Euro area sovereign yield spreads as determinants of private sector borrowing costs

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Abstract

We regress long-term private-sector borrowing rates on a money market rate, a term premium and credit risk. As a contribution to the current debate about European safe assets, our interest is in quantifying the impact of euro area sovereign bond spreads on private-sector lending by employing it as a proxy for private-sector credit risk. Panel estimates show significant, albeit rather small long-run effects. Another finding is large cross-country heterogeneity. Using linear country-specific estimates, we find the effect to be significant in only some countries, but the size of the maximum effect in these countries exceeds the average one more than three-fold. Furthermore, for one country, we find an asymmetrical effect with positive spread changes having greater impact on private-sector borrowing costs than negative ones. Substantial heterogeneity of the spillover effect between euro area countries indicates the presence of financial valuation effects based not only on economic fundamentals. This, in turn, implies that spillovers may entail contagion costs. Overall, our results suggest that these costs are considerable in the euro area and will remain so until an effective form of European safe assets is created.

Keywords: autoregressive distributed lag, composite cost of borrowing, sovereign spread

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1 Introduction

Same-maturity bonds of different issuers are often remunerated with different rates of interest. A key factor for this difference is the credit risk premium that reflects the expectations and uncertainty of financial markets with regard to the fulfillment of debt service obligations. Government bonds of developed economies are generally considered to be virtually risk-free in this respect, because they are ultimately backed by the sovereign's ability to raise funds by taxing private income and wealth. Furthermore, the central bank of developed economies generally acts as a lender of last resort so that a mere shortage of liquidity need not translate into higher sovereign bond yields. Given their ability to efficiently tax private income and wealth, euro area countries qualify as a developed economy whose sovereign bonds generally serve as safe assets and benchmarks for the national economy.[1] Uncertainty in financial markets increased in the wake of the international financial crisis with regard to both the liquidity provision by the European Central Bank and emerging prospects of sovereign debt restructuring (DeGrauwe and Ji, 2013). As a result, several euro countries experienced a substantial increase in their sovereign risk premia. This gave rise to a vicious cycle of rising public borrowing costs, plunging economic activity, capital flight and further soaring risk premia. The role of sovereign yields as benchmarks for private lending rates amplified the economic contraction (Iorgova et al., 2012; Gorton, 2017). Consequently, the surge in sovereign risk premia likely weighed on GDP not only through cuts in government spending, but also by increasing the costs of private investment (Figure 1).

In this paper, we aim to quantify the benchmark role of sovereign bond yields for private sector borrowing rates. We supplement studies investigating the determinants of sovereign yield spreads on the basis of financial market valuation (Bernoth and Erdogan, 2012; Caggiano and Greco, 2012; DeGrauwe and Ji, 2013; Afonso et al., 2014; Dewachter et al., 2015; Kinateder and Wagner, 2017). At the same time, we build on a strand of literature examining the extent to which safe asset constructions limit excessive variation in the valuation function of financial markets (Lane, 2012; Brunnermeier et al., 2016). This is particularly relevant as sovereign spreads not only influence public but also private sector financing conditions (Iorgova et al., 2012; Gorton, 2017). We therefore estimate a simple interest rate model that incorporates sovereign bond yield spreads as one proxy for credit risk and in this way captures potential spill-overs of risk-related changes in government funding costs on private sector funding costs. At the macroeconomic level, the magnitude of such spill-over effects is by no means clear-cut. If specific information

[1] For a general assessment see Gorton (2017). Some observers argue that Greece is an exception, but the available figures indicate an improvement in taxation (Georgakopoulos, 2016). In our analysis, we exclude Greece for data availability reasons.

about a private sector debtor is available, the credit-risk premium demanded by the creditor should in principal be determined by this specific information. Country risk as an additional factor or as a substitute is meaningful only to the extent that it correlates with the earnings and repayment prospects of the respective debtor companies or households.

If sovereign bond spreads are found to play a significant role as proxy for private sector credit risk at the macroeconomic level, it is worthwhile analyzing the spill-overs in the light of financial market valuation variability which is found to play an essential role as a driver of sovereign bond spread changes (Caggiano and Greco, 2012; DeGrauwe and Ji, 2013; Afonso et al., 2014; Dewachter et al., 2015; Kinateder and Wagner, 2017). This is the second objective of the paper at hand. Our approach to identifying variability in financial market valuation is twofold: On the one hand, we investigate the variability of the spillover-effect in terms of cross-country heterogeneity and the potential asymmetry of positive and negative spread changes. On the other hand, we compare the impact coefficients of sovereign bond yield spreads on private sector borrowing rates relative to the impact of the unemployment rate which in our model proxies credit risk not subject to financial market valuation. Substantial variability in the coefficients can point towards the existence of financial market valuation biases so that spillovers may entail contagion costs, as otherwise the financing conditions for private sector real economic activity would be more favorable.

