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Intraday Dynamics of Euro Area Sovereign Credit Risk Contagion

Luboš Komárek, Kristyna Ters, and Jörg Urban *

Abstract

We examine the role of the CDS and bond markets during and before the recent euro area sovereign debt crisis as transmission channels for credit risk contagion between sovereign entities. We analyse an intraday dataset for GIIPS countries as well as Germany, France and central European countries. Our findings suggest that, prior to the crisis, the CDS and bond markets were similarly important in the transmission of sovereign risk contagion, but that the importance of the bond market waned during the crisis. We find flight-to-safety effects during the crisis in the German bond market that are not present in the pre-crisis sample. Our estimated sovereign risk contagion was greater during the crisis, with an average timeline of one to two hours in GIIPS countries. By using an exogenous macroeconomic news shock, we can show that, during the crisis period, increased credit risk was not related to economic fundamentals. Further, we find that central European countries were not affected by sovereign credit risk contagion, independent of their debt level and currency.

JEL Codes: E44, G12, G14, G15.

Keywords: Contagion, credit default swaps, panel VAR, sovereign credit risk, sovereign debt crisis, spillover.

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Abstrakt

Práce se zaměřuje na zkoumání role trhu CDS a dluhopisového trhu jako transmisních kanálů nákazy kreditním rizikem mezi svrchovanými entitami během současné krize svrchovaného dluhu zemí eurozóny a před ní. Analyzujeme mezidenní údaje pro země skupiny GIIPS, Německo, Francii a středoevropské země. Naše výsledky ukazují, že před krizí byl trh CDS a dluhopisový trh podobně významný v rámci transmise šoku finanční nákazy, ale následně během krize významnost dluhopisového trhu slábla. V práci ukazujeme, že efekt „útěku ke kvalitě“ během krize na německém dluhopisovém trhu nebyl v předkrizovém období přítomný. Námi odhadnuté riziko svrchované nákazy bylo vyšší během krize s průměrnou odezvou relevantních informací v délce mezi jednou až dvěma hodinami pro země skupiny GIIPS. Aplikací exogenních makroekonomických informací majících charakter šoku ukazujeme, že během krizového období vzrostlo kreditní riziko, které však nebylo v souladu s vývojem fundamentálních proměnných. Dále ukazujeme, že středoevropské země nebyly dotčeny nákazou svrchovaného kreditního rizika bez závislosti na úrovni jejich veřejného dluhu a na jejich měně.

Nontechnical Summary

This paper analyses the issue of financial shock contagion among countries during the euro area sovereign debt crisis. An important motivation for providing financial support to Greece, despite the no-bailout clause in the Maastricht Treaty, was the fear on the part of policy makers that a Greek default would spill contagiously over to other highly indebted countries. This resulted in a sharp rise in sovereign credit spreads for a number of euro area countries. Due to these developments, policy makers and regulators paid increased attention to the market for sovereign credit default swaps (CDS).

By using CDS and bond data for GIIPS countries, France and Germany as well as central European countries we are able to identify the transmission channels for credit risk contagion between countries and evaluate the relative importance of these markets as transmission channels. Our analysis enables policy makers and regulators to focus on the more important transmission channel in order to promote financial stability.

Our unique intraday dataset allows us to estimate the contagion effects with substantially more accuracy than existing studies have done. Further, we are able to capture the intraday patterns of credit risk contagion. Shocks that may seem to affect several countries simultaneously when viewed at a daily or lower data frequency are revealed, through the lens of intraday data, to have possible origins in one particular country with clear contagion effects on other countries.

We employ a panel VAR methodology that can control for both country-specific risk and contagion effects across countries. Panel VARs are built on the same logic as standard VARs, but, by adding a cross-sectional dimension, they become a much more powerful tool for addressing policy questions related to the transmission of shocks across borders.

Our four main contributions to the existing literature are the following: First, while we find clear contagion during the euro area sovereign debt crisis, the pre-crisis period is rather characterised by comovement dynamics.

Second, our findings suggest that, prior to the crisis, the CDS and bond markets were similarly important in the transmission of financial shock contagion, but that the importance of the bond market waned during the crisis.

Third, markets for sovereign credit risk during the pre-crisis period were driven by macroeconomic news. Positive news led to a decrease in credit spreads and vice versa. The results for the euro area sovereign debt crisis show that movements in sovereign credit spreads did not respond to macroeconomic news but were rather driven by either monetary policy or exaggerations in financial markets due to lack of belief (self-fulfilling crisis).

Finally, countries outside the crisis region (central Europe) were practically unaffected by sovereign risk contagion during the crisis.

1. Introduction

The 2008-09 financial crisis caused investors to look more critically at the fiscal outlook in a number of countries, including several in the euro area. This resulted in a sharp rise in sovereign credit spreads for a number of euro area countries. At their peak, yield spreads on sovereign bonds relative to German bonds reached several hundred basis points, while before the global financial crisis these spreads had averaged only a few basis points. At the same time, market interest in trading credit risk protection on euro sovereign borrowers via credit default swaps (CDS) grew substantially and spreads on such instruments also surged. Due to these developments, policy makers and regulators paid increased attention to the market for sovereign CDS. Of particular interest was, first, the interplay between the pricing of sovereign risk in CDS and bond markets, secondly the possibility that one market could be systematically leading the other, and thirdly the potential for sovereign credit risk contagion effects. In this paper we focus on the latter.

We analyse euro area sovereign credit risk contagion effects in GIIPS[1] countries plus France and Germany from January 2008 to end-December 2011, which we split into a pre-crisis and crisis period. Further, we investigate if and how sovereign credit risk contagion was transmitted from the GIIPS countries to central European countries (Austria, the Czech Republic, Hungary, Poland and Slovakia) during the euro area sovereign debt crisis. Austria is included as low-risk reference country for the Czech Republic. The use of intraday CDS and bond data lets us estimate credit risk contagion effects with substantially more accuracy than existing studies on sovereign credit markets have done. In addition, little is yet known about the transmission channels of credit risk contagion through the CDS and the bond market, and their relative importance in the euro area sovereign debt crisis. As we have data for both the CDS market and the bond market, we are able to assess the contagion impacts conditioned on the credit channel.

The use of intraday data allows us to capture the intraday patterns of credit risk contagion. Indeed, shocks that may seem to affect several countries simultaneously when viewed at a daily or lower data frequency are revealed, through the lens of intraday data, to have possible origins in one particular country with clear contagion effects on other countries. Also, Gyntelberg et al. (2013) discuss the advantages of using intraday data due to the higher accuracy of the results as compared with lower-frequency data. They find that when using daily data, due to the dramatically smaller number of observations, the confidence bands of the estimated coefficients are extremely wide and therefore in most cases not significant. Existing research has differentiated between cross-country and intra-country analysis. Using a panel VAR methodology we can control for both country-specific risk and contagion effects across countries. Panel VARs are built on the same logic as standard

[1] Greece, Ireland, Italy, Portugal, Spain

VARs but, by adding a cross-sectional dimension, they become a much more powerful tool for addressing policy questions of interest related, for example, to the transmission of shocks across borders (Canova and Ciccarelli, 2013). By using the method of Canova and Ciccarelli (2013), we are able to shock the credit risk of an individual country and derive the individual response for each country in the panel.

A large body of literature concerns itself with the potential reasons and transmission channels for contagion as well as with the theoretical modelling of contagion. A whole strand of this literature focuses on empirical tests for the existence of contagion in a given stress period, that is, it asks if there are stronger cross-market linkages in times of crisis. This paper belongs to the latter type, as we focus on testing for the existence of contagion during the euro area sovereign debt crisis. We extent existing research by analysing the relative importance of the CDS and the bond market as transmission channels for sovereign risk contagion.

An important motivation for providing financial support to Greece, despite the no-bailout clause in the Maastricht Treaty, was the fear on the part of policy makers that a Greek default would spill contagiously over to other highly indebted countries in the euro area (Constancio, 2012). As pointed out by Corsetti et al. (2011), there is much disagreement among economists about the exact definition of contagion and how it should be tested. For Constancio (2012) and Forbes (2012) contagion occurs when financial or macroeconomic imbalances (shocks) create a systemic risk beyond that explained by economic fundamentals. Contagion differs from macroeconomic interdependence (comovement) among countries in that the transmission of risk to other countries is different under normal economic conditions. Forbes (2012) defines contagion as spillovers resulting from extreme negative effects. If comovements of markets are similarly high during non-crisis periods and crisis periods, then there is only evidence of strong economic linkages between these economies (Missio and Watzka, 2011). Kaminsky et al. (2003) describe contagion as an episode in which there are significant immediate effects in a number of countries following an event, such as when the consequences are fast and furious and evolve over a matter of hours or days. When the effect is gradual, Kaminsky et al. (2003) refer to it as spillovers rather than contagion. We rely on the contagion and comovement definitions according to Kaminsky et al. (2003), Constancio (2012) and Forbes (2012).

As there is a vast literature on contagion, we limit our discussion to papers that measure contagion among sovereign credit markets. Kamin and von Kleist (1999), Eichengreen and Mody (2000), Mauro et al. (2002), Pan and Singleton (2005), Longstaff et al. (2011), and Ang and Longstaff (2011) concentrate on the relationship between sovereign credit spreads and common global and financial market factors. These papers empirically identify factors that are significant variables for CDS credit spreads, such as the U.S. stock and bond market returns as well as the embedded volatility risk premium.

The issues of financial shock contagion and cross-country spillovers among countries in the euro area during the recent sovereign debt crisis have figured prominently in recent empirical research. Caporin et al. (2012) analyse risk contagion using the CDS spreads of the major euro area countries using different econometric approaches such as Bayesian modelling. They find that the diffusion of shocks in euro area CDS has been remarkably constant, while the risk spillover among countries is not affected by the size of the shock. Other examples are Bai et al. (2015), Neri and Ropele (2013), De Santis (2012), and Arghyrou and Kontonikas (2012). They all employ time series modelling approaches for contagion and include sovereign bond spreads (yield to maturity) to reflect pure credit risk considerations and macroeconomic variables. The results are mostly discriminated in terms of core (e.g. Germany and France) and peripheral countries (GIIPS). In general, they find that the bond spreads of lower-rated countries increase along with their Greek counterparts. However, their results in terms of magnitude, responses to shocks and contagion effects on core countries are somewhat mixed. Similarly to these studies, Koop and Korobilis (2014) employ a panel approach that is superior in empirically modelling financial contagion (Canova and Ciccarelli, 2013). Their findings are at odds with the discrimination between core and peripheral countries, as they also find contagion from GIIPS to core countries, albeit smaller in magnitude. The different results reported in these studies could be due to sample differences or to how bond spreads are calculated. Most empirical research uses the “constant maturity” approach to calculate bond yield differences (relative to Germany). Further, daily or weekly data are used for the empirical analysis, which may lead to inaccurate shock and contagion estimations, especially in periods when activity in sovereign risk markets is high during times of stress.

One of the key contributions of our paper to the existing literature on sovereign risk contagion during the recent euro area sovereign debt crisis is that, in contrast to all the above-mentioned studies, we do not use simple yield differences as our measure of cash spreads. Rather, we use carefully constructed asset swap spreads (ASW) based on estimated zero-coupon government bond prices. This ensures that we are comparing like with like in our empirical analysis for sovereign credit risk, by exactly matching the maturities and the cash flow structures of the CDS and the cash components. The use of ASW is also in line with the practice used in commercial banks when trading the CDS-bond basis. In addition, by calculating ASW we are able to estimate contagion impacts on Germany such as flight to safety. Germany is not included in most contagion studies since German Bund yields are used as the risk-free interest rate in the “constant maturity” approach mentioned above. Moreover, our analysis relies on intraday price data for both CDS and bonds, allowing us to estimate the contagion dynamics and the transmission channel of contagion (CDS or bond market) substantially more accurately than existing studies.[2] Further, by extending our model to include the economic surprise index, we are able to estimate how much of sovereign risk contagion can be attributed to macroeconomic news or overreaction/lack of belief by market participants.

Finally, our findings will improve the understanding of the dynamics in the market for sovereign credit risk. Further, the segregation of the credit risk transmission channels will enable policy makers and regulators to better assess the relative importance of, and the risks arising from, the derivative and cash market.

The rest of the paper is structured as follows. Section 2 discusses our data and the relationship between CDS and bonds. Section 3 discusses the set-up and estimation of the panel VAR (PVAR) model and its extension. Section 4 provides the empirical results and Section 5 concludes.

2. Data

The core data we use in our empirical analysis consists of USD-denominated five-year maturity intraday quotes on CDS contracts and government bonds for France, Germany, Greece, Ireland, Italy, Portugal and Spain (GIIPS). We choose this group of countries as it includes the countries most affected by the euro sovereign debt crisis, as well as Germany, which serves as the near-risk-free reference country, and France, which we consider as a low-risk control country. Further, we include Austria, the Czech Republic, Hungary, Poland and Slovakia in our sample, where Austria serves as a reference country for the Czech Republic.

