sample period for the estimation is January 2000 to March 2013.

The interpolation procedure appears to work well, there are no unusually large spikes in the monthly series and the interpolated figures do not resemble the output from a deterministic interpolation procedure, suggesting the monthly interpolands are informative.

Credit Volumes: I use total loans on balance sheet of MFIs to domestic, non-financial corporations and households. This data is sourced from the ECB's database on MFI balance sheet statistics. I seasonally adjust the data using an X.12 procedure and then convert it into real terms by dividing by the country specific CPI index.

Target 2 Balance: Data for target 2 balances are sourced from the updated dataset of Steinkamp and Westermann (2012). The data is in millions of euros. The balance is converted into percentage of GDP terms by dividing through by nominal GDP linearly interpolated.

C.6 Data releases

Data releases serve two purposes in this paper. First, events which overlap with a twenty minute window about local data releases are excluded from the proxy. Second, the reaction of the market to data releases is aggregated for each month and compared to the reaction about events as a robust check. Here, data releases considered are listed. For the purposes of table A7, those marked with a 1 are used as output releases, those as 2 are inflation releases and those marked as 3 are fiscal releases - the sum of three corresponds to the all data column. Note that only series released monthly are included in this analysis (which is why GDP is not used for example). The first release is always used rather than the final revised number. Descriptions here correspond to those listed on the Bloomberg Economic Calendar.

- **Italian Data Releases:** Budget Balance (3), Business Confidence (1), Consumer Confidence (1), CPI Final, CPI Preliminary (2), Current Account, Deficit to GDP, GDP final, GDP Preliminary, General Government Debt, Hourly Wages, Industrial Orders (1), Industrial

Production (1), Industrial Sales (1), Labor Costs, New Car Registrations, PMI Manufacturing (1), PMI Services (1), PPI (2), Retail Sales (1), Trade Balance , Unemployment Rate (1).

- **Spanish Data Releases:** CPI Final, CPI Preliminary (2), Current Account, GDP final, GDP Preliminary, House Price Index, House transactions (1), Industrial Output (1), Labour Costs, Mortgages on Houses, Producer Prices (2), Retail Sales Volumes (1), Spain Budget Balance (3), Spain Business Confidence, Spain Consumer Confidence (1), Spain Manufacturing PMI (1), Spain Services PMI (1), Total Housing Permits, Trade Balance, Unemployment (1), Unemployment.

- **Portuguese Data Releases:** Construction Works Index, Consumer Confidence, Consumer Price Index, Current Account, Economic Climate Indicator, GDP (YoY) final, GDP Preliminary, Industrial Production (1), Industrial sales (1), Labour Costs, Producer Prices, Retail Sales, Trade Balance, Unemployment Rate.

- **Irish Data Releases:** Consumer Confidence (1), CPI (2), Current Account Balance, GDP, Industrial Production (1), Live Register Level, Manufacturing PMI (1), New Vehicle Licenses (1), PPI (2), Property Prices, Retail Sales Volumes (1), Services PMI (1), Trade Balance, Unemployment Rate.

The following important international and European data releases are also used to exclude overlapping events from the proxy (admittedly this an arbitrary selection):

- **International Releases:** Eurozone Services PMI, Eurozone Manufacturing PMI, German IFO, US Labour Market (non-Farm Payrolls), European Commission Confidence Surveys, Eurozone GDP Final, Eurozone GDP Preliminary.

D Details of the VAR Model and the posterior Sampler

D.1 Modeling the Identification Equation

The variable $m_{c,t}$ constructed as described in Section 2 presents several issues from an econometric perspective. The variable is an aggregation of stochastic high frequency bond market reactions. Events are not continuously observed and certain types of news are omitted. Thus the proxy is best described as the aggregation of censored, random observations. The market reaction is also an imperfect gauge of the informational content of an event. Market-specific factors such as liquidity or large transactions can result in a noisy signal. The informational content of an event can be difficult to process quickly and the market reaction may be slowed by lags in the decision-making of investors. Rumours may also leak in advance, attenuating the response. Thus, there will be measurement error contained in the observed reaction to each event. However, there may be also scaling effects: the initial market reaction may over- or under-shoot in a regular fashion.

Despite these empirical issues, it is possible, by making a few assumptions, to motivate the equation (5) in the main text. The true underlying data-generating process is continuous. However, rather than combine a discrete time VAR and continuous events, I approximate the event-generating process by modeling it on a daily basis (conceptually this could be extended by narrowing the time window). As before let d denote days and t denote months and define $\mathcal{D}_t$ the set days in month t. For notational convenience the country subscript, c, is dropped for the moment. Let $m_{d,t}$ be the recorded market response on day d of month t. I assume this has the following data-generating process:

$$m_{d,t} = \vartheta_{d,t}(\psi\varepsilon_{d,t} + v_{d,t}) \tag{A2}$$

The variable $\vartheta_{d,t}$ is an indicator for censoring, taking a value of one or zero. If an event is not observed on a particular day then $\vartheta_{d,t}$ (and $m_{d,t}$) take a value of zero. If an event occurs, then the observed market reaction is assumed to be the sum of the scalar structural shock that occurs in that time window, $\varepsilon_{d,t}$, scaled by parameter ψ, and some independent measurement error $v_{d,t}$ ($v_{d,t} \sim IID(0, \sigma_v^2)$). I follow the special case in Mertens and Ravn (2013) and assume that the censoring process is random; that is to say $\vartheta_{d,t}$ is an independent variable that takes a value of 1 with probability p and zero otherwise. I assume that the daily series of scalar structural shocks sum to create the monthly shock of interest:

$$\varepsilon_t = \sum_{d \in \mathcal{D}_t} \varepsilon_{d,t}$$

I retain the assumption that the underlying structural shocks are Gaussian so that the daily structural shocks have the property: $\varepsilon_{d,t} \sim NID(0, |\mathcal{D}_t|^{-1})$ such that the monthly shock is Gaussian with unit variance. If one interprets $\varepsilon_{d,t}$ as a structural shock to the bond yield, this aggregation assumption is equivalent to the yield following a process close to a random walk at a high frequency. The data supports this. However, as I describe, $m_{d,t}$ is heavy tailed. To capture this, I assume that $v_{d,t}$ is drawn from a student t distribution with degree of freedom parameter η. I calibrate η such that the excess kurtosis of $m_{d,t}$ conditional on $\vartheta_{d,t} = 1$ matches the cross-country average excess kurtosis of the market reaction in event windows (7.4) assuming $\sigma_v^2 \times |\mathcal{D}_t| = 1$ such that the implicit signal to noise ratio is a half. The choice of a half follows the high relevance prior in Caldara and Herbst (Forthcoming); however, the final calibrated value of ν below is insensitive to this as the conditional kurtosis of $m_{d,t}$ is what matters and that is the target variable. Based on these assumptions, I select $\eta = 5$ to the nearest integer.

The monthly proxy is the sum of the observed market reactions:

$$m_t = \sum_{d \in \mathcal{D}_t} m_{d,t}$$

The exogeneity assumption requires that both $\varepsilon_{d,t}$ and $v_{d,t}$ are uncorrelated with any other structural shocks that hit the economy in month t. These assumptions are sufficient for the relationship between the m_t and the structural shock to satisfy the exclusion restriction. The assumption that $v_{d,t}$ is uncorrelated with other structural shocks has two justifications. First, time windows where important information about other shocks are revealed (e.g. data releases) are excluded. Second, that the time windows are sufficiently small such that any underlying correlation between yields and the state of the economy tends to zero. Furthermore, I assume that $\mathbb{E}(m_{d,t}m_{d+s,t}) = 0$ and $\mathbb{E}(m_t m_{t+s}) = 0 \; \forall s \in \mathbb{N}$ which is backed by the predictability regressions in Appendix B.

Under this setup, it is possible to rewrite the relationship between $m_{c,t}$ and the reduced form residuals, u_t as:

$$m_t = \Upsilon' u_t + \omega_t \tag{A3}$$

where $\Upsilon' = p\psi a_1$ such that $\phi = p\psi$. The distribution of ω is symmetric and zero in expectation but is also heavy tailed due to both the Studentised distribution of $v_{d,t}$ and the censoring process. This motivates the Student-t prior in Section 4.1: $m_t | u_t, \Upsilon, \sigma_w^2; \nu \sim t(\Upsilon' u_t, \sigma_\omega^2; \nu)$. It is straight forward to verify that $\sigma_{c,\omega}^2 = (\frac{\nu-2}{\nu}) p |\mathcal{D}| \sigma_{c,v}^2$ where $|D|$ is the number of days in the average month.

As the data entering the model is at a monthly frequency, jointly estimating the censoring process adds an unnecessary layer of complexity. Instead, p is set deterministically and calibrated to the proportion of trading days in the sample where $m_{d,t}$ is observed. Similarly $|D|$ is deterministic and is set to 30. The extent of the excess kurtosis that is a function of only p, $|D|$, η and the signal noise ratio in equation (A2). Given these values, I simulate the ω in equation A3 for arbitrary σ_v^2 and ψ. I then calculate ν by matching the fourth moment. Plugging in the numbers from above results in $\nu = 11$ to the nearest integer.