Our results can be summarized as follows: Panel estimates show a significant, albeit rather small long-run spillover-effect. We find evidence for cross-country heterogeneity. Using linear country-specific estimates, the spillover-effect is found to be significant only in some countries, but the size of the maximum effect in these countries exceeds the average one more than three-fold. Furthermore, for one country, we find an asymmetrical effect with positive spread changes having greater impact on private-sector borrowing costs than negative ones. Our results contribute to the controversial debate about European Safe Bonds, which mainly focuses on preserving the general capacity of euro area countries to withstand negative macroeconomic and financial shocks (Lane, 2012; Brunnermeier et al., 2016). The linkage estimated in this paper indirectly allows for an evaluation of the very specific potential of safe sovereign assets to limit spillover-effects that negatively impact the economy, first and foremost, private sector investment.

The remainder of the paper is structured as follows. Section 2 reviews selected literature. Section 3 discusses the data and estimation methodology. Section 4 presents the empirical results and Section 5 concludes.

2 Related Literature

The first studies to explain yield differentials of European government bonds after the introduction of the euro in 1999 focus on fundamental national and international determinants. Codogno et al. (2003) attribute the more pronounced yield differential movements of Italian and Spanish bonds relative to other euro area countries in their sample to fiscal vulnerability. They emphasize the further need for convergence of government debt-to-GDP ratios. Hallerberg and Wolff (2008) underline the role of the government budget deficit. According to their analysis, credible deficit signals, especially (rule-based) targets, play a crucial role in reducing the risk premium demanded by financial markets. Reinhart and Rogoff (2011) find that banking crises often precede sovereign debt crises. Here, the presumed transmission channel runs along the fundamentals, from the cost of bank resolution to an increased government debt-to-GDP ratio.

Recent studies underline that, apart from actual changes in the fundamentals, the valuation function in financial markets may vary. Dewachter et al. (2015) estimate European government bond rates in a vector autogressive model of macroeconomic and financial variables. While they basically confirm a dominant role of fundamental factors for spread movements over the period 2005 - 2013, their historical decomposition also shows that the explanatory power of the non-fundamental component has increased since 2011. This component should reflect to a large extent the pricing of risk by market participants. DeGrauwe and Ji (2013) investigate this role more explicitly. They find evidence that "a significant part of the surge in the spreads of the peripheral Eurozone countries during 2010-2011 was disconnected from underlying increases in the debt-to-GDP ratios and fiscal space variables, and was associated with negative self-fulfilling market sentiments". Correspondingly, the coefficients, Caggiano and Greco (2012) estimate for the influence of fiscal and financial variables in ten euro area countries after 2007, are up to four times higher than those before the financial market crisis. Moreover, the statistical significance they find for the interaction terms between fiscal and financial variables points to strong nonlinearities in the financial market valuation of fundamental data.

Similarly, in a time-varying coefficient fixed-effects panel model, Bernoth and Erdogan (2012) find that the impact of fiscal variables and a global risk factor on euro area yield differentials varies considerably over time. They attribute the strong increases of sovereign bond yield spreads after the financial crisis inter alia to a change in the investor's risk aversion. The variability of financial market valuations also plays a central role in studies that estimate the determinants of government bond spreads using non-observable time-varying common factors. Afonso et al. (2014) conclude from an increase in their first principal component that the number of 'risk factors priced by markets has been significantly enriched since March 2009, including international risk, liquidity risk and the risk of the crisis' transmission among EMU member states.' The study by Kinateder and Wagner (2017) explores whether conventional and unconventional monetary policy measures have a decreasing effect on sovereign bond yield spreads. The authors confirm this effect on the basis of widely differing elasticities of the monetary policy variables before and during the crisis, making their results particularly relevant for studies with sample period containing data after 2012.

In this paper, we link the findings about financial market valuation variability to the benchmark function of government bonds for the pricing of other assets. Gorton (2017) comprehensively describes the need for safe assets in an economy. As pointed out by Iorgova et al. (2012), the government yield curve is a traditional benchmark for the pricing and valuation of risky assets. Given that there are neither common sovereign debt instruments nor homogeneous euro area sovereign yield curves, virtually risk-free sovereign bonds such as German or, in our case, French bonds are used as an alternative benchmark (Iorgova et al., 2012). The spread between this benchmark and the government bond rate of a given euro area country serves as a proxy for credit risk.