According to Gyntelberg et al. (2013) when one considers the number of quotes of CDS contracts at the peak of the sovereign debt crisis in 2010, the five-year segment is the most liquid. The use of intraday data in our empirical analysis enables us to obtain much sharper estimates and clearer results with respect to market mechanisms (Gyntelberg et al., 2013). Further, they show that sovereign credit risk dynamics follow an intraday pattern.

Our sovereign bond price data come from MTS (Mercato Telematico dei Titoli di Stato). The MTS data comprise both actual transaction prices and binding bid-offer quotes. The number of transactions of sovereign bonds on the MTS platform is, however, insufficient to allow us to undertake any meaningful intraday analysis. Therefore, we use the trading book from the respective domestic MTS markets.[3] The MTS market is open from 8:15 to

[2] As presented in Gyntelberg et al. (2013). They discuss the advantages of using intraday data due to the higher accuracy of the results as compared with lower-frequency data.

[3] We ignore quotes from the centralised European platform (market code: EBM), as quotes for government bonds on the centralised platform are duplicates of quotes on the domestic platforms.

17:30 local Milan time, preceded by a pre-market phase (7.30 to 8.00) and an offer-market phase (8:00 to 8:15). We use data from 8:30 to 17:30.[4]

The CDS data consist of price quotes provided by CMA (Credit Market Analysis Ltd.) Datavision. CMA continuously gathers information on executable and indicative CDS prices directly from the largest and most active credit investors. After cleaning and checking the individual quotes, CMA applies a time- and liquidity-weighted aggregation so that each reported bid and offer price is based on the most recent and liquid quotes. The CDS market, which is an OTC market, is open 24 hours a day. However, most of the activity in the CMA database is concentrated between around 7:00 and 17:00 London time. As we want to match the CDS data with the bond market data, we restrict our attention to the period from 8:30 to 17:30 CET (CEST during summer).

We construct our intraday data on a 30-minute sampling frequency on our available data set, which spans from January 2008 to end-December 2011. The available number of indicative quotes for CDS does not allow a data frequency higher than 30 minutes. The euro area sovereign CDS markets were very thin prior to 2008, which makes any type of intraday analysis before 2008 impossible. Microstructural noise effects may come into play when high frequency data is used (Fulop and Lescourret, 2007). However, this is unlikely to play any significant role for our data based on a 30-minute sampling frequency because we average the reported quotes over each 30-minute interval (for robustness checks and for a more detailed discussion please refer to Gyntelberg et al. (2013)).

When implementing our analysis we split the data into two subsamples. The first covers the period January 2008 to 19 October 2009 and, as such, represents the period prior to the euro area sovereign debt crisis. While this period includes the most severe phase of the financial crisis, including the default of Lehman Brothers, it is relatively unaffected by market distortions stemming from concerns about the sustainability of public finances in view of rising government deficits and therefore represents the pre-sovereign debt crisis period. The second subsample covers the euro area sovereign debt crisis period and runs from 20 October 2009 to end-December 2011. As the beginning of the crisis period, we designate 20 October 2009, when the new Greek government announced that official statistics on Greek debt had previously been fabricated. Instead of a public deficit estimated at 6% of GDP for 2009, the government now expected a figure at least twice as high.

[4] Due to our sampling frequency we have discard either the first or the last 15 minutes of each trading day. We analysed the data quality in both intervals and found, on balance, that the first 15 minutes have a lower data quality than the interval at the end of the trading day. Consequently, we discarded the first 15 minutes of each trading day.

We employ CDS and bond data in our analysis in order to be able to differentiate between the transmission of contagion according to the credit risk channel from one country to another. Duffie (1999) argues that, since the CDS and the bond yield spread both price the same default of a given reference entity, their price should be equal if markets are perfect and frictionless. Thus, in a perfect market, due to arbitrage, the CDS spread equals the bond yield over the risk-free rate. However, for this parity to hold, a number of specific conditions must be met, including that markets are perfect and frictionless, that bonds can be shorted without restrictions or cost and that there are no tax effects, etc. Only floating rate notes should be used for the bond yield computation, because these bonds (unlike plain vanilla bonds) carry only credit rate risk and no interest rate risk. However, they are relatively uncommon, in particular for sovereign entities. A further complication linked to the use of fixed-rate or plain vanilla bonds as substitutes is that it is unlikely that the maturity of these instruments exactly matches that of standard CDS contracts.

To ensure proper comparability with CDS, Gyntelberg et al. (2013) employ synthetic par asset swap spreads (ASW) for the bond leg of the basis. The use of ASW is in line with the practice used in commercial banks when trading the CDS-bond basis. By calculating ASW for our empirical analysis, we ensure an accurate cash flow matching, as opposed to studies that use simple "constant maturity" yield differences for credit risk.

An asset swap is a financial instrument that exchanges the cash flows from a given security - e.g. a particular government bond - for a floating market rate.[5] This floating rate is typically a reference rate such as Euribor for a given maturity plus a fixed spread, the ASW. This spread is determined such that the net value of the transaction is zero at inception. The ASW allows the investor to maintain the original credit exposure to the fixed rate bond without being exposed to interest rate risk. Hence, an asset swap on a credit risky bond is similar to a floating rate note with identical credit exposure, and the ASW is similar to the floating-rate spread that theoretically should be equivalent to a corresponding CDS spread on the same reference entity. Specifically, the ASW is the fixed value A required for the following equation to hold[6] (O'Kane, 2000):

$$\underbrace{100-P}_{\substack{\text{Upfront payment for bond}\\ \text{asset in return for par}}} + \overbrace{\underbrace{C\sum_{i=1}^{N_{\text{fixed}}} d(t_i)}_{\text{Fixed payments}} = \underbrace{\sum_{i=1}^{N_{\text{float}}} \big(L_i+A\big)d(t_i)}_{\text{Floating payments}}}^{\text{Interest rate swap}}, \tag{2.1}$$

5 See O'Kane (2000) and Gale (2006) for detailed discussions of the mechanics and pricing of asset swaps.

6 This assumes that there is no accrued coupon payment due at the time of the trade; otherwise, an adjustment factor would need to be added to the floating payment component.

where P is the full (dirty) price of the bond, C is the bond coupon, L_i is the floating reference rate (e.g. Euribor) at time t_i and $d(t_i)$ is the discount factor applicable to the corresponding cash flow at time t_i.

In order to compute the ASW A, several observations and simplifications have to be made. First, in practice it is almost impossible to find bonds outstanding with maturities that exactly match those of the CDS contracts and second, the cash-flows of the bonds and the CDS will not coincide. To overcome these issues, in what follows we use synthetic asset swap spreads based on estimated intraday zero-coupon sovereign bond prices. Specifically, for each interval and each country, we estimate a zero-coupon curve based on all available bond price quotes during that time interval using the Nelson and Siegel (1987) method. With this procedure, we are able to price synthetic bonds with maturities that exactly match those of the CDS contracts, and we can use these bond prices to back out the corresponding ASW. As this results in zero coupon bond prices, we can set C in Equation (2.1) to zero.

A CDS contract with a maturity of m years for country j in time interval k of day t, denoted as $S_j(t_k,m)$, has a corresponding ASW $A_j(t_k,m)$:

$$100 - P_j(t_k,m) = \sum_{i=1}^{N_m} \left(L_i(t_k) + A_j(t_k,m)\right) \cdot d(t_k,t_i), \tag{2.2}$$

with $P_j(t_k,m)$ as our synthetic zero coupon bond price.

For the reference rate L_i in Equation (2.2), we use the 3-month Euribor forward curve to match as accurately as possible the quarterly cash flows of sovereign CDS contracts. We construct the forward curve using forward rate agreements (FRAs) and euro interest rate swaps. We collect the FRA and swap data from Bloomberg, which provides daily (end-of-day) data. 3-month FRAs are available with quarterly settlement dates up to 21 months ahead, i.e. up to 21×24. From two years onwards, we bootstrap zero-coupon swap rates from swap interest rates available on Bloomberg and back out the corresponding implied forward rates. Because the swaps have annual maturities, we use a cubic spline to generate the full implied forward curve, thereby enabling us to obtain the quarterly forward rates needed in Equation (2.2).

Given our interest in intraday dynamics, we follow Gyntelberg et al. (2013) and generate estimated intraday Euribor forward rates by assuming that the intraday movements of the Euribor forward curve are proportional to the intraday movements of the German government forward curve.[7] To be precise, for each day, we calculate the difference between our

[7] Euribor rates are daily fixing rates, so we are actually approximating the intraday movements of the interbank interest rates for which Euribor serves as a daily benchmark.

Euribor forward curve and the forward curve implied by the end-of-day Nelson-Siegel curve for Germany.[8] We then keep this difference across the entire curve fixed throughout that same day and add it to the estimated intraday forward curves for Germany earlier on that day to generate the approximate intraday Euribor forward curves. This approach makes the, in our view, reasonable assumption that the intraday variability in Euribor forward rates will largely mirror movements in corresponding German forward rates.

Finally, we need to specify the discount rates $d(t_k, t_i)$ in Equation (2.2). The market has increasingly moved to essentially risk-free discounting using the overnight index swap (OIS) curve. We therefore take $d(t_k, t_i)$ to be the euro OIS discount curve, which is constructed in a way similar to the Euribor forward curve. For OIS contracts with maturities longer than one year, we bootstrap out zero-coupon OIS rates from interest rates on long-term OIS contracts. Thereafter, we construct the entire OIS curve using a cubic spline. We use the same technique as described above to generate approximate intraday OIS discount curves based on the intraday movements of the German government curve.

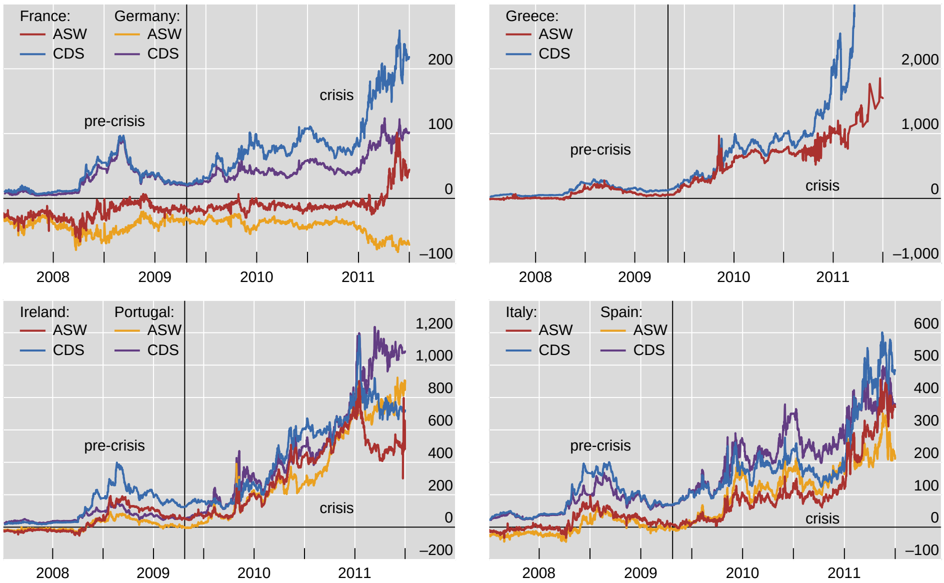
To gauge the potential impact of this assumption on our empirical results, we reestimate our model using an alternative assumption that the Euribor and OIS curves are fixed throughout the day at their observed end-of-day values. Under this alternative assumption, we obviously fail to capture any movements in money market rates within the day when we price our synthetic asset swaps. Our results remain robust.

Please refer to Gyntelberg et al. (2013) for an in-depth discussion of the construction of our intraday ASW and to O'Kane (2000) for a general discussion.

According to different panel unit root tests (see Appendix C) our CDS and ASW price data (displayed in Figure 1) is I(1). Therefore, we estimate our subsequent models in first differences.

[8] Here we use the second to last 30-minute interval, because the last trading interval is occasionally overly volatile.