The identification equation can be estimated for a subset of the full sample period: Let $1 \leq t_1$ be the earliest date for which the $m_{c,t}$ available and $t_2 \leq T$ be the latest available date with $T_m = t_2 - t_1 + 1$. Letting $\mathcal{M}_c$ be the matrix of observations on $m_{c,t}$ stacked over time and U_c^m be the matrix of residuals for $t \in \{t_1, \ldots, t_2\}$, one can approximate the conditional distribution of the proxy as:

$$\mathcal{M}_c | U_c^m, \Upsilon_c, \sigma^2_{c,w}; \nu \sim t(U_c^m \Upsilon_c, \sigma^2_\omega I_{T_m}; \nu) \tag{A4}$$

Where $t(.,.)$ denotes a multivariate scaled Student's-t distribution with a scalar degrees of freedom parameter ν. Last, Define the residuals from the proxy model as: $V_c = (\mathcal{M}_c - U_c^m \Upsilon_c)$.

D.2 Estimation

To generalise notation I rewrite the priors on λ_β and λ_Υ as

$$\lambda_\beta | \varrho, \varsigma \sim IG_2 \propto \lambda_\beta^{\frac{-\varsigma+2}{2}} exp\{-\frac{1}{2}\frac{\varrho}{\lambda_\beta}\} \tag{A5}$$

$$\lambda_\Upsilon | \varrho, \varsigma \sim IG_2 \propto \lambda_\Upsilon^{\frac{-\varsigma+2}{2}} exp\{-\frac{1}{2}\frac{\varrho}{\lambda_\Upsilon}\} \tag{A6}$$

where $\varrho = 0$ and $\varsigma = -1$ are the shape and scale parameters in the inverse Gamma distribution. In addition, let $\kappa = N + 2$ denote the degrees of freedom parameter from the inverse-Wishart prior in equation (4).

An advantage of working with the $\mathcal{A}$ matrix is that the joint likelihood of the data is hierarchical and straightforward to define. To simplify the notation define the parameter space in the model (Θ), the set of data used in the reduced form VAR (Y) and the external instruments variables as ($\mathcal{M}$):

$$\Theta = \{\beta_1, \ldots, \beta_C, \Sigma_{1,u}, \ldots, \Sigma_{C,u}, \gamma_1, \ldots, \gamma_C, \bar{\beta}, \lambda_\beta, \bar{S}, \Upsilon_1, \ldots, \Upsilon_C, \sigma_{1,\omega}, \ldots, \sigma_{C,\omega}, \bar{\Upsilon}, \lambda_\Upsilon\},$$

$$Y = \{Y_1, \ldots, Y_C, X_1, \ldots, X_C, Z_1, \ldots, Z_C\}$$

$$\mathcal{M} = \{\mathcal{M}_1, \ldots, \mathcal{M}_C\}.$$

By Bayes rule, the joint likelihood of the data is equal is $p(\mathcal{M}, Y|\Theta) = p(\mathcal{M}|Y, \Theta)p(Y|\Theta) = \prod_c p(\mathcal{M}_c|Y_c, \Theta)p(Y_c|\Theta)$. The form of $p(Y_c|\Theta)$ is given in equation (3). The condition density of the proxy, $p(\mathcal{M}_c|Y_c, \Theta)$, is defined in equation (A4). As is well known (see Geweke (1993)), by expanding the parameter space it is possible to rewrite the conditional density as a Gaussian regression model with heteroskedastic errors:

$$\mathcal{M}_c|Y_c \sim N(U_c^m \Upsilon_c, \sigma_{\omega,c}^2 \Xi_c)$$

Where the matrix Ξ_c is a diagonal vector of unknown parameters equal to $diag\{\xi_{c,t_1}, \ldots, \xi_{c,t_2}\}$. With the additional prior assumption $\nu/\xi_{c,i} \sim \chi^2(\nu)\ \forall i = t_1, \ldots, t_2$, where ν are the degrees of freedom on the student-t errors. Therefore $p(\mathcal{M}_c|Y_c, \Theta)$ is Gaussian and from above $p(Y_c|\Theta)$ is Gaussian; hence, the joint density, $p(\mathcal{M}_c, Y_c|\Theta)$, is also Gaussian.

For estimation, the unconditional densities of the parameters cannot be determined analytically; hence they are computed numerically using Markov Chain Monte Carlo methods. The functional forms of the priors lead to a set of conditional posterior distributions that are standard; this motivates the use of a Gibbs Sampler to construct the posteriors.

An important feature of using the joint likelihood $p(\mathcal{M}, Y|\Theta)$ is that the estimates of the VAR parameters, β_c, are conditioned on the proxy and take into account the relationship between $\mathcal{M}$ and Y. The conditional posterior density of β_c contains a term that acts to reduce the probability that a draw of the reduced form slope coefficients will produce reduced form residuals that have a weak relationship with the proxy. The reduced form estimates therefore incorporate the information contained in the proxy despite its omission from the reduced form model. This is in contrast to the frequentist proxy SVAR estimation strategy.

This gives:

$$p(\mathcal{M}_c|Y_c,\Theta)=\sigma_{c,\omega}^{-T_m-1}\prod_{t=t_1}^{t_2}\xi_{c,t}^{-1/2}exp\left[-\sum_{t=t_1}^{T_m}\frac{(m_{c,t}-\Upsilon_c u_{c,t})^2}{2\xi_{c,t}\sigma_{c,\omega}^2}\right]=\sigma_{c,\omega}^{-T_m-1}|\Xi_c|^{-\frac{1}{2}}exp\left\{-\frac{1}{2}\left\{\sigma_{c,\omega}^{-2}V_c'\Xi_c^{-1}V_c\right\}\right\}$$

The likelihood of the reduced form VAR model, $p(Y|\Theta)$, is the product of the country specific Gaussian distributions as defined in the main text. Combining these two densities with the priors gives a joint posterior density, $p(\mathcal{M},Y|\Theta)p(\Theta)$, proportional to:

$$\begin{aligned}&|\bar{S}|^{\frac{C\kappa-(N+1)}{2}}(\lambda_\Upsilon\lambda_\beta)^{-\frac{\varsigma+2+C}{2}}\prod_c\left(\sigma_{c,\omega}^{-T_m-1}|\Xi_c|^{-\frac{1}{2}}|\Sigma_{c,u}|^{-\frac{T+\kappa+N+1}{2}}\right)exp\left\{\frac{\varrho}{\lambda_\beta}+\frac{\varrho}{\lambda_\Upsilon}\right\}\dots\\&exp\left\{-\frac{1}{2}\left(\sum_c\left\{tr\left[(U_c'U_c\Sigma_{c,u}^{-1})+\bar{S}\Sigma_{c,u}^{-1}\right]+(\beta_c-\bar{\beta})'(\lambda_\beta L_{c,\beta})^{-1}(\beta_c-\bar{\beta})\right\}\right)\right\}\dots\\&exp\left\{-\frac{1}{2}\left(\sum_c\left\{(\Upsilon_c-\bar{\Upsilon})'(\lambda_\Upsilon L_{c,\Upsilon})^{-1}(\Upsilon_c-\bar{\Upsilon})+\sigma_{c,\omega}^{-2}V_c'\Xi_c^{-1}V_c-\nu tr(\Xi_c^{-1})\right\}\right)\right\}\end{aligned}$$

This is a convenient way to express the posterior. However, it is also apparent that the VAR data and the proxy are jointly Gaussian:

$$\begin{pmatrix} y_c \\ \mathcal{M}_c \end{pmatrix}|\Theta\sim N\begin{pmatrix} (I_N\otimes X_c)\beta_c+(I_N\otimes Z_c)\gamma_c & , \begin{bmatrix}\Phi_{c,11} & \Phi_{c,12}\\ \Phi_{c,12}' & \Phi_{c,22}\end{bmatrix} \\ 0 & \end{pmatrix}$$

where

$$\Phi_{c,11}=\Sigma_{c,u}\otimes I_T$$

$$\Phi_{c,22}=\Upsilon_c'\Sigma_{c,u}\Upsilon_c\otimes I_{T_m}+\sigma_{\omega,c}^2\Xi_c$$

$$\Phi_{c,12}=\Sigma_{c,u}\Upsilon_c\otimes\begin{pmatrix}0_{(t_1-1)\times T_m}\\ I_{T^m}\\ 0_{(T-t_2)\times T_m}\end{pmatrix}$$

For all the parameters in the model the conditional densities used in the Gibbs Sampler are in the form of classical distributions. The conditional density of the slope coefficients is:

$$p(\beta_c|Y,\mathcal{M},\Theta\setminus\beta_c) = exp\left\{-\frac{1}{2}\left(\left\{tr\left[(U_c'U_c\Sigma_{c,u}^{-1})\right] + (\beta_c-\bar{\beta})'(\lambda_\beta L_{c,\beta})^{-1}(\beta_c-\bar{\beta}) + \sigma_{c\omega}^{-2}V_c'\Xi_c^{-1}V_c\right\}\right)\right\}$$

Using the joint Gaussian density of the proxy and the reduced form one can show:

$$p(\beta_c|Y,\mathcal{M},\Theta\setminus\beta_c) \propto N(D_c^{-1}d_c, D_c^{-1}) \tag{A7}$$

where

$$D_c = (I_N\otimes X_c)'(\Phi_{c,11}-\Phi_{c,12}\Phi_{c,22}^{-1}\Phi_{c,22})^{-1}(I_N\otimes X_c) + \lambda_\beta^{-1}L_{c,\beta}^{-1}$$

$$d_c = (I_N\otimes X_c)'(\Phi_{c,11}-\Phi_{c,12}\Phi_{c,22}^{-1}\Phi_{c,21})^{-1}(y_c-(I_N\otimes Z_c)\gamma_c-\Phi_{c,12}\Phi_{c,22}^{-1}\mathcal{M}_c) + \lambda_\beta^{-1}L_{c,\beta}^{-1}\bar{\beta}$$