3 Methodology and Data

To estimate the impact of sovereign yield spreads on private-sector borrowing rates, we decompose long-term composite costs of borrowing (CCB) into (i) a short-term money market rate (EONIA), (ii) a term premium (TERM) capturing the expected future changes in short-term interest and inflation rates as well as the corresponding risk premia (interest-rate risk, inflation risk) and (iii) credit-risk premia (CREDIT). Including an error term, we obtain

$$CCB_{t-1} = \beta^{E} EONIA_{t-1} + \beta^{T} TERM_{t-1} + \beta^{C} CREDIT_{t-1} + u_{t-1} \tag{1}$$

Two econometric issues have to be taken into account. First, we should consider lagged values of the dependent and independent variables to avoid omitted variable bias. This directly leads us to autoregressive distributed lag models, ARDL$[p, q]$, where p denotes the maximum number of lags of the dependent variable and q the maximum number of lags of the independent variables. Second, if the variables in Equation (1) are co-integrated, a level estimator should be modified to avoid a spurious relationship between the interest rates.[2] Moreover, if a co-integration relation

[2]See for instance Phillips and Hansen (1990) suggesting to apply fully modified ordinary least squares (FMOLS).

between the variables is comprehensively tested and identified, ARDL models can be transformed to error-correction form. This is useful for estimating both short-term and long-term effects of the explanatory variables, in particular the sovereign bond spread, on private sector borrowing rates. The term *error* relates to the fact that last-period's deviation from the long-run equilibrium, the error u in Equation (1), influences the first difference of the dependent variable, i.e.

$$\begin{aligned}\Delta CCB_t = \alpha + \rho\Big(CCB_{t-1} - (\beta^E EONIA_{t-1} + \beta^T TERM_{t-1} + \beta^C CREDIT_{t-1})\Big) \\ + \phi^{CCB}\Delta CCB_{t-1} + \phi^E \Delta EONIA_t + \phi^T \Delta TERM_t + \phi^C \Delta CREDIT_t + \ldots + \epsilon_t,\end{aligned} \tag{2}$$

where the first line displays the deviation from the long-run equilibrium and the second line the short-term dynamics. The term *correction* relates to the fact that these deviations should be corrected, i.e. the coefficient ρ, the so-called speed of adjustment, should be negative, if the ARDL model is well designed.[3]

As two proxies for the country-specific credit risk premia, we employ a specific sovereign bond spread, GS5Y, as well as the unemployment rate, U, i.e. $CREDIT = (GS5Y, U)$. In the ARDL model, the overall effect of the sovereign bond spreads on private sector borrowing costs can then be described by the interaction of long-run and short-run coefficients. This is captured by the dynamic multiplier

$$m_h^k = \sum_{j=0}^{h} \frac{\partial CCB_{t+j}}{\partial GS5Y_t}, \quad h = 0, 1, 2, \ldots \tag{3}$$

which by construction converges to the long-run coefficient β^{GS5Y}, if $h \to \infty$. For the sake of brevity, our analysis therefore primarily focuses on the long-run coefficients of different ARDL specifications. Only in the case of nonlinear specifications do we explicitly refer to the dynamic multiplier.

Our estimation strategy involves three stages. Summarizing all explanatory variables in the vector x, we start with panel data models, employ the co-integration tests by Westerlund (2007), and then apply the pooled mean group (PMGE) estimator, developed by Pesaran et al. (1999)

$$\Delta CCB_{i,t} = \alpha_i + \rho_i\left(CCB_{i,t-1} - \sum_{k=1}^{m}\beta^k x_{i,t-1}^k\right) + \sum_{j=1}^{p-1}\gamma_{i,j}\Delta CCB_{i,t-j} + \sum_{i=1}^{m}\sum_{j=0}^{q-1}\phi_{i,j}^k \Delta x_{i,t-j}^k + \epsilon_{i,t} \tag{4}$$

for different sub-samples, both in terms of time and country dimensions. We robustify the results by applying dynamic fixed effects and mean-group (MG) estimations, as is common in the

[3] If the composite costs of borrowing are greater than the sum of the explanatory interest components, this lowers the change of the composite costs. If it is lower, this increases the change of the composite costs.

literature (Blackburne and Frank, 2007). Since the MG-estimator, unlike the PMG estimator, does not require homogeneity of the long-run coefficients, we use the Hausman test to evaluate PMG as the preferred model. If we find evidence for cross-country heterogeneity, we apply a country-specific estimation strategy in the second stage.