Figure 1: CDS and ASW Spreads in Basis Points



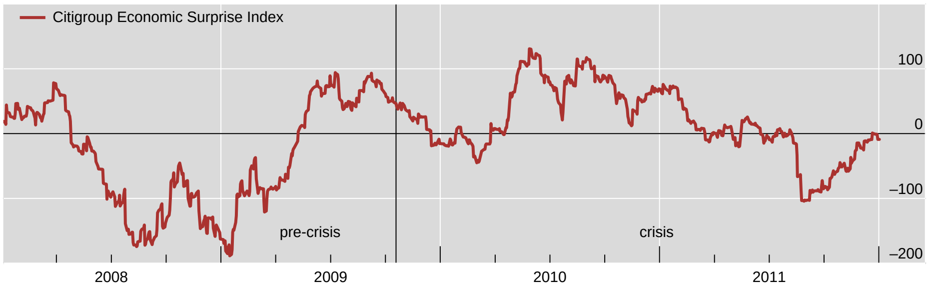
Notes: The figures are based on data with a 30-minute sampling frequency. Our split into the pre- and the crisis period is indicated by the vertical line in each figure. Due to the Greek debt restructuring the data for Greece ends in September 2011.

In our PVARX model extension (Section 3.2), we make use of the Citigroup economic surprise index for the euro zone[9] as an exogenous variable (see Figure 2). This index is widely recognised in academia and by practitioners for measuring unexpected economic news (such as in Goldberg and Grisse (2013), Scotti (2013), and Paulsen (2014)).

The Citigroup economic surprise index measures how economic news/data is developing relative to the anticipated consensus forecasts of market economists. According to Citigroup, the index captures objective and quantitative measures of economic news, defined as weighted historical standard deviations of data surprises (actual releases versus the Bloomberg survey median). A positive reading of the economic surprise index for the euro zone suggests that economic releases have on balance beaten the consensus, while a negative reading indicates the opposite. The index captures economic news on macroeconomic and fiscal variables such as employment change, the housing market, retail sales, debt-to-GDP, the budget deficit and consumer confidence in the euro zone. Thus, the Citigroup economic surprise index does not include news on monetary policy decisions.

[9] Bloomberg ticker: CESIEUR Index

Figure 2: Citigroup Economic Surprise Index



Notes: The figure shows the daily values of the Citigroup Economic Surprise Index (Bloomberg ticker: CESIEUR Index) from the beginning of 2008 until the end of 2011. The vertical line corresponds to 19 October 2009, the end of our pre-crisis period. Source: Bloomberg.

The economic surprise index has a different daily frequency from the intraday data that we are analysing in this paper. However, as market participants are exposed to economic news throughout the whole day, we disperse the actual Citigroup economic surprise index data given at the end of the trading day over that entire day. We conducted several simulations with different distributions to generate a pseudo-intraday economic surprise index. Our results remain extremely robust to this experiment. The experiment design is justified because we are interested not in the exact time line of the absorption of unexpected macroeconomic news, but rather in a qualitative picture of whether markets react to fundamental macroeconomic news in the pricing of sovereign credit risk.

3. Modelling Sovereign Credit Spread Contagion

To empirically measure the impact of euro area sovereign credit risk contagion effects according to the credit risk channel (CDS and bond market), we employ a panel vector autoregressive (PVAR) model. PVARs have the same structure as VAR models, in the sense that all variables are assumed to be endogenous and a cross-sectional dimension is added to the representation. We define our PVAR model following Binder et al. (2005) with fixed effects when N is finite and T is large, as $i = 1, ..., N$ is the cross-sectional dimension and $t = 1, ..., T$ is the time-series dimension in our model. According to Koop and Korobilis (2014) and Canova and Ciccarelli (2013), in this setup the PVAR is the ideal tool for examining the international transmission of macroeconomic or financial shocks from one country to another.

The PVAR has several advantages over individual country VARs in a time series framework. By analysing a panel of countries, we can more accurately model contagion from one country to another since the panel approach captures country-level heterogeneity. We control for cross-sectional heterogeneity by including fixed effects in the regression. By using CDS and ASW as endogenous variables for each country in our cross-section,[10] we can differentiate the credit risk channel of contagion, which improves the understanding of the market microstructural dynamics. With an extension of the PVAR using a purely exogenous variable, we can assess the effect of unexpected economic news[11] on credit risk contagion for the countries in our sample.

3.1 Panel VAR

In vector autoregressive models (VAR) all variables are treated as endogenous and interdependent in both a dynamic and static sense. The VAR model is formally defined as:

$$Y_t = A_0 + A(\mathrm{L})Y_{t-1} + u_t, \tag{3.3}$$

where Y_t is a G x 1 vector of endogenous variables and $A(\mathrm{L})$ is a polynomial in the lag operator, A_0 is a G x 1 vector and u_t is a G x 1 vector of i.i.d. shocks.

Panel VARs (PVAR) have the same structure as VAR models in Equation (3.3), as all variables are assumed to be endogenous and independent. However, a cross-sectional dimension i, in our case across countries, is added to the representation. Thus, Y_t is the stacked version of y_{it}, the vector of G variables for each country $i = 1,...,N$, i.e. $Y_t = (y'_{1t}, y'_{2t}, ..., y'_{Nt})'$ and $t = 1,...,T$. The major difference between a VAR and the PVAR is that the covariance σ_{ij} of the residuals is zero by definition for country i different from country j in a VAR model. The PVAR is defined as follows:

$$y_{it} = A_{0i} + A_i(\mathrm{L})Y_{t-1} + u_{it}. \tag{3.4}$$

A_{0i} are G x 1 vectors and A_i are G x GN matrices. We allow for country-specific heterogeneity by including a country-specific intercept. Further, lags of all endogenous variables of all entities enter the equation of country i. Canova and Ciccarelli (2013) call this feature "dynamic interdependencies". The residual u_{it} is a G x 1 vector and $u_t = (u_{1t}, u_{2t}, ..., u_{Nt})$. u_{it} is generally correlated across the cross-sectional dimension i. Canova and Ciccarelli (2013) call this feature "static interdependencies". Thus the variance-covariance matrix for a PVAR has the following property $E(u_{it}u'_{jt}) = \sigma_{ij} \neq 0$ for $i \neq j$, i.e. static interdependencies occur when the correlations between the errors in two countries' VARs are non-zero. On the other hand, dynamic interdependencies occur when one country's lagged

[10] bi-variate estimation per country

[11] By using the economic surprise index as a predetermined purely exogenous variable in the PVARX model.

variables affect another country's variables. Hence, the PVAR is more flexible compared to a VAR ($\sigma_{ij} = 0$ for $i \neq j$).[12]

In our bivariate case, i.e. $G = 2$, we can rewrite the PVAR in Equation (3.4) as:

$$\begin{pmatrix} \Delta CDS \\ \Delta ASW \end{pmatrix}_{it} = \begin{pmatrix} A_{01} \\ A_{02} \end{pmatrix}_{i} + \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}_{ji} (\mathrm{L}) \begin{pmatrix} \Delta CDS \\ \Delta ASW \end{pmatrix}_{jt-1} + \begin{pmatrix} u_1 \\ u_2 \end{pmatrix}_{it}, \tag{3.5}$$

where A_{mn} are scalars and $i, j = 1, ..., N$ is the number of countries in the cross-sectional dimension.

For the estimation, we follow the approach proposed by Canova and Ciccarelli (2009) of an unrestricted PVAR which allows for the selection of restrictions involving dynamic interdependencies, static interdependencies and cross-section heterogeneities.[13] According to an empirical model comparison by Koop and Korobilis (2014), the proposed methodology by Canova and Ciccarelli (2009) shows the best properties compared to other PVAR approaches. Canova and Ciccarelli (2009) suggest adopting a flexible structure through a factorisation of the coefficients in Equation (3.4). Through the flexible coefficient factorisation, the PVAR can be rewritten as a reparametrised multicountry VAR and estimated using SUR (Canova and Ciccarelli, 2009). The advantage of this flexible factorisation is that the overparametrisation of the original PVAR is dramatically reduced while, in the resulting SUR model, estimation and specification searches are constrained only by the dimensionality of the estimated coefficient matrix (for a more in-depth discussion please refer to Canova and Ciccarelli (2009) and Koop and Korobilis (2014)).

3.2 Panel VARX

As an extension to the previous analysis, we consider the response of credit risk in CDS and bond markets in our GIIPS and low-risk country sample to unexpected macroeconomic news. We follow Canova and Ciccarelli (2013) by extending the PVAR model in Equation (3.4) with a predetermined purely exogenous variable X_t which results in a PVARX model which takes the following form:

$$y_{it} = A_{0i} + A_i(\mathrm{L})Y_{t-1} + F_i(\mathrm{L})X_t + u_{it}, \tag{3.6}$$

[12] According to Canova and Ciccarelli (2013), these features distinguish a panel VAR typically used for macroeconomics and finance from a panel VAR used in microeconomics.

[13] We use demeaned and standardised first differences.

with X_t as a M x 1 vector (M is equal to the number of exogenous variables) common to all entities i. The PVARX can also be rewritten as:

$$\begin{pmatrix} \Delta CDS \\ \Delta ASW \end{pmatrix}_{it} = \begin{pmatrix} A_{01} \\ A_{02} \end{pmatrix}_i + \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}_{ji} (\mathrm{L}) \begin{pmatrix} \Delta CDS \\ \Delta ASW \end{pmatrix}_{jt-1} + \begin{pmatrix} \tilde{F}_1 \\ \tilde{F}_2 \end{pmatrix}_i (\mathrm{L}) X_t + \begin{pmatrix} u_1 \\ u_2 \end{pmatrix}_{it} . \tag{3.7}$$

We employ the economic surprise index as a predetermined exogenous variable X_t for unexpected macroeconomic news in the euro zone, i.e. $M = 1$.

The extension to the PVARX model allows us to analyse whether credit risk responses can be attributed to macroeconomic fundamental news or if exaggerations in terms of lack of belief of economic agents also contributed to credit risk responses.

4. Results

We carry out an impulse response analysis to investigate contagion of financial shocks across the euro area countries that were most affected in the sovereign debt crisis. Further, we present results on shock contagion to central European countries. We focus on individual country shocks propagating from GIIPS countries and analyse the impact of an unexpected one-unit shock to credit risk in both the CDS and ASW markets from country i to j. Finally, we present results of exogenous economic news shocks and the effect on sovereign risk in GIIPS countries.

In standard VAR models (see Equation (3.3)), shock identification is performed by imposing a Choleski decomposition in all countries. To reduce the number of identification restrictions in a VAR model, it is assumed that $E(u_{it}u'_{jt})$ is block diagonal, with blocks corresponding to each country. Canova and Ciccarelli (2009) state that block diagonality implies differences in the responses within and across countries. Within a country, variables are allowed to move instantaneously. But across entities, variables can only react with one lag.

The identification of shocks for PVAR models as defined in Equation (3.4) is more complicated, given that the PVAR model allows for static interdependencies, as u_{it} is correlated across entities i. Thus, cross-entity symmetry in shock identification cannot be assumed. We compute the impulse responses following Canova and Ciccarelli (2009) as the difference between two conditional forecasts: one where a particular variable is shocked and one where the disturbance is set to zero. For a more in-depth discussion of shock identification using conditional forecasts in PVAR models allowing for static interdependencies please refer to Canova and Ciccarelli (2009).

4.1 Results for GIIPS and Low-risk Countries

As a general result, we find that, pre-crisis, the bond and CDS markets are of similar importance, i.e. the response function of country i to a one-unit shock to the ASW and CDS markets of country j is of a comparable size in the two markets (see Figure 3). These results are as expected, as both markets should price the countries' credit risk equally (Duffie, 1999). During the crisis period, the CDS market becomes more relevant on balance (see Figure 4). Interestingly, the inter-market shock transmission, i.e. from CDS to ASW and vice versa, is not important during the pre-crisis period. This weak connection between the two markets during the pre-sovereign debt crisis period can be explained by different market participants and their distinct investment horizons. Insurance firms active in the bond market have a longer investment horizon than, for example, hedge funds in the CDS markets. Shocks to ΔCDS or ΔASW are very short-lived and may not be seen by participants in the other market. However, CDS and ASW levels are persistently affected, as shown in Appendix D. During the crisis period, shock transmission between markets becomes relatively more important, suggesting a stronger inter-market connectivity. Market participants get more vigilant to potential bad news, which may spill over from other markets.

Further, we find that the decay of a shock is faster on average in the pre-crisis period than in the crisis period (see Figures 3 and 4). The timelines of our estimated shock contagion and absorption are dramatically shorter than in existing empirical studies, such as Koop and Korobilis (2014), who find that shock contagion spreads on average within one to two months in the case of shocks that do not decay over a timeline of 10 months. We find for both sample periods that contagion propagates within the first 30-minute time interval. Therefore, responses to shock contagion are typically not lagged as found, for example, in Koop and Korobilis (2014). Further, the average response for shock absorption is around one hour in the pre-crisis period and slightly longer at one to two hours on average during the crisis period. This result is clearly in line with the generally accepted notion that financial markets react very fast to new information (Gyntelberg et al., 2013). The slower speed of shock absorption during the crisis seems to contradict our statement above that market participants are more reactive to news during crisis periods. This can be explained by the fact that the estimated timeline of shock absorption during the crisis period is strongly affected by turmoil in financial markets, while the pre-crisis period represents a relative normal market environment for European sovereign states without fast and furious shock contagion but rather with comovements across markets as defined by Constancio (2012) and Forbes (2012).