The coefficients on the deterministic terms in the reduced form VAR have a similar form, the conditional density is given by:

$$p(\gamma_c|Y,\mathcal{M},\Theta\setminus\gamma_c) = exp\left\{-\frac{1}{2}\left(\left\{tr\left[(U_c'U_c\Sigma_{c,u}^{-1})\right] + \sigma_{c,\omega}^{-2}V_c'\Xi_c^{-1}V_c\right\}\right)\right\}$$

This is also Gaussian:

$$p(\gamma_c|Y,\mathcal{M},\Theta\setminus\gamma_c) \propto N(F_c^{-1}f_c, F_c^{-1})$$

where

$$F_c = (I_N\otimes Z_c)'(\Phi_{c,11}-\Phi_{c,12}\Phi_{c,22}^{-1}\Phi_{c,21})^{-1}(I_N\otimes Z_c)$$

$$f_c = (I_N\otimes Z_c)'(\Phi_{c,11}-\Phi_{c,12}\Phi_{c,22}^{-1}\Phi_{c,21})^{-1}(y_c-(I_N\otimes X_c)\beta_c-\Phi_{c,12}\Phi_{c,22}^{-1}\mathcal{M}_c)$$

The conditional posterior of Σ_c is proportional to:

$$p(\Sigma_{c,u}|Y,\Theta\setminus\Sigma_{c,u}) \propto |\Sigma_{c,u}|^{-\frac{T+\kappa+N+1}{2}} exp\{-\frac{1}{2}tr\left[(U_c'U_c)+\bar{S}\right]\Sigma_{c,u}^{-1}\}$$

which is consistent with an inverse-Wishart distribution:

$$p(\Sigma_c|Y,\Theta_1 \setminus \Sigma_{c,u}) \propto iW((U_c^{'}U_c)+\bar{S}, T+\kappa)$$

In terms of the cross-country hyper-parameters, $\bar{\beta}$, has a conditional posterior proportional to a Normal:

$$p(\bar{\beta}|Y,\Theta \setminus \bar{\beta}) \propto N(\left[\sum_c G_c\right]^{-1}\left[\sum_c g_c\right], \left[\sum_c G_c\right]^{-1})$$

$$G_c = (\lambda_\beta L_{c,\beta})^{-1}$$

$$g_c = (\lambda_\beta L_{c,\beta})^{-1}\beta_c$$

The conditional posterior of $\bar{S}$ is proportional to:

$$p(\bar{S}|Y,\Theta \setminus \bar{S}) \propto |\bar{S}|^{\frac{C\kappa-N-1}{2}} exp\{-\frac{1}{2}tr\bar{S}\left[\sum_c \Sigma_{c,u}^{-1}\right]\}$$

which corresponds to a Wishart distribution:

$$p(\bar{S}|Y,\Theta \setminus \bar{S}) \propto W\left(\left[\sum_c \Sigma_{c,u}^{-1}\right]^{-1}, C\kappa\right)$$

Note that $E(\bar{S}|Y,\Theta \setminus \bar{S}) = C\kappa(\sum_c\left[\Sigma_{c,u}^{-1}\right])^{-1}$. This implies that the expected value of $\bar{S}$ is the harmonic mean of the individual country covariance matrices scaled by the degrees of freedom parameter κ. This is used to determine the covariance of the cross-country model, $\bar{\Sigma}$, for use in impulses etc. By setting $\bar{\Sigma} = \bar{S}/\kappa$, one obtains a matrix that is analogous to a covariance matrix and in (conditional) expectation is equivalent to the harmonic mean of the estimated country covariances.

The conditional posterior for the shrinkage parameter, λ_1, is proportional to:

$$p(\lambda_\beta|Y,\Theta \setminus \lambda_\beta) \propto \lambda_\beta^{-\frac{CN^2L+\varsigma+2}{2}} exp\left\{-\frac{1}{2}\left(\frac{\varrho}{\lambda_\beta}+\sum_c\left[(\beta_c-\bar{\beta})'\lambda_\beta^{-1}L_{c,\beta}^{-1}(\beta_c-\bar{\beta})\right]\right)\right\}$$

or

$$p(\lambda_\beta|Y,\Theta\setminus\lambda_\beta) = iG_2\left(\varrho + \sum_c\left[(\beta_c-\bar{\beta})'L_{c,\beta}^{-1}(\beta_c-\bar{\beta})\right], CN^2L+\varsigma\right)$$

Where iG_2 refers to an inverted Gamma-2 distribution. For computational convenience, it is easier to draw from the posterior distribution of the inverse of λ_β which is easily shown to be proportional to a standard Gamma distribution.

In terms of the identification parameters, the slope terms have the following conditional densities:

$$p(\Upsilon_c|Y,\Theta\setminus\Upsilon_c) \propto N(Q_c^{-1}q_c, Q_c^{-1})$$

where:

$$Q_c = \sigma_{\omega,c}^{-2}(U_c^m)'\Xi_c^{-1}U_c^m + \lambda_\Upsilon^{-1}L_{c,\Upsilon}^{-1}$$

$$q_c = \sigma_{\omega,c}^{-2}(U_c^m)'\Xi_c^{-1}\mathcal{M}_c + \lambda_\Upsilon^{-1}L_{c,\Upsilon}^{-1}\bar{\Upsilon}$$

And the conditional posterior of $\sigma_{\omega,c}^2$ is proportional to:

$$p(\sigma_{\omega,c}^2|Y,\Theta\setminus\sigma_{\omega,c}^2) \propto \sigma_{c\omega}^{-T^m-1}exp\{-\frac{1}{2}\left[(V_c'\Xi_c^{-1}V_c)\right]\sigma_{\omega,c}^{-2}\}$$

which is consistent with an inverse-Gamma distribution:

$$p(\sigma_{\omega,c}^2|Y,\Theta\setminus\sigma_{\omega,c}^2) \propto iG((V_c'\Xi_c^{-1}V_c), T^m)$$

The conditional posterior of $\xi_{c,t}$, the diagonals in Ξ_c, can be expressed as:

$$p(\xi_{ct}|Y,\Theta) \propto \xi_{c,t}^{-(\nu+3)/2}exp\left[-(\sigma_{\omega,c}^{-2}(m_{c,t}-\Upsilon' u_{c,t})+\nu)/2\xi_{c,t}\right]$$

Which is consistent with each diagonal element, ξ_{ct}, being related to the inverse of a χ^2, specifically:

$$p((\sigma_{\omega c}^{-2}(m_{ct} - \Upsilon' u_{ct}) + \nu)/\xi_{ct}|Y, \Theta) \propto \chi^2(\nu + 1)$$

In terms of the cross-country hyper-parameters, $\bar{\Upsilon}$ has a conditional posterior proportional to a Normal:

$$p(\bar{\Upsilon}|Y, \Theta_1 \setminus \bar{\Upsilon}) \propto N(\left[\sum_c R_c\right]^{-1} \left[\sum_c r_c\right], \left[\sum_c R_c\right]^{-1})$$

$$R_c = (\lambda_\Upsilon L_{c,\Upsilon})^{-1}$$

$$r_c = (\lambda_\Upsilon L_{c,\Upsilon})^{-1} \Upsilon_c$$

Last, the posterior of λ_Υ is proportional to:

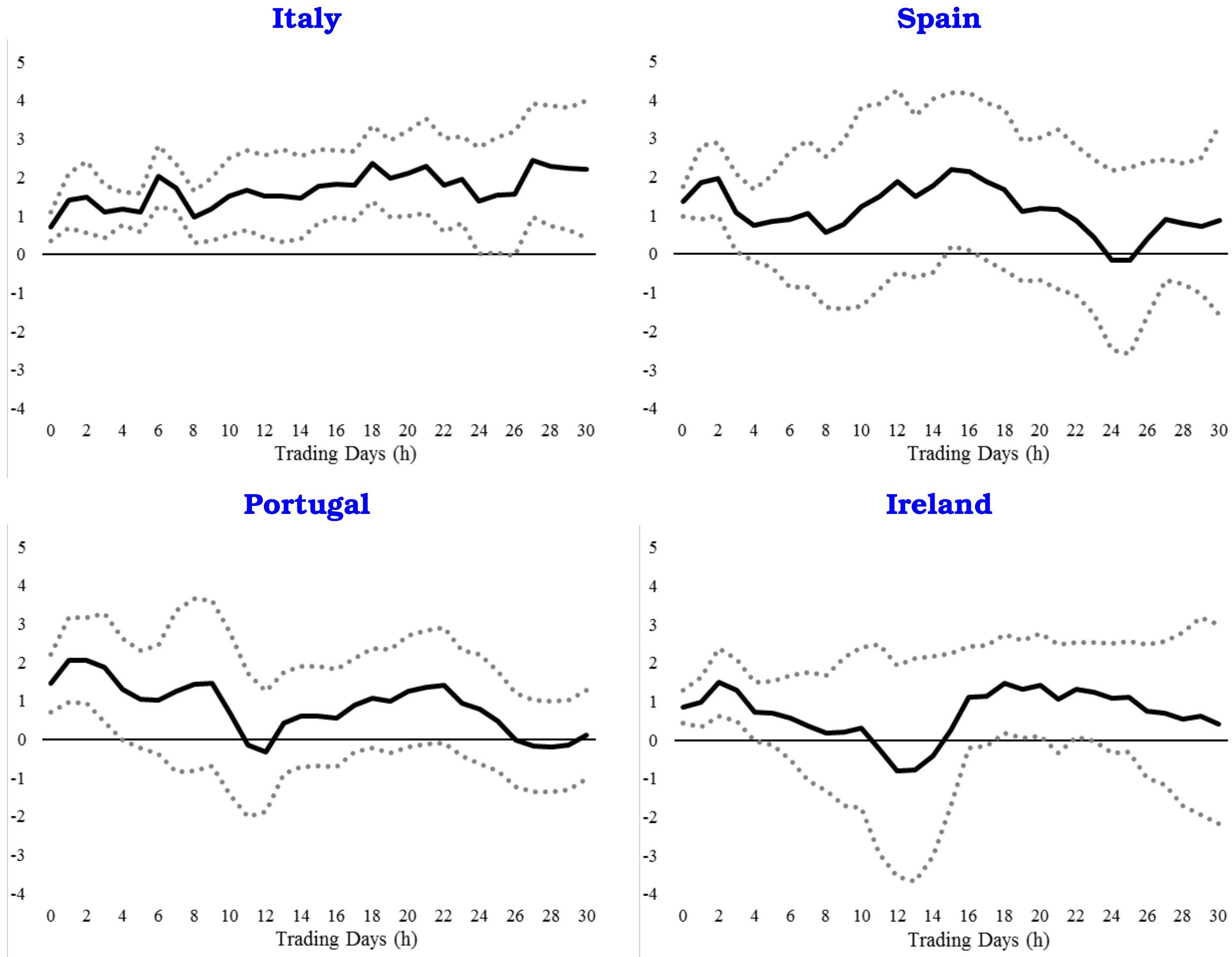
$$p(\lambda_\Upsilon|Y, \Theta \setminus \lambda_\Upsilon) = iG_2\left(\varrho + \sum_c \left[(\Upsilon_c - \bar{\Upsilon})'(L_{c,\Upsilon})^{-1}(\Upsilon_c - \bar{\Upsilon})\right], CN + \varsigma\right)$$