To identify the most robust co-integration relation, we employ three different co-integration tests: (i) The bounds test developed by Pesaran et al. (2001) examines whether the ARDL model contains a level relationship between the dependent variable and the regressors using non-standard critical values as bounds that also cover the case of regressors being a mixture of I(0) and I(1). (ii) The Hansen (1992) instability test uses Lagrange multiplier theory to explore time-variation in the scores of the estimated equation. For the scores, the residuals of the level estimation are essential. Here we use fully modified ordinary least squares (FMOLS) and check whether the long run-coefficients of the ARDL are close to the FMOLS results. (iii) The Phillips and Ouliaris (1990) test examines whether the residuals of the OLS level estimation are integrated. Overall, the test results are expected to fall into three groups: for some countries, they may indicate no stable co-integration relation, for other countries the inclusion of both credit risk proxies, the sovereign bond yield spread and the unemployment rate, and, for the remaining countries the exclusion of one credit risk proxy. In cases where we do not obtain a stable co-integration relationship, we do not interpret the coefficients because the ARDL model is not well specified. Based on the test result, we estimate single equation error correction models including either one or both credit risk proxies.

$$\Delta CCB_t = \alpha + \rho\left(CCB_{t-1} - \sum_{k=1}^{m}\beta^k x_{t-1}^k\right) + \sum_{j=1}^{p-1}\gamma_j \Delta CCB_{t-j} + \sum_{i=1}^{m}\sum_{j=0}^{q-1}\phi_j^k \Delta x_{t-j}^k + \epsilon_t \tag{5}$$

using the Schwarz information criterion for the lag selection.

In the third stage, we estimate country-specific nonlinear autoregressive distributed lag (NARDL) models developed by Shin et al. (2014)

$$\begin{aligned}\Delta CCB_t = \alpha + \rho\left(CCB_{t-1} - \sum_{k=1}^{m-1}\beta^k x_{t-1}^k - \left(\beta^{*+} x_{t-1}^{*+} + \beta_j^{*-} x_{t-1}^{*-}\right)\right) + \sum_{j=1}^{p-1}\gamma_j \Delta CCB_{t-j} \\ + \sum_{i=1}^{m-1}\sum_{j=0}^{q-1}\phi_j^k \Delta x_{t-j}^k + \sum_{j=0}^{q-1}\left(\phi_j^{*+}\Delta x_{t-j}^{*+} + \phi_j^{*-}\Delta x_{t-j}^{*-}\right) + \epsilon_t\end{aligned} \tag{6}$$

to identify possible asymmetries between positive and negative changes in the credit risk variables for those countries found to have a stable co-integration relation.[4] We evaluate the cumu-

[4]More precisely, x_t^{*+} accumulates positive changes of the variable x, x_t^{*} negative changes.

lative dynamic multiplier effect on CCB after one year (12 periods).

Our data set contains monthly observations of the respective series for ten euro area countries for the period 2003:1 - 2017:12.[5] The long-term composite costs of borrowing, CCB, calculated by the European Central Bank stands for aggregate private-sector bank lending rates. Five years is the presumed average maturity of the ECB data for the CCB, as this does not only incorporate yields for mortgages with longer maturity, but also yields of corporate and consumer credit with shorter maturity. Table 1 shows descriptive statistics of the CCB for the individual countries in our sample as well as their country codes. We seek to explain the substantial heterogeneity of the national lending rates on the basis of the following data: The short-term interest rate is proxied by the EONIA money market rate (EONIA) and the term premium (TERM) by the spread between the French 5-year government bond yield and the EONIA. As proxies for the country-specific credit risk premia, we employ the spread between the respective country's sovereign bond yield and the yield on French 5-year sovereign bonds, GS5Y, as well as the unemployment rate, U. Table 2 summarizes the variables used, their data source and frequency and the expected signs in the long-run level relation embedded in the ARDL model. All time series used show a unit root according to Augmented Dickey-Fuller tests.

Figure 1 illustrates the data used as well as the GDP of the euro area countries examined, based on the nexus of our estimation to real economic activity. Sub-figure top left shows that the composite costs of borrowing was at nearly the same level in all countries until 2008. During the financial market crisis, interest rates began to diverge; these differences then dramatically amplified in the wake of the euro area crisis starting in 2010. Borrowing cost divergence declined from 2012 to mid-2014, largely as a result of monetary policies including ECB President Draghi's game-changing 'What-ever-it-Takes' speech in July 2012. Until 2017, interest rate divergence hardly changed but remained elevated compared to pre-crisis levels. Money market rates and virtually risk-free yields of German and French bonds rose more during the financial market crisis than during the euro crisis (sub-figure bottom left). Sub-figure bottom right shows that the development of and divergence between 5-year government bond yields are very similar to those of the composite costs of borrowing, the synchronization being stronger in some countries than in others. This has to do with liquidity influences, as explained in the next section in the context of our estimation results. Looking at GDP trends (sub-figure top right), major divergences between euro area countries can be observed before and after the crisis years. Discussing all interdependencies goes beyond the scope of the paper. However, for the role of euro area