In the pre-crisis period, a credit risk shock spreading from the ASW to the CDS market and vice versa had more or less the same impact in terms of magnitude and shock absorption. Thus, the derivatives market and the spot market were about equally significant in terms of shock contagion prior to the euro area sovereign debt crisis. However, during the crisis period we find that shock transmission from the ASW to the CDS market had a dramatically lower impact than vice versa. This leads to the assumption that the importance of the spot market as a channel of financial shock contagion decreased during the euro area sovereign debt crisis. Thus, the contagion of shocks to credit risk has been transmitted predominantly through the derivatives market.

During the pre-crisis period, a one-unit shock to either the ASW or CDS of country i results in a spread widening for all countries. However, during the crisis, we find evidence of a flight to safety to German bonds, as Germany is considered a safe haven for investors. This effect is visible in the inter-market connection, i.e. a positive shock to a GIIPS country's CDS or ASW leads to spread tightening in German ASW, while we cannot report a similar effect for German CDS. Similar behaviour is not visible for France, despite it being considered a low-risk control country.

During the pre-crisis period, we find that the magnitude of the impulse responses is similar across all countries, while during the crisis period, GIIPS countries exhibit much larger impulse responses than the rest of our sample countries do.

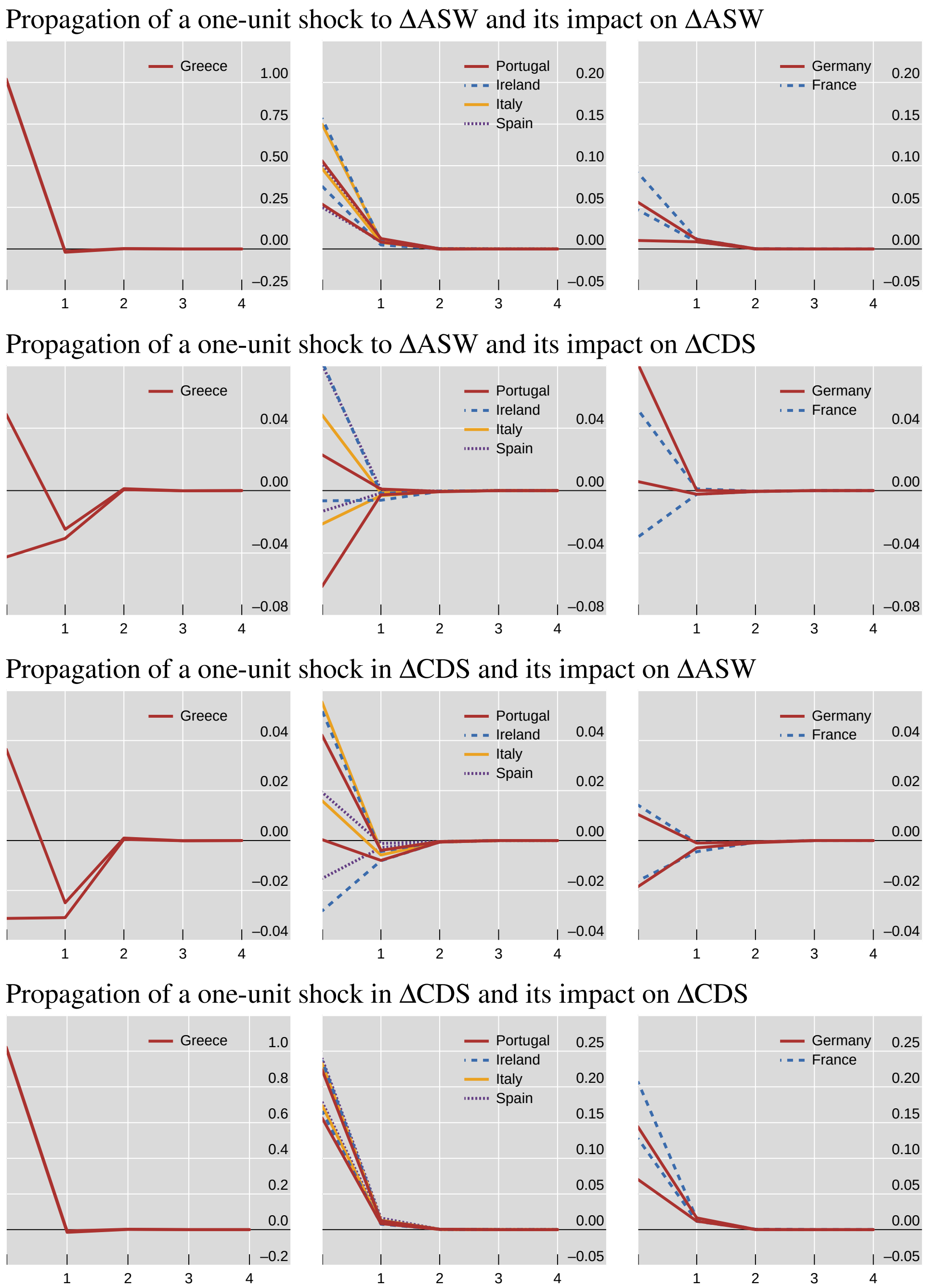
The forecasting precision is much more accurate during the crisis period, as the confidence bands are much tighter than in the pre-crisis period.

In contrast to the other empirical studies using this methodology, Koop and Korobilis (2014) find confidence bands for their impulse responses that all lie between positive and negative reactions to a one-unit shock to Greek bond yields relative to Germany. The advantage of our approach, using ASW and intraday data, dramatically increases the precision of the results during the crisis period.

In addition to the impulse response functions for a shock to Greek ΔASW and ΔCDS in Figures 3 and 4, we present impulse response functions for a shock to Spanish and Portuguese ASW and CDS in Appendix A[14].

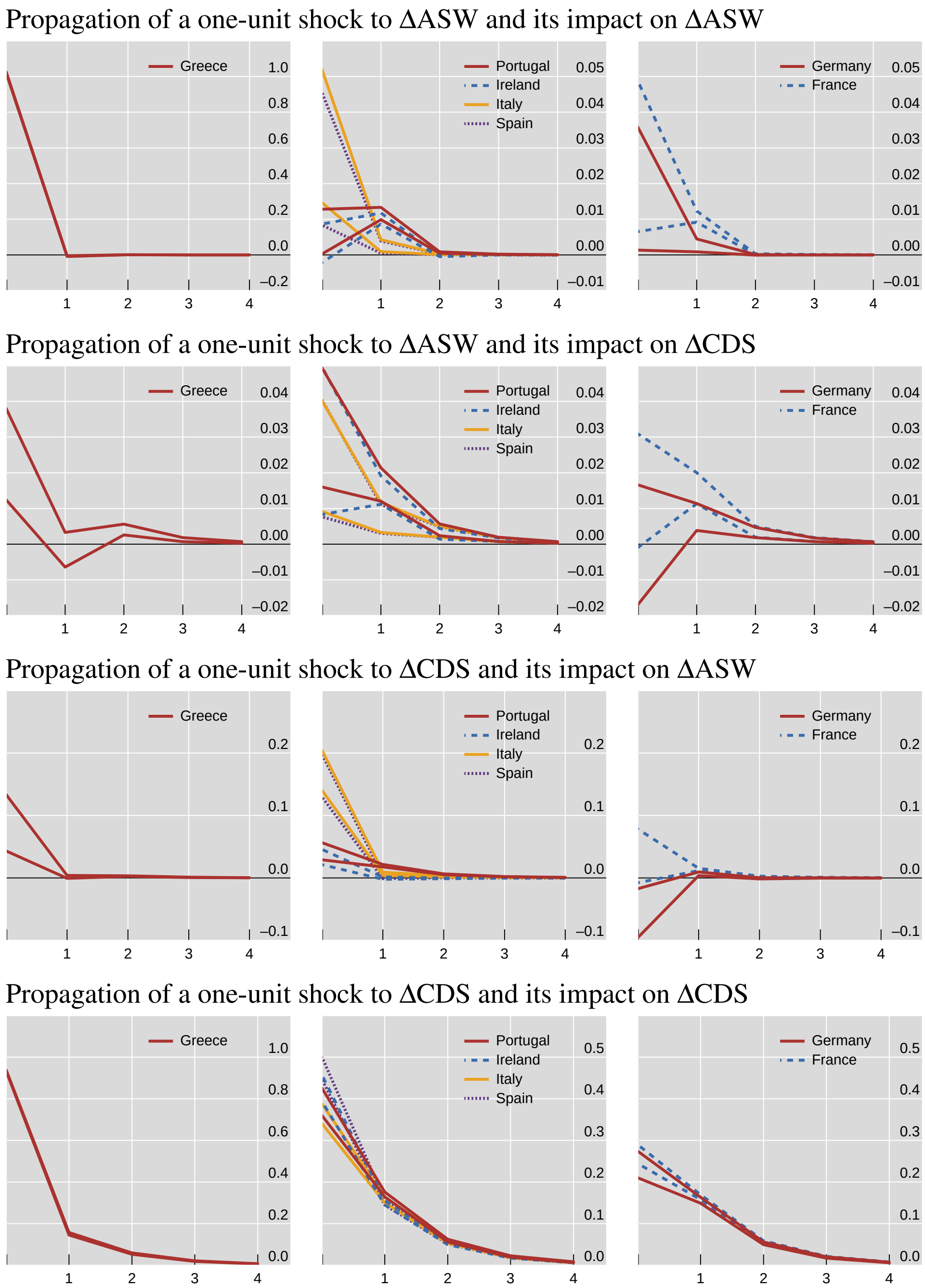
[14] Impulse response functions for a shock to Irish and Italian ASW and CDS show similar results and can be provided on request.

Figure 3: Impulse Responses in the Pre-crisis Period - Shock in Greece



Notes: This figure illustrates the impulse response for ΔCDS and ΔASW to a one-unit shock (increase) for the period from January 2008 to 19 October 2009. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the impulse response to a one-unit shock to Greek ΔCDS or ΔASW. The number of 30-minute time intervals is described by the x-axis. For each impulse response we plot the upper and lower 95% confidence bands.

Figure 4: Impulse Responses in the Crisis Period - Shock in Greece

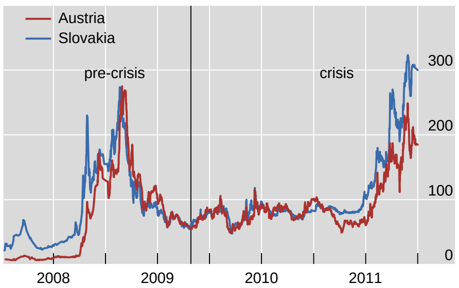


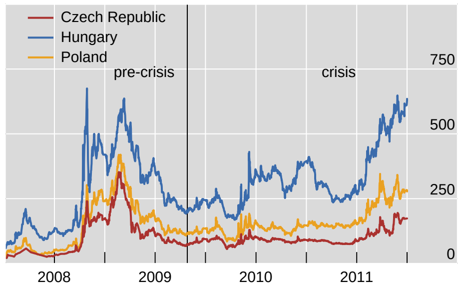
Notes: This figure illustrates the impulse response for ΔCDS and ΔASW to a one-unit shock (increase) for the period from 20 October 2009 to end-December 2011. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the impulse response to a one-unit shock to Greek ΔCDS or ΔASW. The number of 30-minute time intervals is described by the x-axis. For each impulse response we plot the upper and lower 95% confidence bands.

4.2 Results for Central European Countries during the Crisis Period

This section presents the results of an unexpected one-unit shock to CDS credit risk propagating from GIIPS countries and the shock response in central European countries. Due to the illiquidity of the bond markets in the central European countries in our sample, we were only able to conduct an intraday analysis based on CDS data during the crisis period. This, however, does not limit the validity of our analysis, because the results for GIIPS and low-risk countries in Section 4.1 strongly indicate that bond markets were not the main venue of sovereign credit shock contagion during the crisis. Thus, the PVAR model in Equation (3.4) applied in this section is estimated with $G = 1$.

Figure 5: CDS Spreads in Basis Points





Notes: The figures are based on data with a 30-minute sampling frequency. Our split into the pre- and the crisis period is indicated by the vertical line in each figure.