Where iG_2 refers to an inverted Gamma-2 distribution.

For the baseline specification the posterior is simulated using 1,500,000 draws from the MCMC sampler; the first 500,000 are discarded as a burn-in and the remaining chain is thinned by a factor of 100 leaving 10,000 draws for inference. Results presented are the median of the 10,000 retained draws and credible intervals are computed using standard Bayesian Monte-Carlo methods. This length of chain ensures the convergence diagnostics on λ_Υ and λ_β are passed. However, a shorter chain leads only to imperceptible differences in results. Hence, for other specifications used for robustness checks or extensions I use a shorter chain of length 150,000 where the first 50,000 are discarded as burn-in with the remaining draws thinned by a factor of 100 leaving 1000 draws for inference.

Tables and Figures

Figure A1: Local Projection Estimates of the Persistence of Foreign Events: Country Specific Models



Notes: Impulse response to the following local projection: $s_{c,d+h} - s_{c,d-1} = \alpha_c + \beta_c^h m_{c,d} + w_{c,d+h}$, where: (i) $s_{c,d}$ is the 2 year sovereign bond spread at 16:30 London time on day d for country c; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Country specific estimates for reaction countries Spain, Ireland, Italy and Portugal. Sample period 1st July 2009 to 31st of March 2013. The y-axis is the estimates of β^h and can be interpreted as a cumulative percent change in the spread following a foreign event that provokes a 100bps market reaction on day 0; x-axis is the horizon h. Grey dashed lines are 90% confidence intervals constructed from Newey-West standard errors with a 30 day lag window.

Table A1: Contribution of Foreign Events to Daily Sovereign Spread: Country Specific Regressions

	(1)	(2)	(3)	(4)
Country c:	Spain	Ireland	Italy	Portugal
m_{cd}	1.5245***	1.0280***	0.6425	1.7792***
	(0.231)	(0.305)	(0.538)	(0.565)
N	44	67	43	65
r2	0.37	0.17	0.05	0.40

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}$, where: (i) $\Delta s_{c,d}$ is the change in the 2 year sovereign bond yield between 08:30 and 16:30 London time on trading day d for country c relative to the change in the German 2 year yield; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. Columns (1)-(4) country specific estimates for reaction countries Spain, Ireland, Italy and Portugal.

Table A2: Contribution of Foreign Events to Daily Sovereign Spread: 10 year market reactions

	(1) Open-Close, event days	(2) Close-Close, event days	(3) All Days	(4) Two Year
$m_{c,d}$	1.2516***	0.854***	0.914***	1.751***
	(0.399)	(0.190)	(0.267)	(0.36)
N	183	303	3908	183
R^2	0.17	0.16	0.03	0.09
Robust F-stat	9.9	20.3	15.0	5.3

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,d} = \alpha + \beta m_{c,d} + w_{c,d}$, where: (i) $\Delta s_{c,d}$ is the change in the 10 year sovereign bond yield between 08:30 and 16:30 London time on trading day d for country c relative to the change in the German 10 year yield; (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (relative to Germany and excluding events that overlap with other news). Sample period 1st July 2009 to 31st of March 2013. Column (1): Pooled regression on days where $|m_{c,d}| > 2bp$ such that a meaningful quantity of foreign news is observed. Column (2): includes the overnight time period such that $\Delta s_{c,d}$ is calculated between 16:30 $d-1$ and 16:30 on trading day d; similarly $m_{c,d}$ is computed including events that happen overnight. Column (3): includes all trading days with $m_{c,d}$ set as zero when no news is observed. Column (4): equivalent to Column (1) with $\Delta s_{c,d}$ measured as the change in the 2 year sovereign spread.

Table A3: Proxies for redenomination risk using the bonds in Krishnamurthy et al. (2018)

	(1) Italy Sovereign USD spread less EUR spread	(2) Italy Corporate USD spread less EUR CDS	(3) Spain Sovereign USD spread less EUR spread	(4) Spain Corporate USD spread less EUR CDS	(5) Portugal Sovereign USD spread less EUR spread	(6) Portugal Corporate GBP spread less EUR CDS
Impact						
$m_{c,d}$	0.1203	0.0820	-1.0492	-0.5076	0.1394	0.0248
	(0.172)	(0.112)	(1.050)	(0.319)	(0.263)	(0.120)
N	68	69	56	56	90	98
R^2	0.01	0.00	0.01	0.04	0.00	0.00
5 day horizon						
$m_{c,d}$	0.3357	0.0106	-0.1847	-0.1127	0.4683	0.0970
	(0.489)	(0.336)	(0.851)	(0.483)	(0.509)	(0.349)
N	68	69	56	56	90	98
R^2	0.01	0.00	0.00	0.00	0.01	0.00

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, ***$p < 0.01$

Notes: Estimates for β_c^h from the following regression: $s_{c,d+h} - s_{c,d-1} = \alpha_c + \beta_c^h m_{c,d} + w_{c,d+h}$, where: (i) impact denotes $h = 0$ and 5 day horizon denotes $h = 4$ and (ii) $m_{c,d}$ is the accumulation of the country c market reactions to foreign events on trading day d (2 year benchmark bond yield, relative to Germany and excluding events that overlap with other news). Column (1) $s_{c,d}$ is the difference between the spread (to swaps) on the Italian Sovereign Euro and Dollar denominated bond. Column (2) $s_{c,d}$ is the difference between the spread (to swaps) on Italian Corporate (ENI) Dollar denominated bond and ENI EUR CDS. Column (3) $s_{c,d}$ is the difference between spread (to swaps) on the Spanish Sovereign Euro and Dollar denominated bond. Column (4) $s_{c,d}$ is the spread between the yield on Portuguese Sovereign Euro and Dollar denominated bond. Column (5) $s_{c,d}$ is the difference between the spread (to swaps) on Spanish Corporate (Telefonica) Dollar denominated bond and Telefonica EUR CDS. Column (6) $s_{c,d}$ is the difference between the spread (to swaps) on Portuguese Corporate (EDP) Sterling denominated bonds and EDP EUR CDS. In all cases, c refers to the country that issued the bonds used to construct $s_{c,d}$. Sample restricted to days where $|m_{c,d}| > 2bp$.

Table A4: Contribution of Foreign Events to Monthly Change in 2 Year Sovereign Spreads: Alternative measures of $m_{c,t}$

	(1)	(2)	(3)	(4)
	Excluding Overnight Events		Using 10Yr Spreads	
	Pooled	Pooled ex-Ireland	Pooled	Pooled ex-Ireland
$m_{c,t}$	2.2730**	3.0409***	0.9510	0.9811
	(0.881)	(1.005)	(0.857)	(0.636)
N	180	135	180	135
R^2	0.07	0.17	0.01	0.02
Robust F-stat	6.66	9.16	1.23	2.38

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,t} = \alpha + \beta m_{c,t} + w_{c,t}$. Where: (i) $\Delta s_{c,t}$ is the change in the average 10 year sovereign bond spread relative to Germany in month t; (ii) $m_{c,t}$ is the accumulation of the country c market reactions to foreign events in month t (relative to Germany and excluding events that overlap with other news). Sample period July 2009 to 31st of March 2013. Column (1)-(2): $m_{c,t}$ measured using 2 year spreads but excludes events outside the market open. Column (3)-(4): $m_{c,t}$ measured using 10 year spreads.