We find that the central European countries in our sample were much less affected by shocks compared to the GIIPS countries during the euro area sovereign debt crisis. We do not find differences in the impulse responses for central European euro area member countries (Austria and Slovakia) and non-euro member countries (see Figure 6, lower panel). Interestingly, we also do not find a difference in the response functions according to the debt-to-GDP levels of central European countries. The level of response for the central European countries (see Figure 6, lower panel) is almost identical. We would have expected a stronger response to shocks in central European countries with higher debt levels, such as Hungary (see Table 1).

Claeys and Vasicek (2014) are in line with our results, as they also find substantial contagion effects only between countries that were most affected by the euro area sovereign debt crisis.

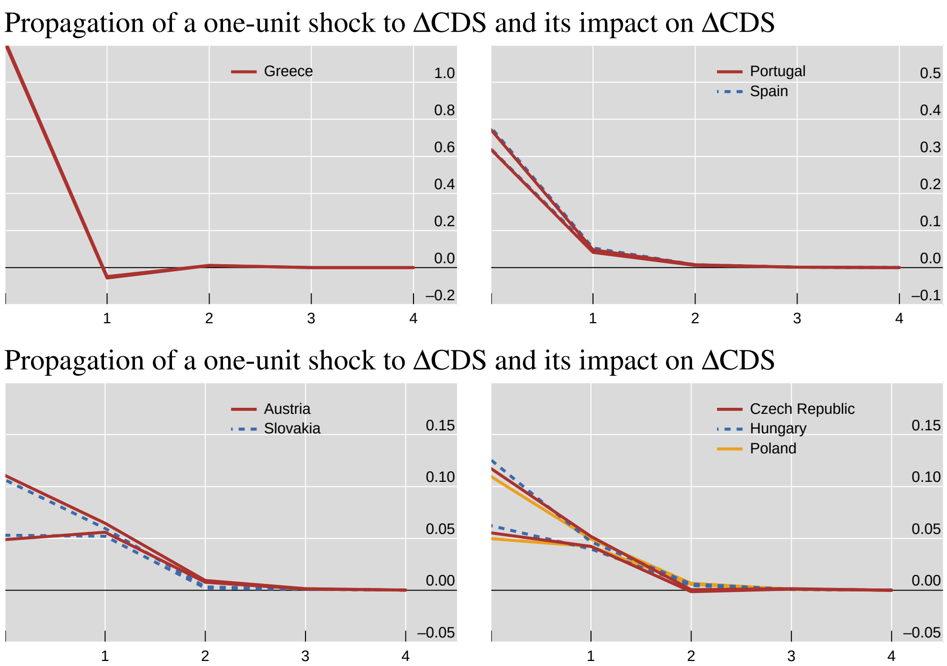
This leads to the conclusion that countries that lie geographically outside the crisis region are dramatically less sensitive to shocks propagating from the euro zone crisis regions.

Table 1: Debt-to-GDP Levels in Percent, Market Adjusted

year	Austria	Czech Republic	Hungary	Poland	Slovakia
2008	67.4	27.2	65.5	44.0	29.2
2009	77.7	32.2	77.2	49.1	36.6
2010	88.3	39.1	82.6	53.1	41.6
2011	87.1	40.8	80.7	54.9	44.9
2012	91.8	47.6	79.4	54.8	54.4
2013	91.4	49.8	82.2	56.4	57.3
2014	93.3	49.2	86.7	49.2	56.6

Source: National Data, Authors' Calculations

Figure 6: Impulse Responses in Central European Countries in the Crisis Period - Shock in Greece



Notes: This figure illustrates the impulse response of central European countries for ΔCDS to a one-unit shock (increase) in Greece for the period from 20 October 2009 to end-December 2011. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the impulse response to a one-unit shock to Greek ΔCDS. The number of 30-minute time intervals is described by the x-axis. For each impulse response we plot the upper and lower 95% confidence bands.

The speed of shock absorption is similar to that found for the GIIPS countries in our bivariate PVAR model discussed in Section 4.1.

Further impulse responses to shocks to Portuguese and Spanish CDS and their impact on central European countries can be found in Appendix B.

4.3 The Impact of Unexpected Macroeconomic News on Sovereign Credit Risk: Results from a PVARX Experiment

In this section, we conduct an experiment with the aim of analysing whether responses to shocks and shock contagion can be attributed to economic fundamentals or if overreactions in credit risk during the crisis period might also be due to self-fulfilling prophecies. For Constancio (2012) and Forbes (2012), contagion occurs when financial or macroeconomic imbalances (shocks) create a systemic risk beyond that explained by economic fundamentals. Contagion differs from macroeconomic interdependence among countries in that transmission of risk to other countries is different under normal economic conditions. Gibson et al. (2012) explain the effect of self-fulfilling prophecies by interest rate spreads that were lower than justified by fundamentals prior to the crisis, owing to the role played by Greece's euro area membership in biasing investor expectations. During the crisis period, Gibson et al. (2012) define this self-fulfilling prophecy effect that interest rate spreads were higher than those predicted by fundamentals in terms of the market's disbelief that sustainable financial consolidation measures and structural reforms would be implemented.

Our experiment is designed in a similar way to that of Canova and Ciccarelli (2009), as follows: we distribute the data of the economic surprise index over each trading day (18 time intervals). The distribution is chosen such that the maximum is reached at noon, and the sum of the 18 different intraday values is equal to the value reported by the Citigroup economic surprise index. We experimented with different distributions and, despite the arbitrary distribution assumption, we found robust results. Next, the last seven values are removed from all time series in order to be close to the last maximum (in the case of a positive reading of the surprise index) or close to the last minimum (in the case of a negative reading of the surprise index). We then fit the PVARX model from Equation (3.6) and produce an out-of-sample forecast for eight intervals beyond the last data point,[15] which is in the case of the pre-crisis period 15 October 2009 and in the case of

[15] We have chosen the forecast length of eight intervals in order to be slightly longer than the number of removed values (seven). We experimented with different values and found that the qualitative results remained robust. Again, the choice to remove the last seven values is motivated to be close to the last maximum/minimum of the surprise index, reached at midday by construction.

the crisis period 26 May 2011[16]. We call this forecast the “real forecast". Further, we repeat this same procedure, but now set the data of the surprise index of the last day to zero, i.e. we artificially remove the last positive or negative “shock" given by the data. We again produce an out-of-sample forecast, which we call the “counterfactual forecast". The difference between the real and the counterfactual forecast captures the impact of the positive or negative values of the Citigroup economic surprise index on the last day. In other words, the experiment mimics what would have happened if the last positive or negative economic news had not occurred and thus helps answer the question of whether macroeconomic fundamental news can explain changes in sovereign credit risk.

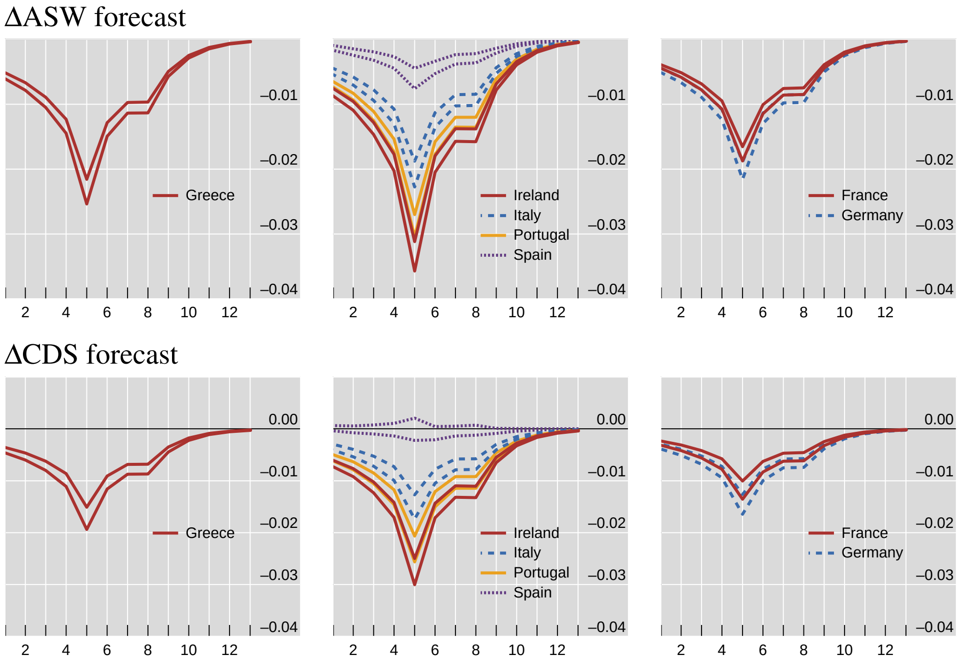
During the pre-crisis period, we find for all countries in the sample that a positive (negative) shock from the economic surprise index on the last day (15 October 2009) leads to an expected decrease (increase) in credit risk (see Figure 7). Prior to the crisis, the magnitude of the effect following an unexpected macroeconomic news shock is similar in the bond and CDS markets. Our pre-crisis results indicate that markets reacted to macroeconomic news in pricing sovereign credit risk.

During the euro area sovereign debt crisis period, a negative reading of the economic surprise index on the last available day (26 May 2011) leads surprisingly to a decrease in credit spreads in most countries (see Figure 8). In rational markets, a negative economic news shock should lead to an increase in sovereign credit risk and thus to an increase in spreads. Our results are counterintuitive, unlike those for the pre-crisis period. For the crisis period, they show that credit markets were driven not by macroeconomic news, but most likely by monetary policy, political decisions and speculations. Figure 9 displays the individual components (the real and the counterfactual forecasts) of our unexpected economic news shock experiment. Subtracting the counterfactual forecast in row 2 from the real forecast in row 1 of Figure 9 produces the forecast in row 1 of Figure 8. The same applies to the remaining rows in Figures 8 and 9. Surprisingly, in most cases a negative economic shock leads to a tightening of credit spreads (row 1 and 3 of Figure 9).

The shapes of the curves in Figures 7, 8 and 9 are due to our particular choice of decomposing the daily Citigroup economic surprise index into intraday intervals. However, other choices leading to different shapes of our curves do not change the results. This gives support to the self-fulfilling crisis theory, that changes in sovereign credit risk during the euro area sovereign debt crisis were only partially driven by economic fundamentals, as markets did not react to economic news in contrast to the pre-crisis period.

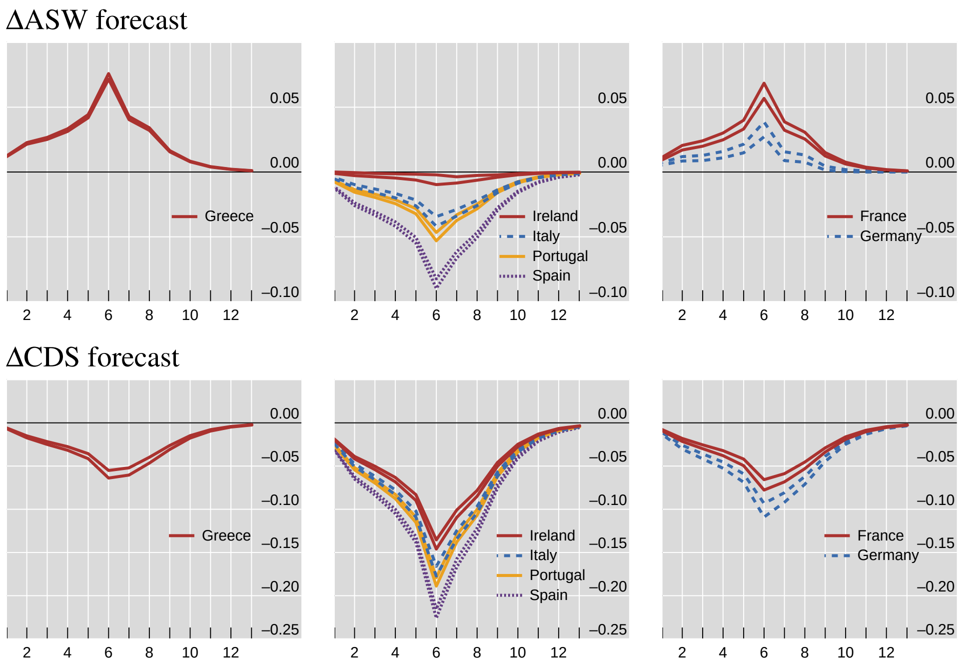
[16] We chose end-May as the last time stamp in our experiment for the crisis period because the liquidity in the Greek bond market deteriorates. The lack of pricing data from May onwards does not allow to generate a sensible intraday forecast for our experiment.