Table A5: Contribution of Foreign Events to Monthly Change in 10 Year Sovereign Spreads

	(1)	(2)	(3)	(4)
	2 Year Market Reaction		10 Year Market Reaction	
	Pooled	Pooled ex-Ireland	Pooled	Pooled ex-Ireland
$m_{c,t}$	0.5524**	0.8102**	0.9385*	0.9364**
	(0.279)	(0.341)	(0.481)	(0.373)
N	180	135	180	135
R^2	0.03	0.07	0.04	0.05
Robust F-stat	3.92	5.65	3.81	6.31

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Notes: Estimates from equations $\Delta s_{c,t} = \alpha + \beta m_{c,t} + w_{c,t}$. Where: (i) $\Delta s_{c,t}$ is the change in the average 10 year sovereign bond spread relative to Germany in month t; (ii) $m_{c,t}$ is the accumulation of the country c market reactions to foreign events in month t (20mins before vs 20mins after, relative to Germany and excluding events that overlap with other news). Sample period July 2009 to 31st of March 2013. Column (1-2): $m_{c,t}$ measured using 2 year spreads. Column (3-4): $m_{c,t}$ measured using 10 year spreads.

Figure A2: Mean Country Impulse Responses to a Sovereign Spread Shock: Fully Pooled Model

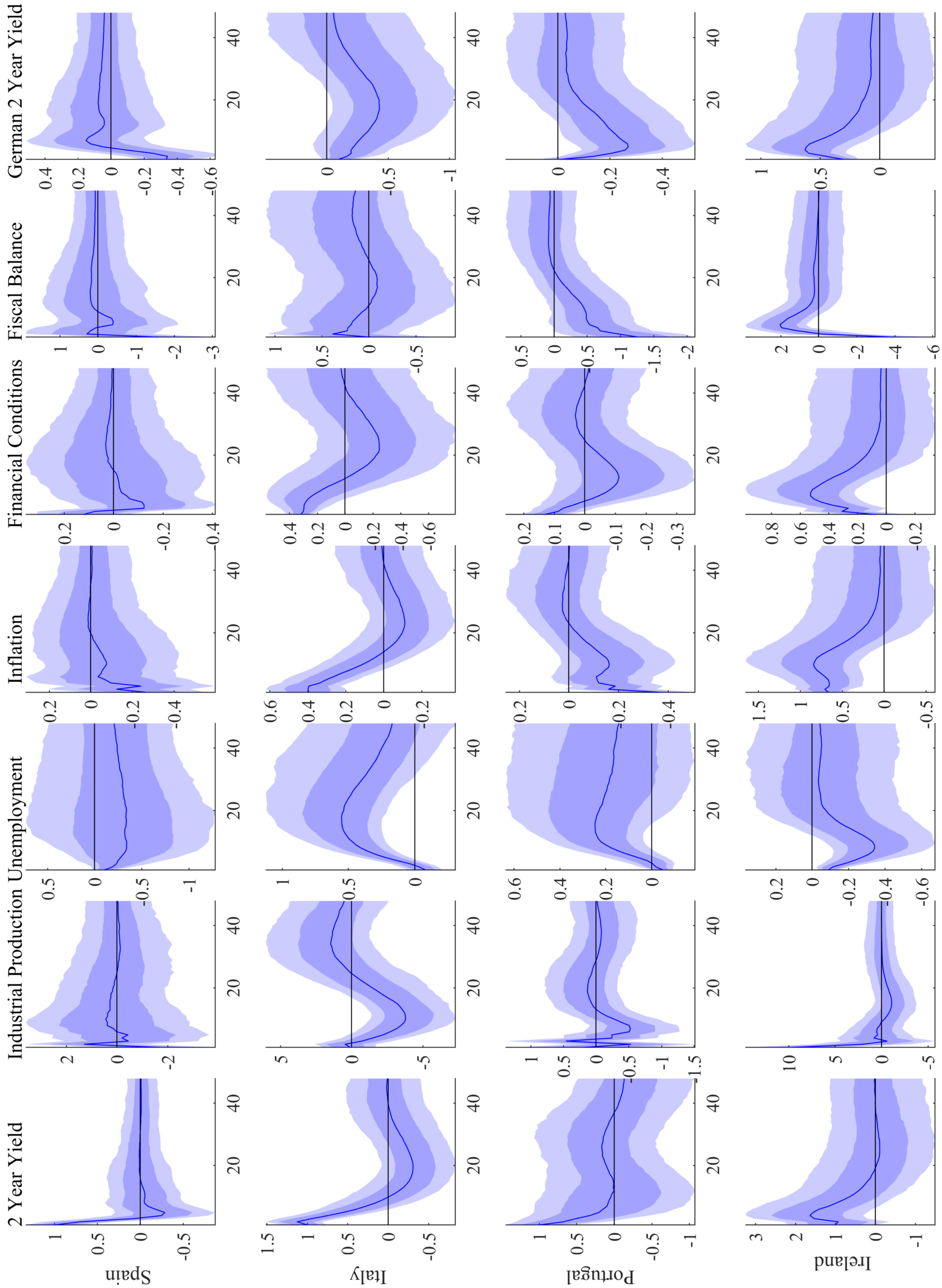
Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

Table A6: Predictability of the proxy series

	Weekly Event Data	Biweekly Event Data	Monthly Event Data	Monthly Event Data + VAR data
Country	F-stat from Regression			
Italy	0.30	0.02	0.16	0.67
	(0.74)	(0.98)	(0.85)	(0.80)
Portugal	2.56**	2.76*	0.89	1.99*
	(0.04)	(0.07)	(0.42)	(0.06)
Spain	0.26	0.16	0.55	1.20
	(0.77)	(0.85)	(0.58)	(0.34)
Ireland	0.58	1.74	1.35	2.02*
	(0.56)	(0.18)	(0.27)	(0.05)
	Adj R^2			
Italy	0.00	0.00	0.00	0.0
Portugal	0.03	0.04	0.00	0.30
Spain	0.00	0.00	0.00	0.08
Ireland	0.00	0.02	0.02	0.30

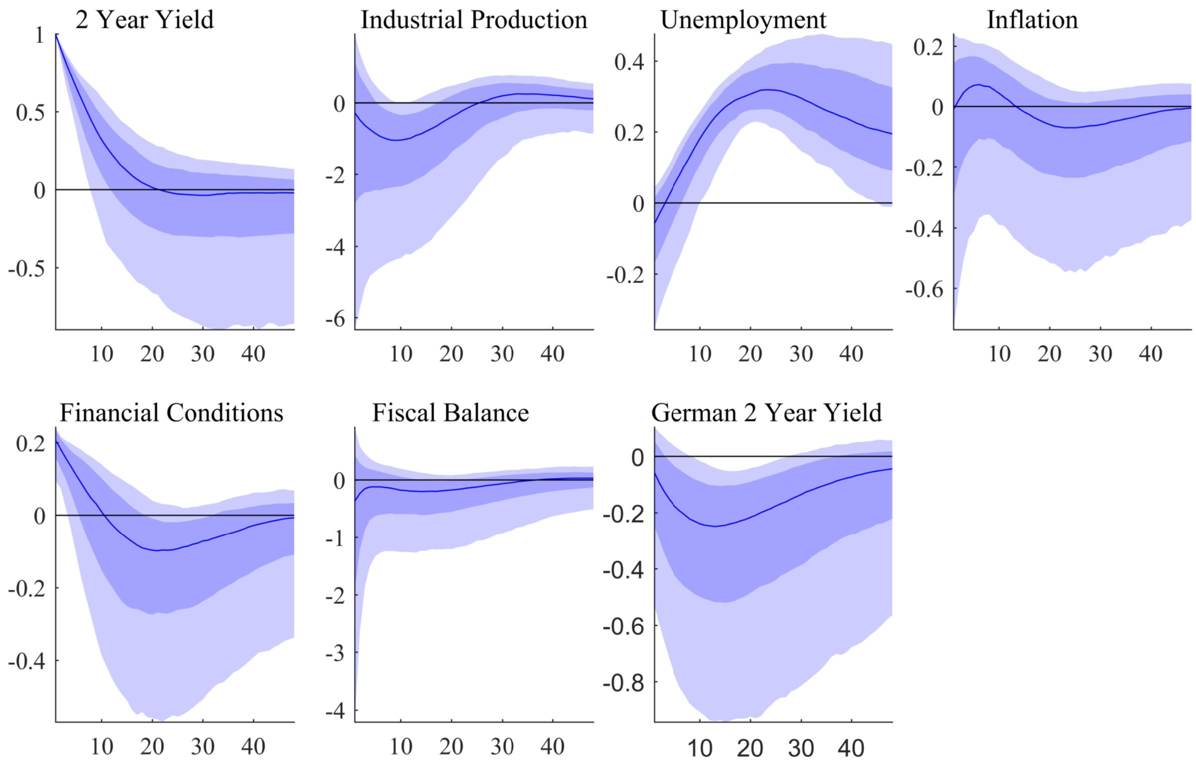
Notes: Regression statistics from univariate predictive regressions of the market reactions to foreign events aggregated at different frequencies: weekly, biweekly, monthly. Sample Period: 1st July 2009 - 31st March 2013. Explanatory variables include a constant, lags of proxy variable and aggregated market reactions to local events. Lag order selected by the Bayesian Information Criterion on a country and frequency specific basis. Maximum lag order set to 12 weeks/6 fortnights/3 months. Monthly model with VAR data includes explanatory variables from the reduced from the benchmark VAR model in section in the main text. with two lags P-values in parenthesis, *** Denotes significance at 1% level, ** 5%, *10%. Negative adjusted R^2 are normalised to zero.

Figure A3: Country Specific Impulse Responses to a Sovereign Spread Shock: Unpooled Model



Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; for exact data definitions see the data appendix. Centre line is the median of draws from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

Figure A4: Mean Country Impulse Responses to a Sovereign Spread Shock: 1 Lag



Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

Table A7: Correlations between market reaction to events and data news

	Correlation with Foreign Events				
	All Data	Output	Inflation	Fiscal	ECB Meetings
Italy	-0.12	-0.25	0.28*	0.18	0.34***
Portugal	-0.09	0.07	-0.12	-	-0.02
Spain	0.02	-0.04	0.17	0.11	0.05
Ireland	0.19	0.09	0.25	-	-0.14

* $p < 0.1$, ** $p < 0.05$,*** $p < 0.01$

Notes Sample correlation coefficients between market reactions in 20 minute windows aggregated into monthly series about events, data and ECB meetings. Foreign events refer to the proxy series as described in the main text; market reactions to local events are aggregated in a similar fashion excluding events that overlap with data releases, ECB meetings or pan-European policy interventions are omitted. Data releases organised by relevant month. Output: IP, Confidence Surveys, PMI's, Unemployment. Inflation: CPI and PPI, Fiscal: Monthly fiscal data and Government Debt (where applicable). Sample Period: July 2009 - March 2013.

Figure A5: Mean Country Impulse Responses to a Sovereign Spread Shock: 3 Lags

Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

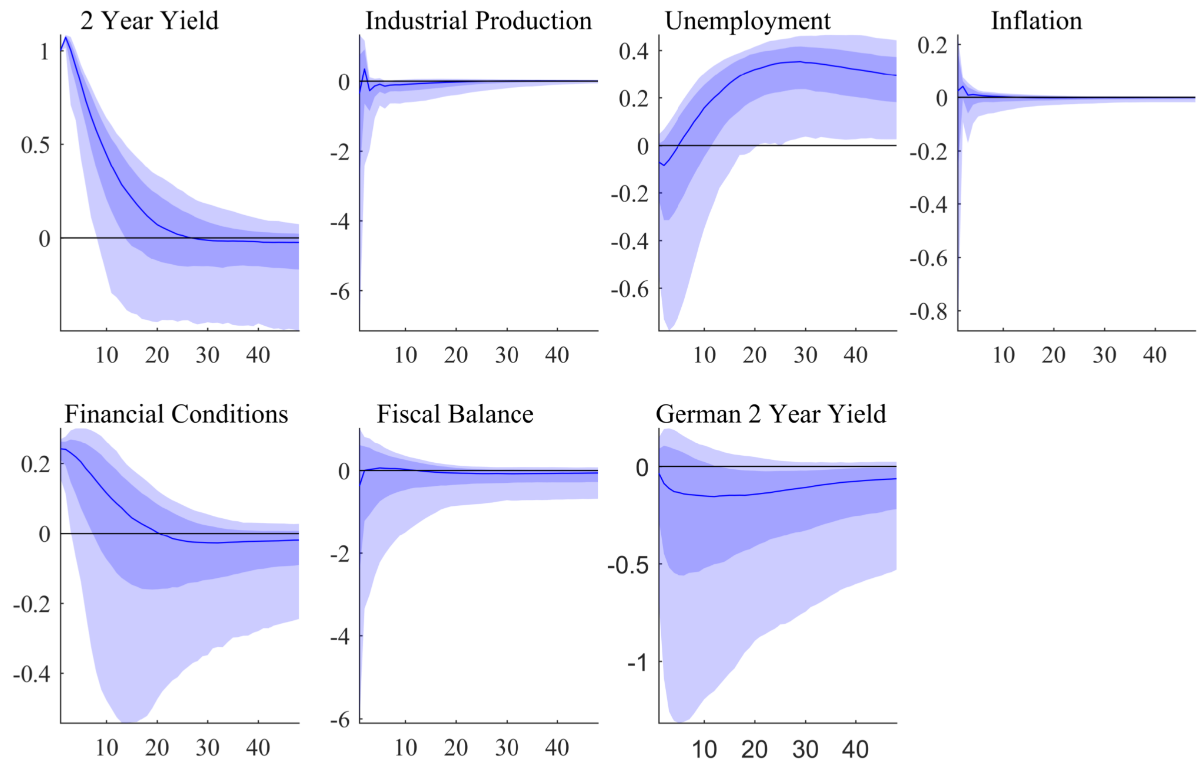
Table A8: Estimates of the effect of trade and financial linkages on the relative market reactions to events

	Benchmark	incl. non-crisis	lag interlinkages	log interlinkages
	(1)	(2)	(3)	(4)
Trade	2.707	0.668	2.728	-0.652
	(2.541)	(1.465)	(2.391)	(1.926)
Financial	0.141	0.027	0.188	0.387
	(0.166)	(0.097)	(0.162)	(0.819)
Observations	763	1958	763	763
R^2	0.503	0.373	0.503	0.502

Cluster robust standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$,*** $p < 0.01$

Notes: Results from panel regressions of absolute local market reactions to foreign events on real and financial linkages between the local and foreign economy. All specifications include event and reaction country/event country pair fixed effects. The identified events are in Greece, Cyprus, Portugal, Spain, Ireland and Italy as determined during the construction of the proxy. Reaction countries are Spain, Ireland, Italy and Portugal. Overlapping events, "non-headline" overnight events and local reactions to local events are not included. Coefficient estimates correspond to change in market reaction in basis points for every 1% change in GDP of linkages. *Financial* is reaction country bank claims on the event country as a percentage of reaction country GDP. *Trade* is reaction country exports to the event country as a percentage of reaction country GDP. The different specifications are as follows *(1):* Benchmark specification. *(2)* Reaction countries expanded to include non-crisis countries (Austria, Belgium, Germany, France, Netherlands). *(3)* Trade and financial linkages are lagged by one period. *(4)* Real and financial linkages expressed in logarithimic terms.

Figure A6: Mean Country Impulse Responses to a Sovereign Spread Shock: Trended Series in First Differences



Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

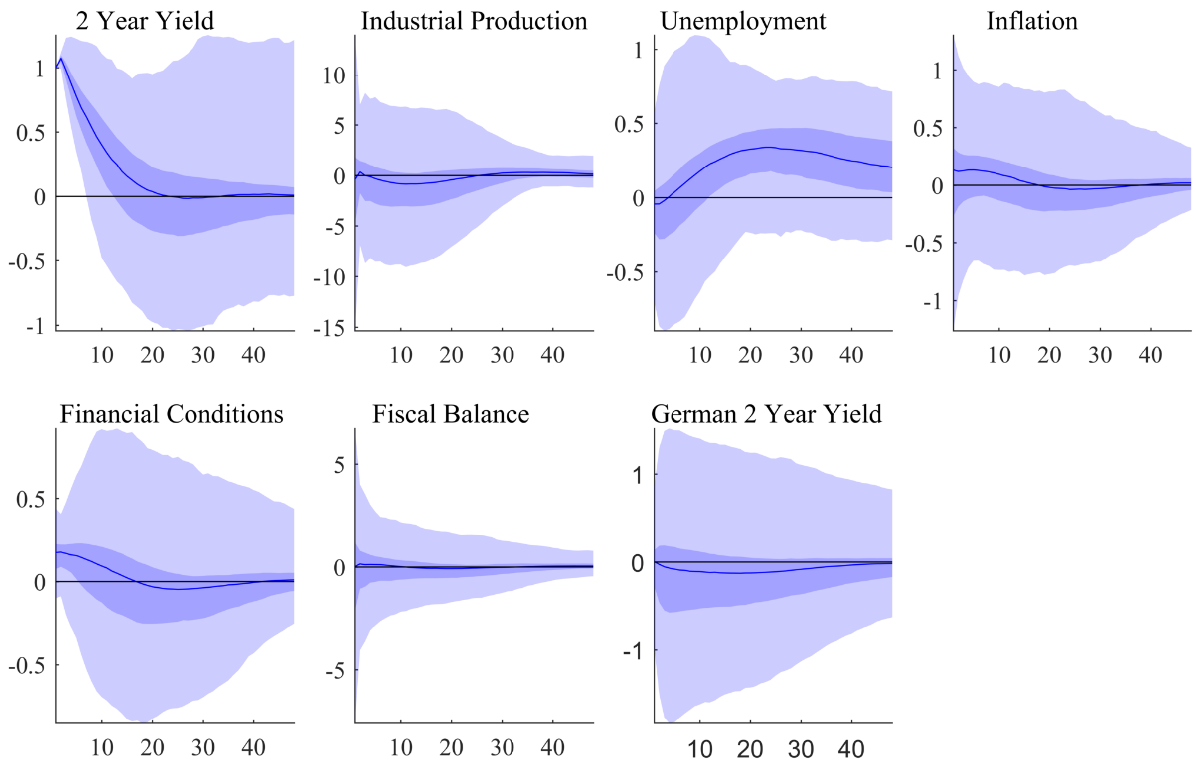
Figure A7: Mean Country Impulse Responses to a Sovereign Spread Shock: Treating pre-July 2009 as Missing

Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

Figure A8: Mean Country Impulse Responses to a Sovereign Spread Shock: Using 10 Year Bonds