Figure 7: Positive Shock to the Economic Surprise Index During the Pre-crisis Period



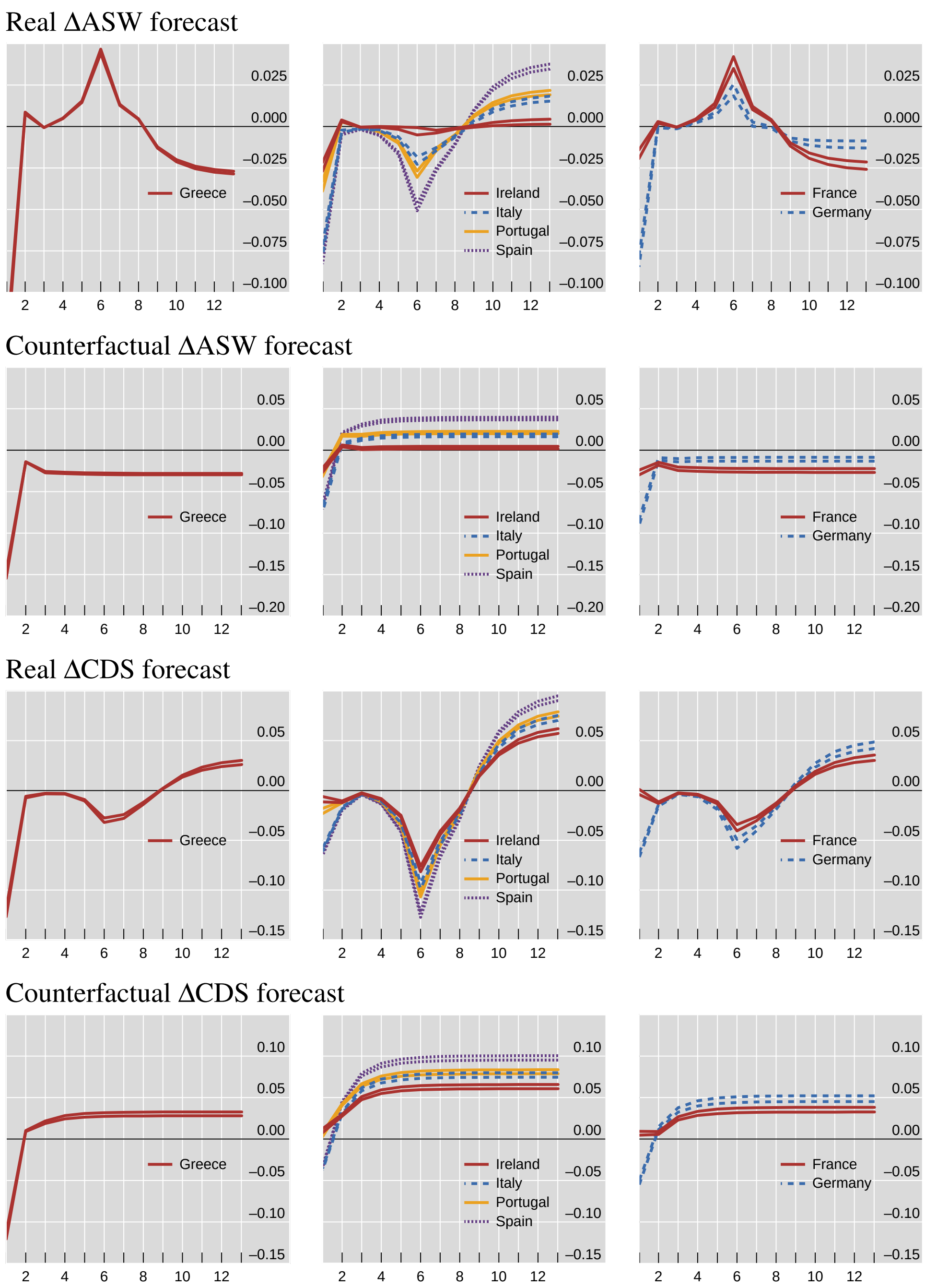
Notes: This figure illustrates a scenario of a real positive shock to the economic surprise index minus a counterfactual scenario where we assumed that the shock did not happen. The period under consideration is from January 2008 until 15 October 2009. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the response of ΔASW (upper part) and ΔCDS (lower part) in basis points. The number of 30-minute time intervals is described by the x-axis. We plot the upper and lower 95% confidence bands for each country.

Figure 8: Negative Shock to the Economic Surprise Index During the Crisis Period



Notes: This figure illustrates a scenario of a real negative shock to the economic surprise index minus a counterfactual scenario where we assumed that the shock did not happen. The period under consideration is from 20 October 2009 until end-May 2011. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the response of ΔASW (upper part) and ΔCDS (lower part) in basis points. The number of 30-minute time intervals is described by the x-axis. We plot the upper and lower 95% confidence bands for each country.

Figure 9: Real and Counterfactual Forecast Decomposition



Notes: This figure presents the individual components, i.e. the real and the counterfactual forecasts, of our experiment for the crisis period 20 October 2009 until end-May 2011. The figures are based on 5-year tenor data with a 30-minute sampling frequency. The y-axis represents the response of ΔASW (upper part) and ΔCDS (lower part) in basis points. The number of 30-minute time intervals is described by the x-axis. We plot the upper and lower 95% confidence bands for each country.

5. Conclusion

The CDS market was the main venue for the transmission of sovereign credit risk contagion during the euro area sovereign debt crisis. In contrast, we find that, prior to the crisis, the two markets (CDS and bond) were similarly important in the transmission of financial contagion, while the importance of the bond market decreased relative to the CDS market during the crisis period. We find evidence for sovereign credit risk contagion during the euro area sovereign debt crisis period, as our results show more drastic reactions to shocks in terms of magnitude and absorption compared to the pre-crisis period. Thus, our results on the responses to sovereign credit risk shocks during the crisis period confirm the contagion across euro area countries, as they result from extreme negative, systemic effects and are much larger in magnitude compared to the pre-crisis period, a fact which cannot be explained by macroeconomic fundamentals.[17] We find comovement effects rather than contagion during the pre-crisis period, as markets react rationally to economic fundamentals, while the responses to sovereign credit risk shocks remain moderate in magnitude. The use of intraday data substantially increases the precision of the results, as we find average timelines of financial shock contagion of one to two hours during the crisis period and 30 minutes to one hour prior to the crisis. This is a clear indication of the efficiency of financial markets.

We find a flight to safety during the crisis period in the German bond market. This is not present prior to the crisis and, interestingly, is also not visible in the French bond market. The flight-to-safety effect can be explained by market participants' lack of belief in the future path of public finances (a self-fulfilling crisis), which cannot be explained by macroeconomic news.

Our results using an unexpected exogenous macroeconomic news shock suggest that, during the pre-crisis period, markets for sovereign credit risk were driven by macroeconomic news. Positive news led to a decrease in credit spreads and negative news to an increase. Using the same experiment for the euro area sovereign debt crisis period, our results show that movements in sovereign credit spreads did not respond to macroeconomic news but were rather driven by either monetary policy or exaggerations in financial markets due to lack of belief (a self-fulfilling crisis).

We find that central European countries were practically unaffected by sovereign risk contagion during the crisis. Our model further indicates no difference in the responses to shocks according to debt levels or whether the country belongs to the monetary union or

[17] See the contagion definitions according to Constancio (2012), Forbes (2012), and Kaminsky et al. (2003) in the Introduction.

not. This implies that, in general, countries that lie geographically outside of the crisis region were much less affected by sovereign risk contagion.

As stated by Gyntelberg et al. (2013), the fact that CDS premia are more responsive to new information may reflect the fact that the market participants in these markets on average are more highly leveraged, are more aggressive in taking positions and hence respond more quickly to new information. Thus it is crucial for policy makers and regulators to understand the dynamics in the market for sovereign credit risk, especially in the derivative market, where contagion effects are more severe during our analysed crisis sample.

Even though policy makers may not be interested in intraday movements in credit risk, our results show that the level impacts from the short-term dynamics are persistent (see Appendix D). Hence, our results are important with regard to financial stability.

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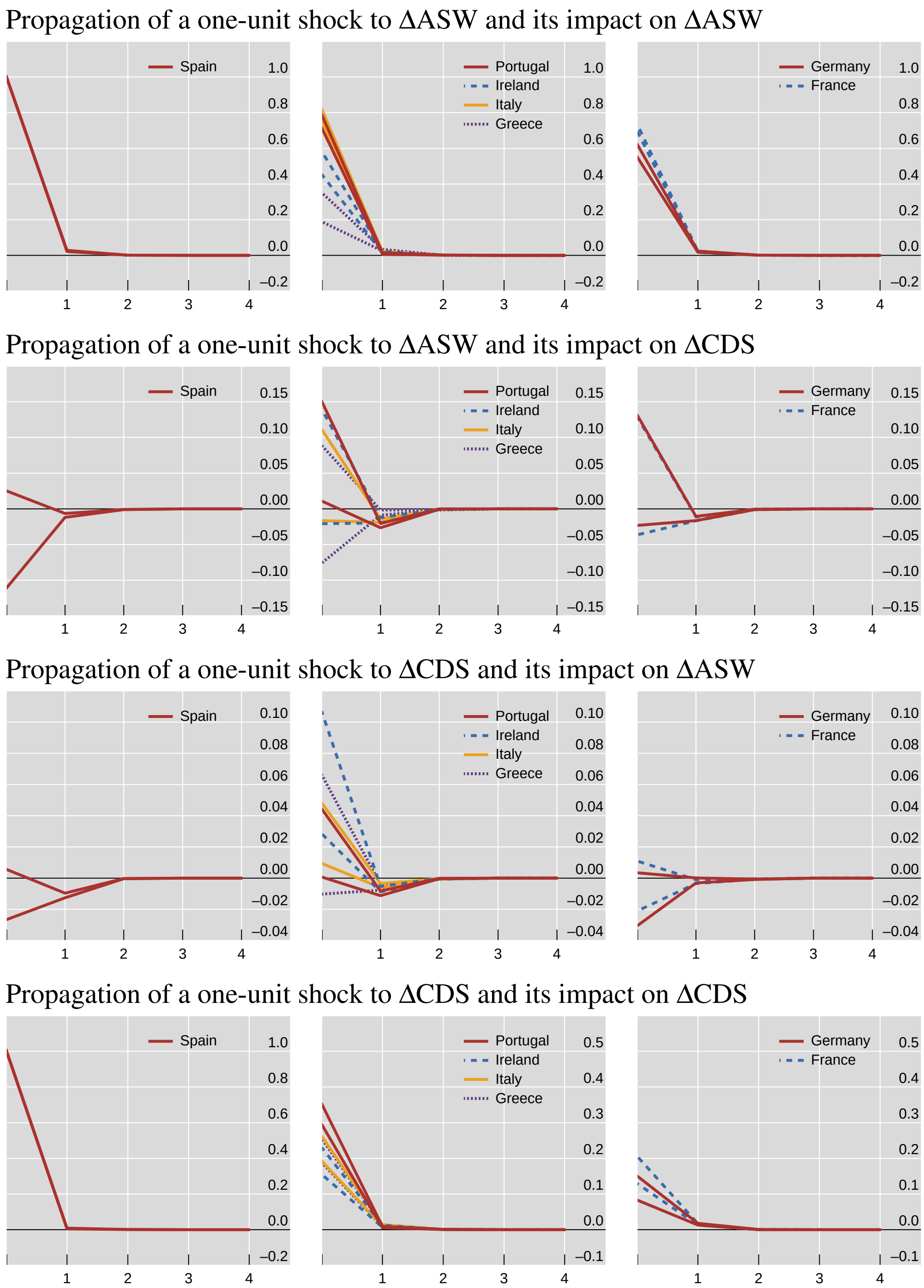
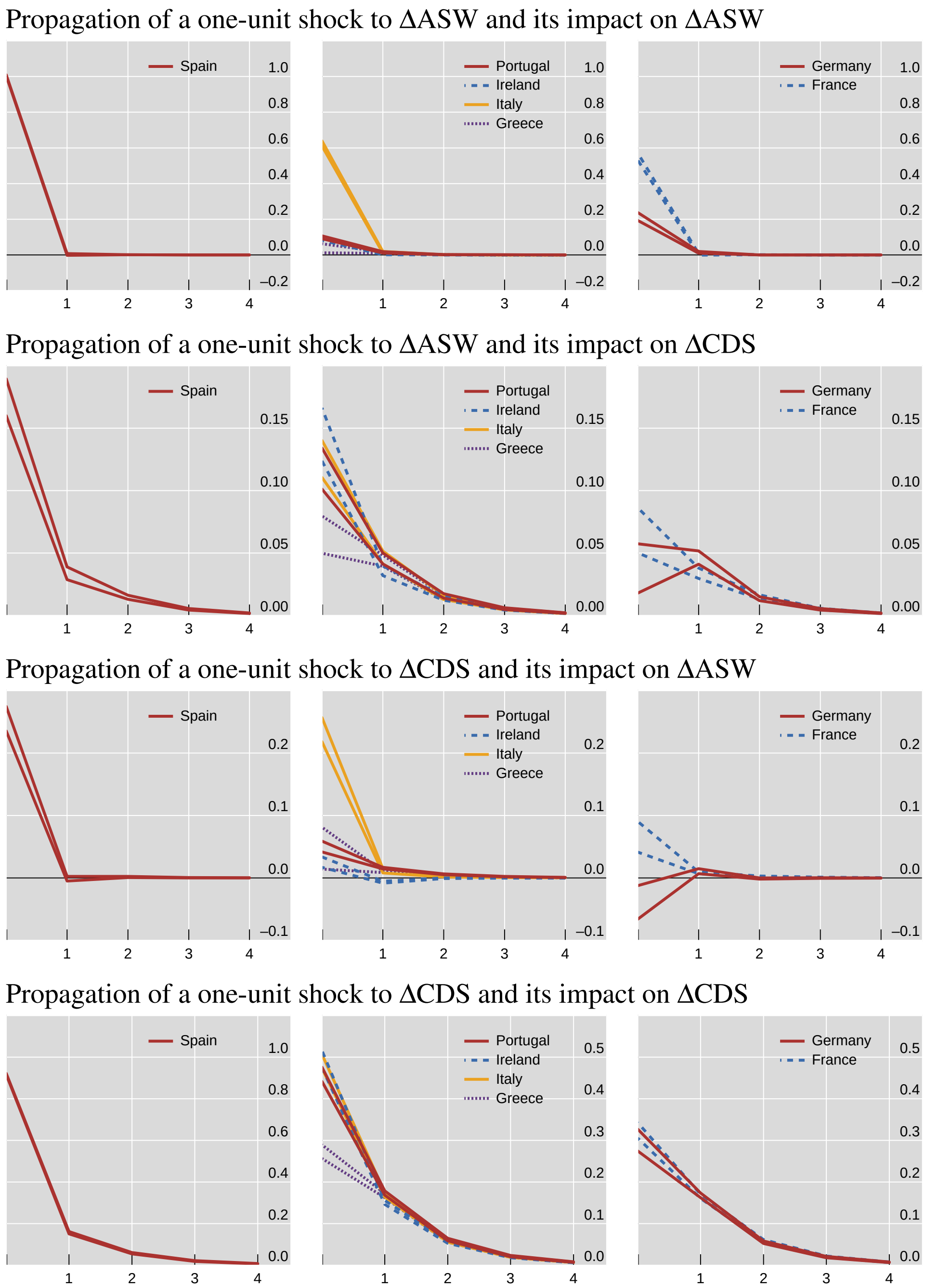
Appendix A: Impulse Response Functions for Spain and Portugal

Figure A.1: Impulse Responses in the Pre-crisis Period - Shock in Spain



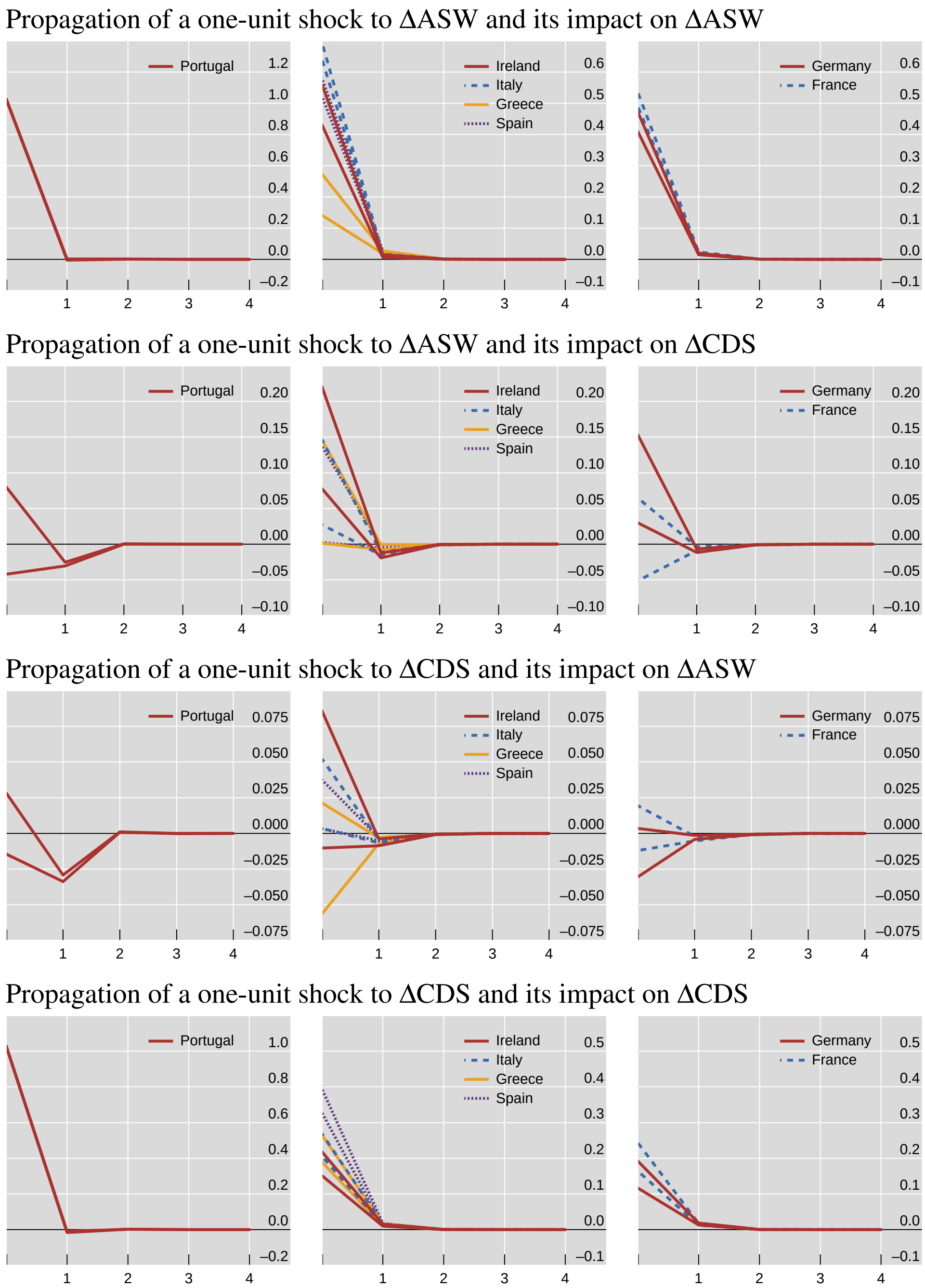
Notes: For details see Figure 3.

Figure A.2: Impulse Responses in the Crisis Period - Shock in Spain



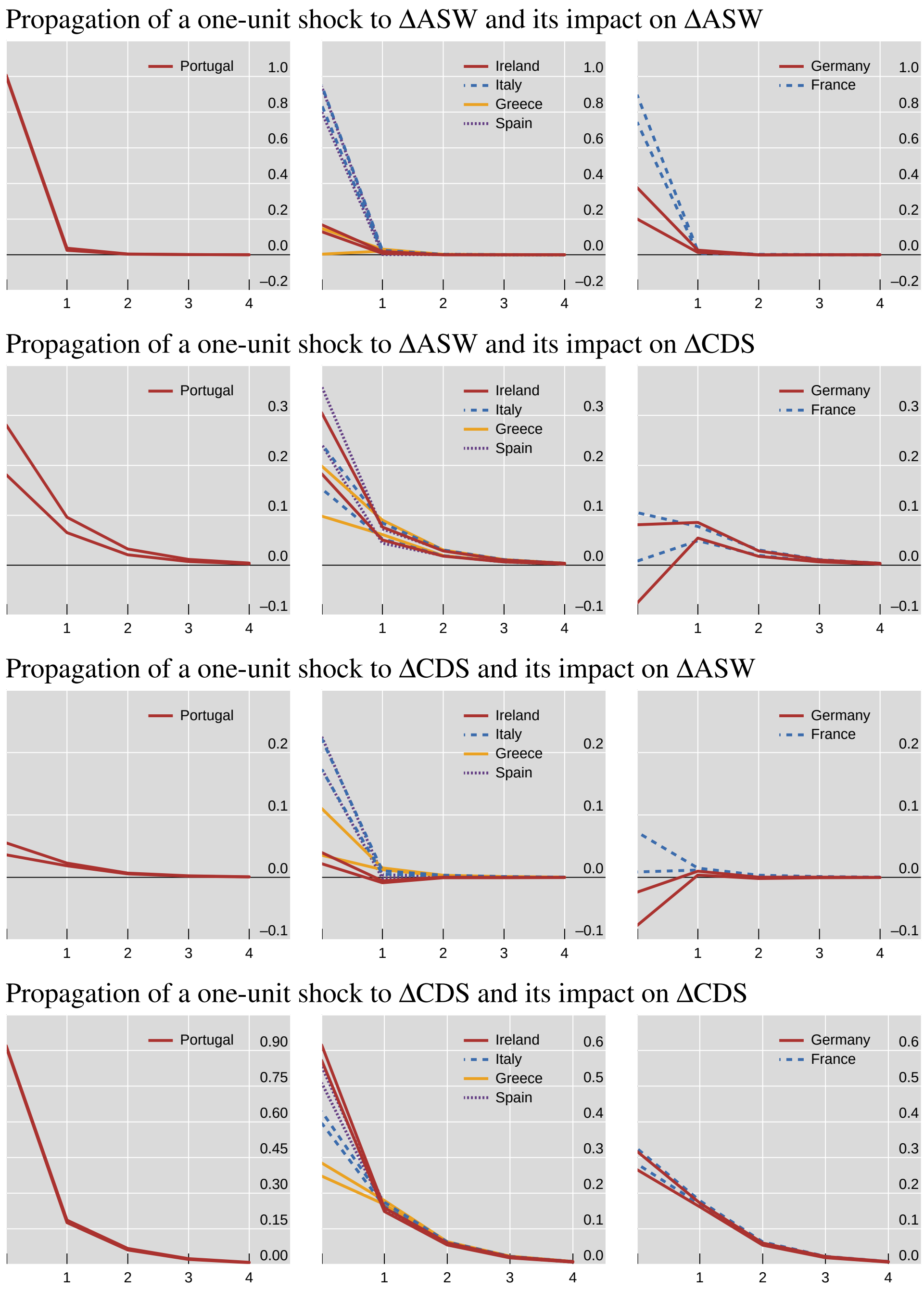
Notes: For details see Figure 4.

Figure A.3: Impulse Responses in the Pre-crisis Period - Shock in Portugal



Notes: For details see Figure 3.

Figure A.4: Impulse Responses in the Crisis Period - Shock in Portugal

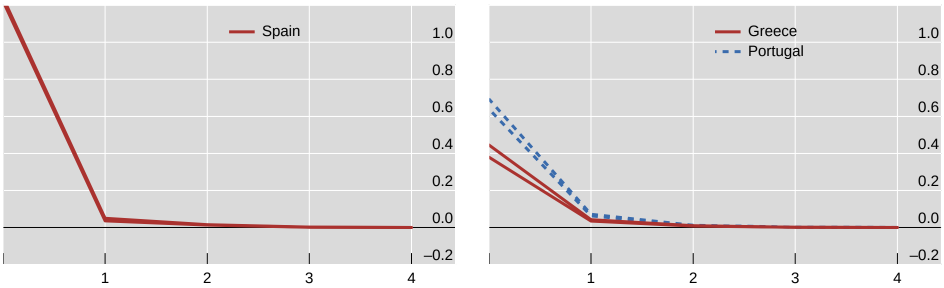


Notes: For details see Figure 4.