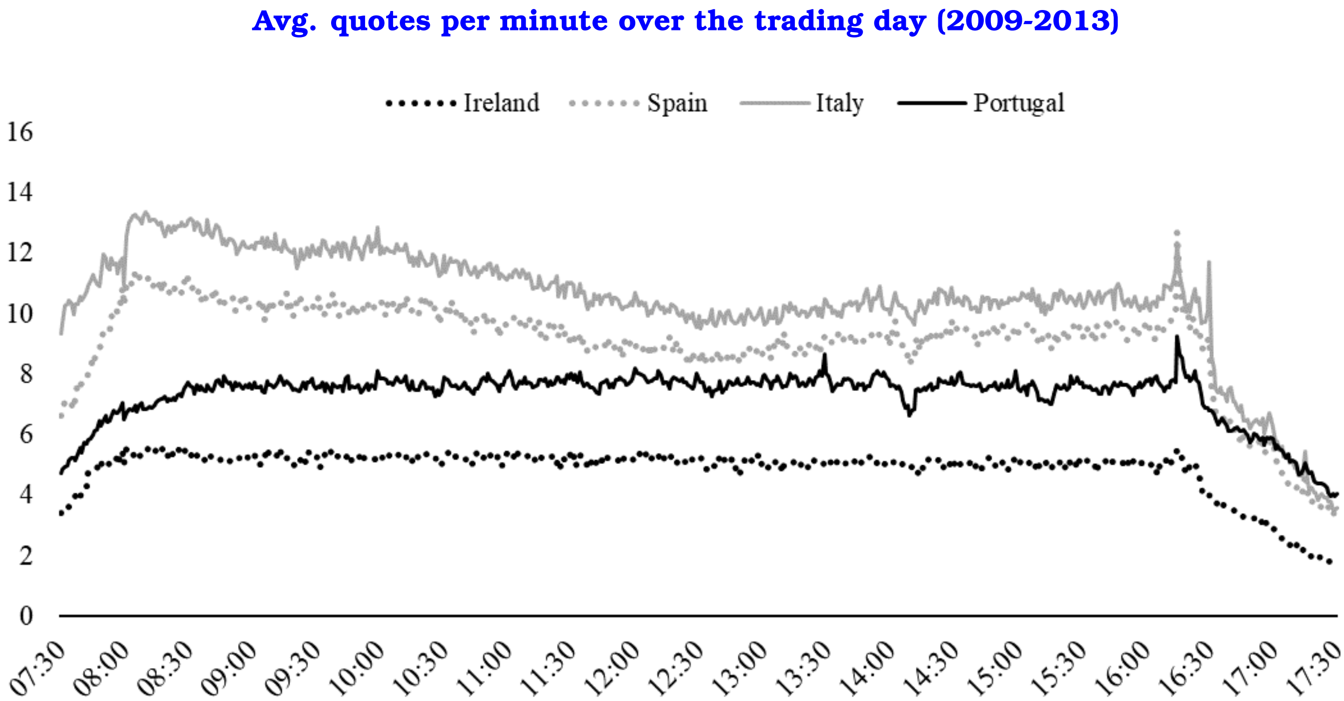
Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 10 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

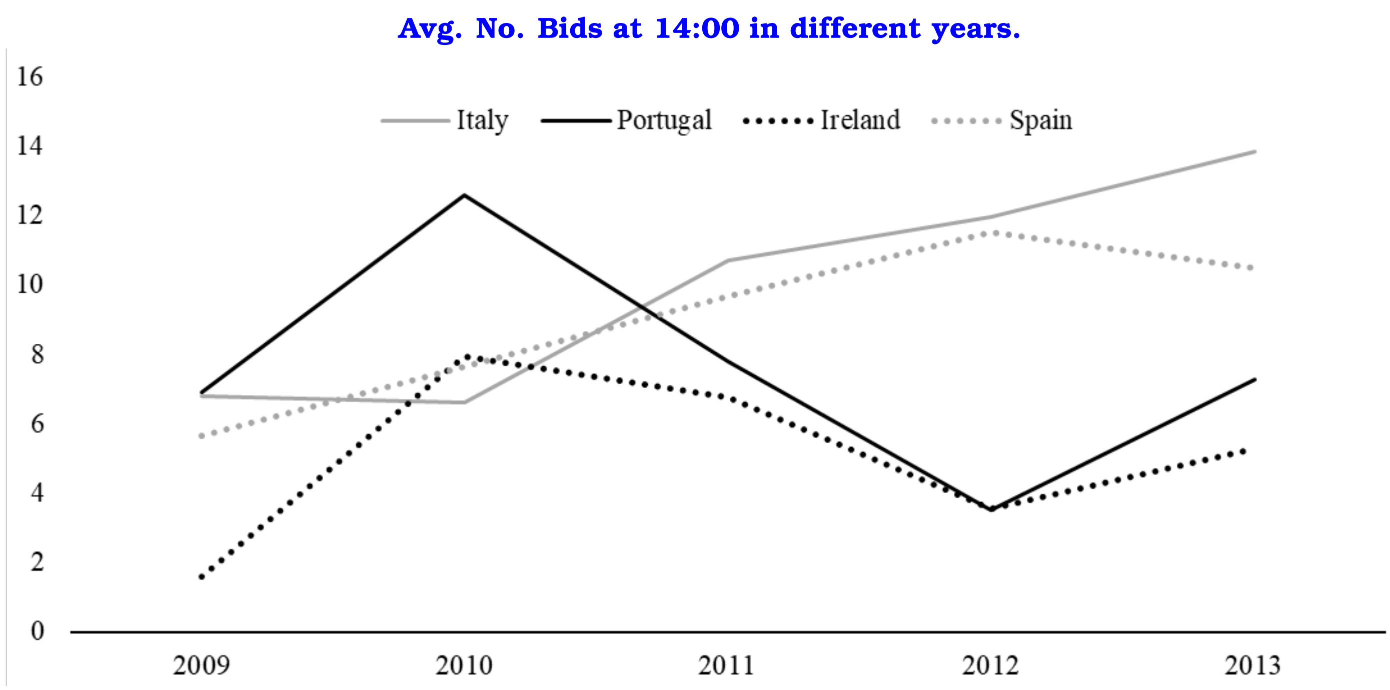
Figure A9: Mean Country Impulse Responses to a Sovereign Spread Shock: No Overnight Events



Notes: Impulse responses to a sovereign spread shock scaled to be consistent with a 100bp increase in the 2 year sovereign yield on impact. X-axis is months. Y-axis is percentages in all cases; Mean country model refers to impulse responses estimated using $\bar{\beta}$, $\bar{\Upsilon}$ and $\bar{\Sigma}$. Centre line is the median from the simulated posterior. Error bands are 68% and 90% Bayesian credible intervals.