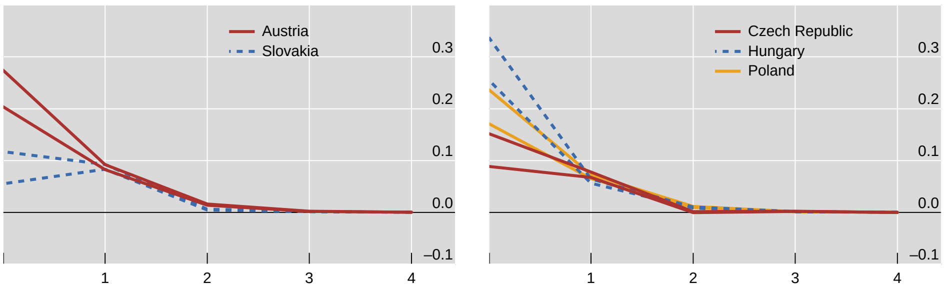
Appendix B: Impulse Response Functions of Central European Countries

Figure B.1: Impulse Responses in the Crisis Period - Responses of Central European Countries

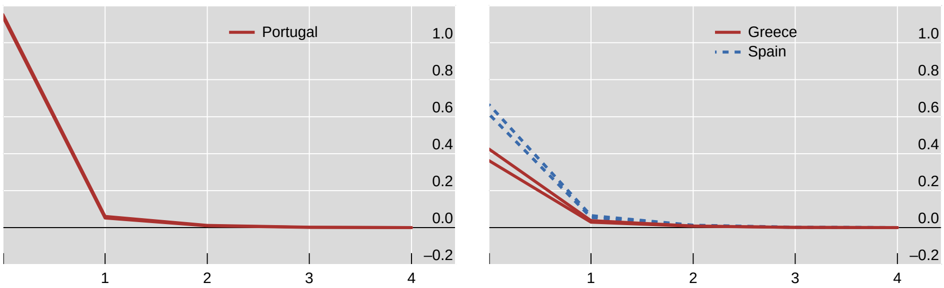
Propagation of a one-unit shock to Spanish ΔCDS and its response in ΔCDS



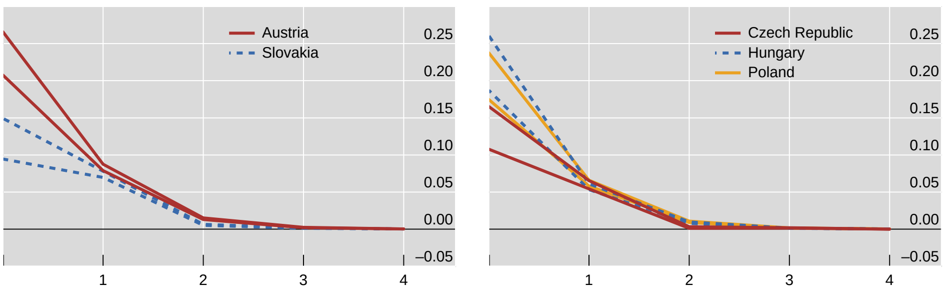
Propagation of a one-unit shock to Spanish ΔCDS and its response in ΔCDS



Propagation of a one-unit shock to Portuguese ΔCDS and its response in ΔCDS



Propagation of a one-unit shock to Portuguese ΔCDS and its response in ΔCDS



Notes: We present the responses of central European countries to a one-unit shock to ΔCDS in Spain and Portugal. For details see Figure 6.

Appendix C: Panel Unit Root Tests

Before analysing contagion effects within a panel framework, we perform unit root and stationarity tests on our CDS and ASW price data. Canova and Ciccarelli (2013) suggest that panel-based unit root or stationarity tests have a higher power than univariate tests. For our ASW and CDS data, we cannot reject the $\mathscr{H}_0$ of a common unit root according to the Levin, Lin-, and Chu test. Further, we also can not reject the $\mathscr{H}_0$ of individual unit root processes according to the Im, Pesaran and Shin panel unit root test for our data (see Table C.1). Since all of our country series are considered simultaneously and our data for CDS and ASW are non-stationary (I(1)), we use first differences for our model estimations.

Our panel unit root test takes the following form:

$$\Delta y_{it} = \alpha_i + \rho_i y_{i,t-1} + u_{it} \qquad \text{with} \qquad \mathscr{H}_0 : \rho_1 = ... = \rho_N = 0,$$

where $i = 1,...,N$ is the cross-sectional dimension and $t = 1,...,T$ is the time-series dimension. Hence, all series are independent random walks under the $\mathscr{H}_0$ and non-stationary.

We perform the Levin, Lin, and Chu test, which assumes a common unit root process where the homogenous alternative takes the following form:

$$\mathscr{H}_{1a} : \rho_1 = ... = \rho_N = \rho < 0,$$

where all series are stationary under the $\mathscr{H}_1$.

Further, we perform individual panel unit root tests based on Im, Pesaran and Shin where the heterogeneous alternative takes the following form:

$$\mathscr{H}_{1b} : \rho_1 < 0, ..., \rho_{N_0} < 0, \qquad \text{where} \qquad N_0 \leq N.$$

Hence, $N_0 \leq N$ series are stationary, with potentially different AR parameters.

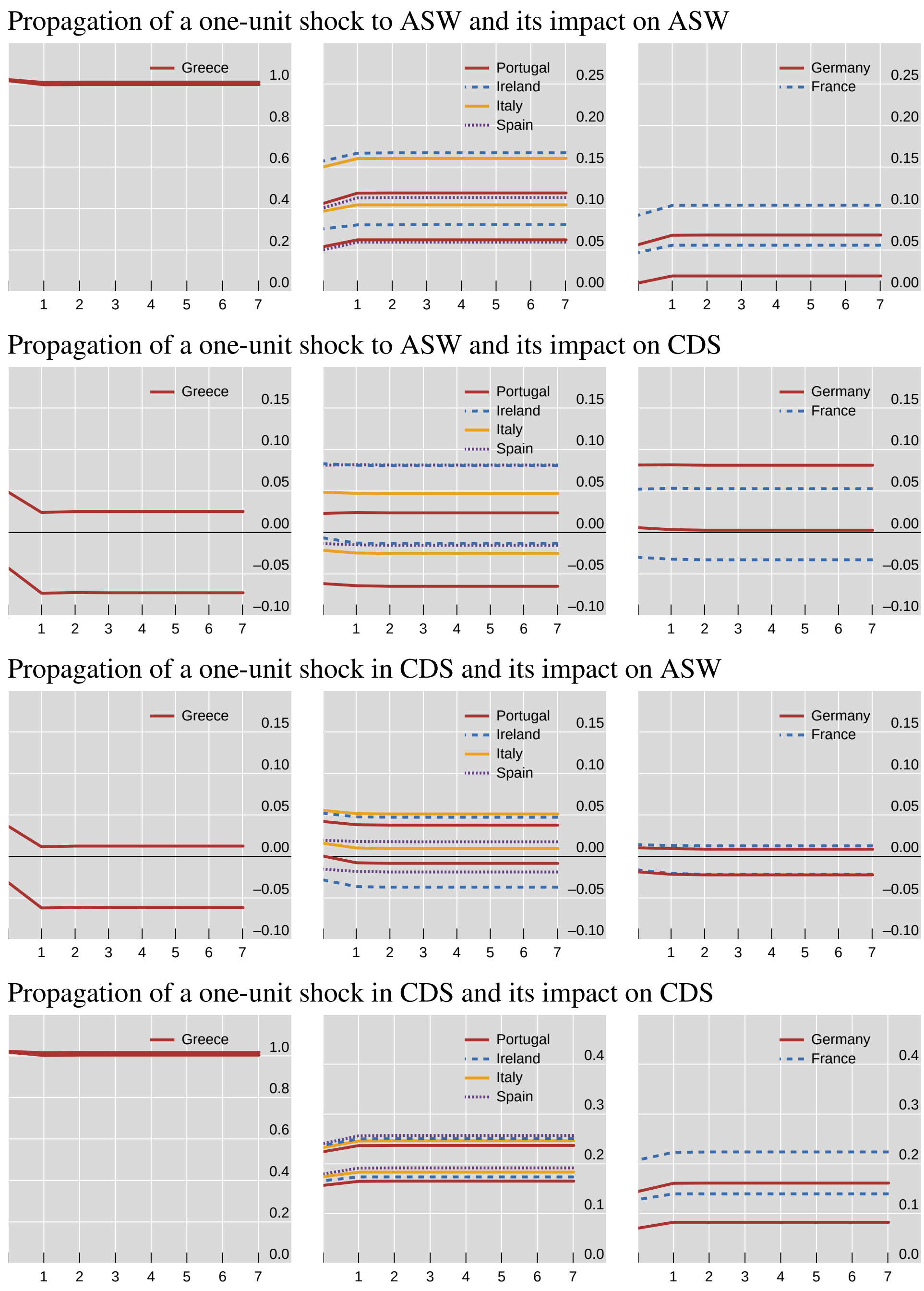
Table C.1: Panel Unit Root Tests - p-values

	Levin, Lin, Chu	Im, Pesaran, Shin
ASW	1.00	0.23
CDS	0.59	0.15

Notes: This table reports the p-values of the panel unit root test for ASW and CDS with individual intercepts for the period from January 2008 to end-December 2011 and 30-minute sampling frequency. The cross-section consist of the seven countries in our sample.

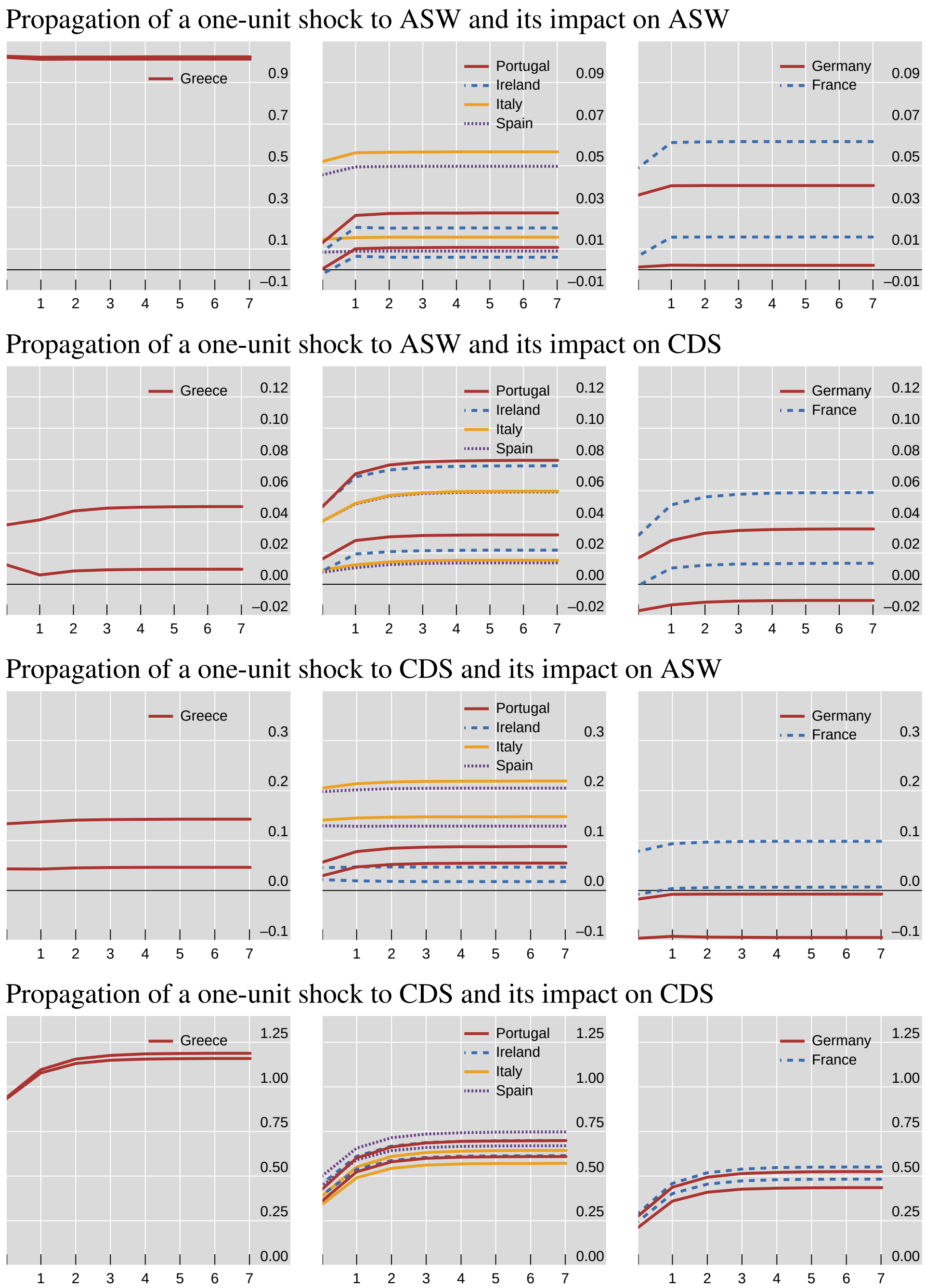
Appendix D: Accumulated Impulse Response Functions for Greece

Figure D.1: Accumulated Impulse Responses in the Pre-crisis Period - Shock in Greece



Notes: This figure illustrates the accumulated impulse response for CDS and ASW to a one-unit shock (increase) for the period from January 2008 to 19 October 2009. For further details see Figure 3.

Figure D.2: Accumulated Impulse Responses in the Crisis Period - Shock in Greece



Notes: This figure illustrates the accumulated impulse response for CDS and ASW to a one-unit shock (increase) for the period from 20 October 2009 to end-December 2011. For further details see Figure 4.

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