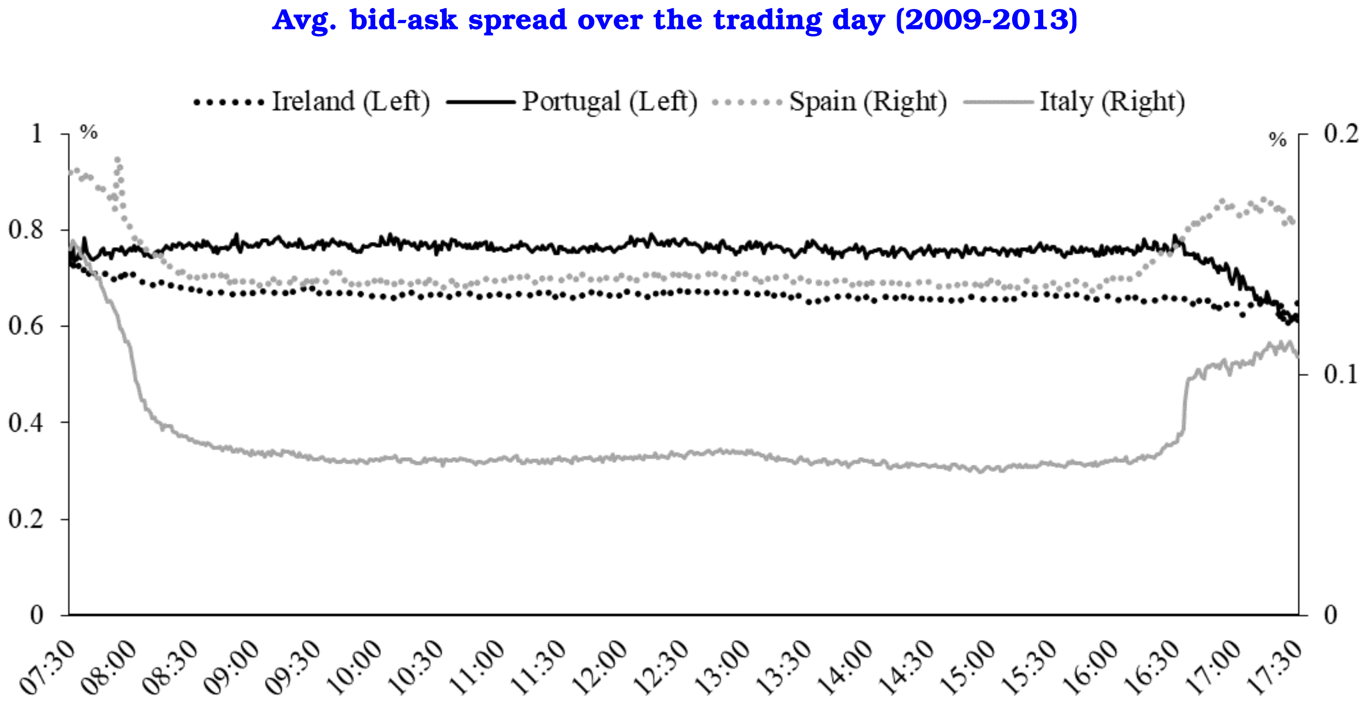
Figure A10: Quotes per minute

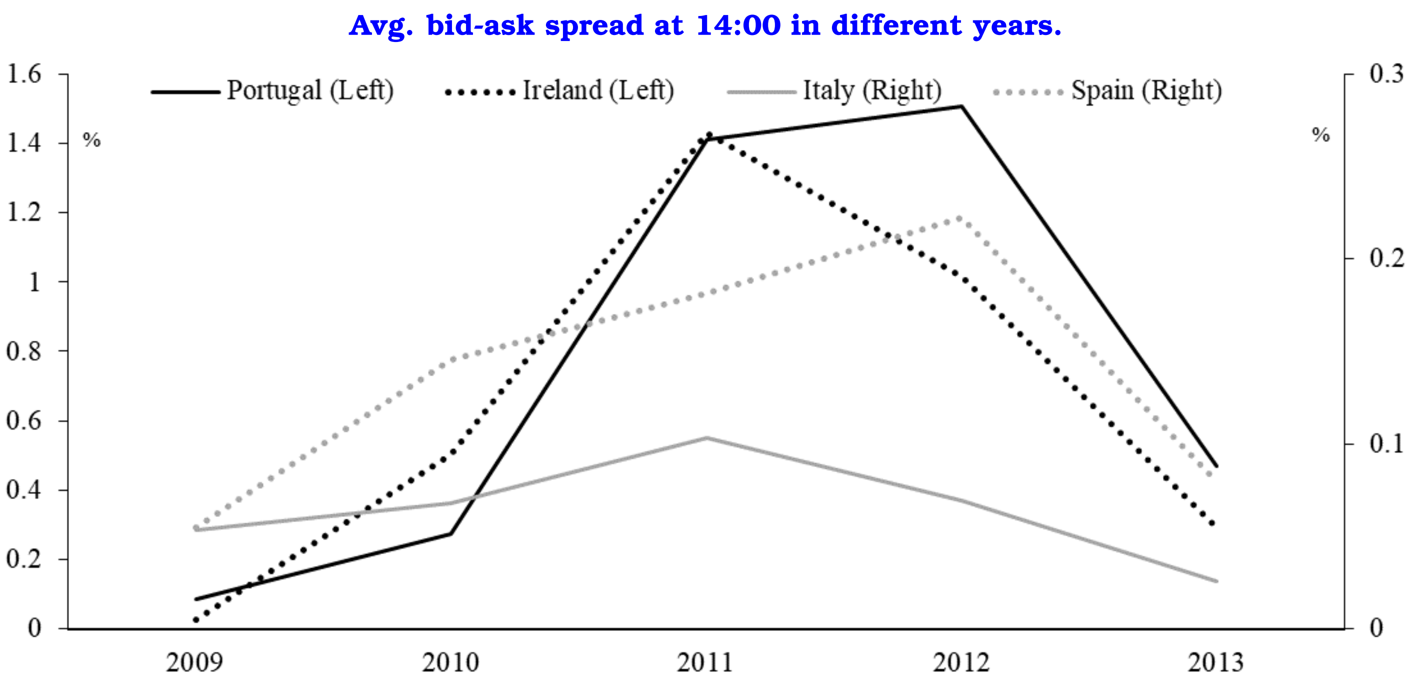




Notes: Data is taken from the "number of bids" field in Thomson Reuters Datascope. Sample is all trading days over the period 2009-2013 where trading day exclude weekends, UK holidays and local holidays. All times refer to London.

Figure A11: Bids-Ask Spreads

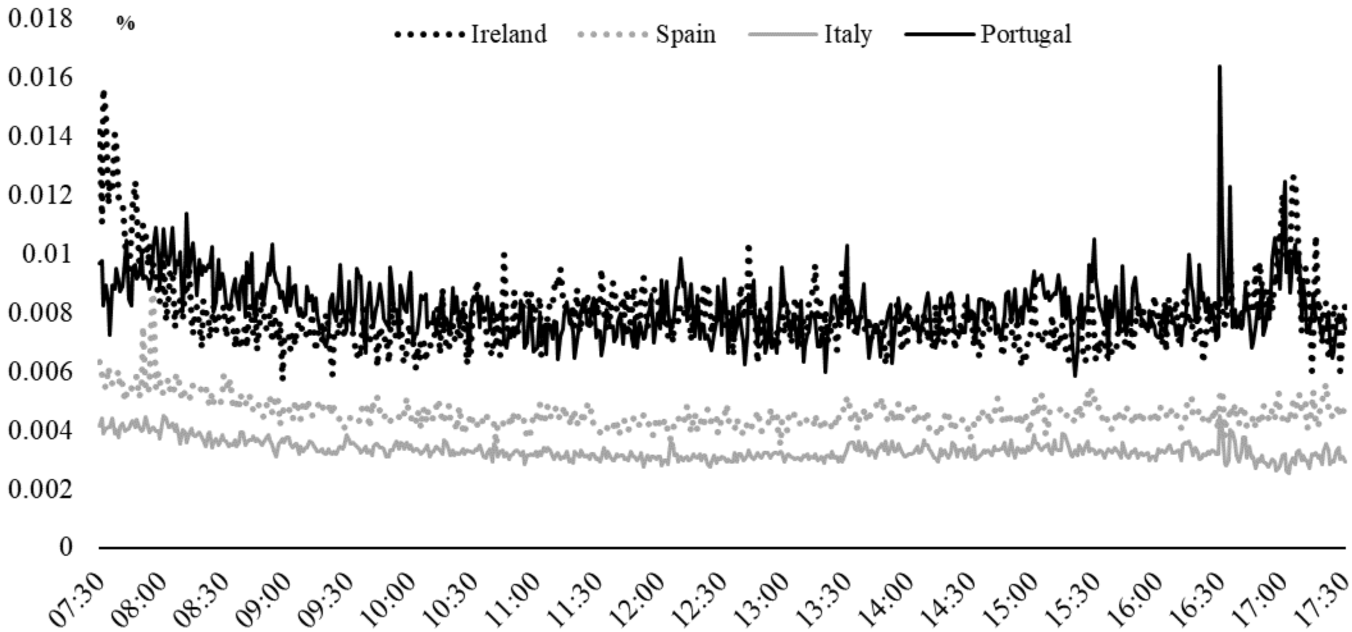




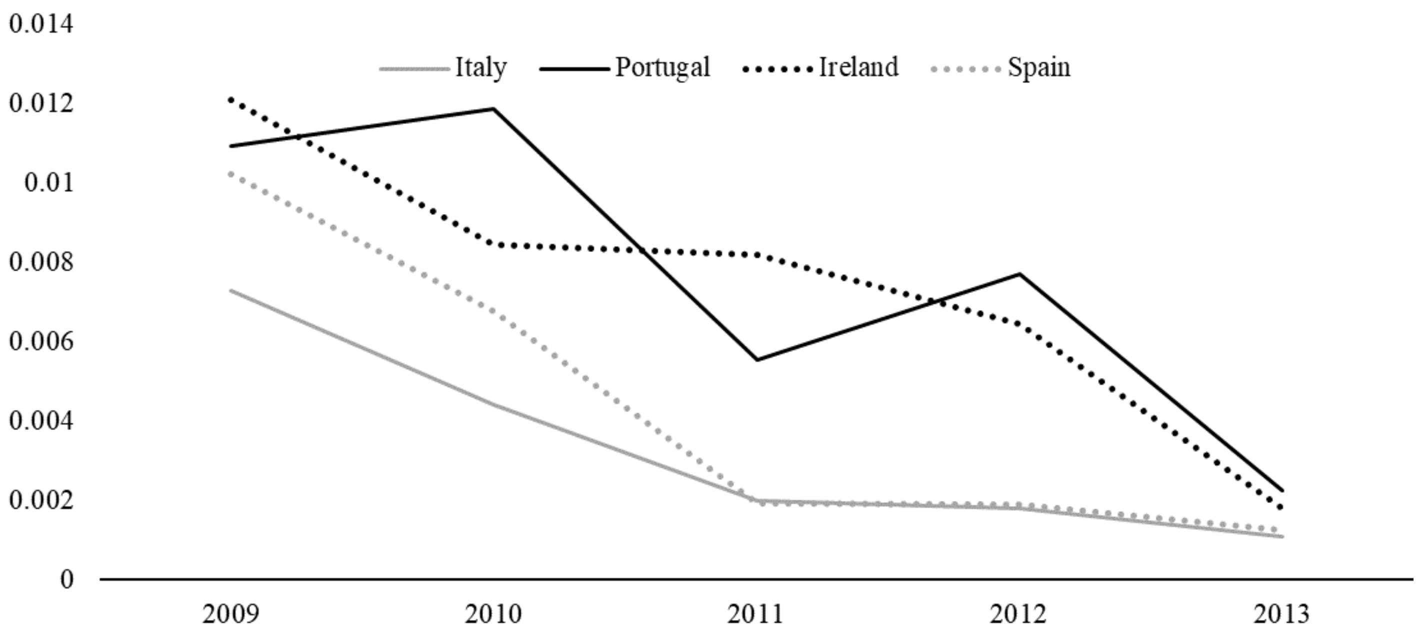
Notes: Data taken at one minute intervals the average of the “open bids” and “close bid” fields in Thomson Reuters Datascope less the average of the “open ask” and “close ask”. Irish and Portuguese bonds correspond to the left hand axis, Italian and Spanish the right hand axis. Sample is all trading days over the period 2009-2013 where trading days exclude weekends, UK holidays and local holidays. All times refer to London.

Figure A12: Average absolute change in yields

Avg. absolute minutely change in yields over the trading day (2009-2013)



Avg. absolute minutely change in yields 13:30-14:30 mean in different years.



Notes: Average absolute change in the mid-yield. Sample is all trading days over the period 2009-2013 where trading days exclude weekends, UK holidays and local holidays. All times refer to London.

Figure A13: An example of a Bloomberg newswire output to illustrate event timing