[5]The sample includes ten of the original twelve euro countries. Luxembourg and Greece are not included because their government bonds with 5-year maturity lack the minimum liquidity necessary for a valid analysis.

sovereign bond spreads, developments after 2010 are particularly relevant. One can outline the edges of our data sample by focusing on the comparison of France and Portugal. At the height of the euro crisis, the risk premium on 5-year government bonds issued by Portugal exceeded France's by more than 15 percentage points in early 2012, the average spread amounting to 2.2% during the sample period. In real terms, Portugal regained its production level of 2008 only in 2018, while France already did in 2011.

Two aspects of our data and variable selection deserve special motivation. The first concerns the selection of the French bond as the benchmark for sovereign bond spreads, whereas other studies mainly use German bonds (Bernoth and Erdogan, 2012; Caggiano and Greco, 2012; Afonso et al., 2014; Kinateder and Wagner, 2017). As can be seen from Figure 1, bottom left, the differences between the two bond rates are generally small.[6] However, German sovereign bonds seem to be more often overpriced on account of safe-haven effects. The reason we prefer the French bond is that its yield exhibits fewer periods in which it is below the EONIA as this undershooting theoretically contradicts our interest rate decomposition. The second aspect is our use of the unemployment rate. We consider the unemployment rate to be a suitable macroeconomic proxy for fundamental credit that is directly observable, not subject to financial market valuation and available on a monthly basis. Unemployment affects both household incomes and the sales opportunities of enterprises. The unemployment rate therefore allows for a meaningful comparison between fundamental credit risk and credit risk subject to valuation effects as measured by the sovereign bond spread. As can be seen in Table 2, the expected sign in our interest rate model is also positive for the unemployment rate, even though the unemployment rate may not always be a good proxy credit risk in all events.[7] In principle, there is a link between our two proxies for credit risk. For example, an increase in government debt may, on the one hand, increase the sovereign bond spread. On the other hand, additional government spending can lead to lower unemployment rates. The most likely outcome in our estimation approach, if both proxies are taken into account in the co-integration relationship, would be that the total pricing of default risk would offset opposing movements in the single proxies. If the co-integration relationship is only identified as stable with one of the proxies, an omitted variable bias may play a role. However, we decide in favor of a stable co-integration relation.

[6]The sample correlation is about 95%.

[7]A negative coefficient could theoretically emerge if an increase in unemployment lowers inflation expectations and induces the central bank to reduce nominal interest rates. The same applies to safe-haven effects. Hence, we only interpret the unemployment rate as a credit risk premium, if its coefficient is positive. Given our interest rate decomposition, endogeneity issues may arise, but our results are roughly the same when using only lagged values for the explanatory variables including the short-run dynamics.

4 Results

As shown in Table 3, the Westerlund test rejects the null hypothesis of 'no cointegration' for all of the panel data models (Table 3, Model 1-11). Furthermore, the speed of adjustment, the coefficient of the co-integration relation COINT, is significantly negative as was to be expected. Model 1 is based on the sample 2009M1 to 2017M12 and includes both credit risk proxies, GS5Y and U, in addition to the structural interest components EONIA and TERM. All regressors show a significant long-run coefficient. As would be expected from our interest rate decomposition, the long-run coefficients of EONIA and TERM premium are close to 1. More precisely, in Model 1 a 100 basis point increase in EONIA or TERM causes an increase in the private-sector borrowing rate of 75 basis points and 80 basis points, respectively. The long-run coefficients of the two credit risk proxies are almost identical. The aggregate effect is located slightly above 20 basis points. This changes when we look at the results for the entire sample (Model 2). The coefficient for the unemployment rate is now negative and close to 0; the coefficient of the sovereign bond spread is now 18 basis points and the coefficients of EONIA and TERM rise to 84 and 96 basis points, respectively. Since the coefficient of the unemployment rate should no longer be interpreted as a credit risk premium, being negative, we omit it in Models 3 and 4 despite its previous significance, and re-estimate using both the reduced and the entire sample. The impact of the credit risk premium as measured by a 100 basis point increase of the sovereign bond spread is now 15 basis points irrespective of the sample size. Given our interest rate decomposition, we can draw a first conclusion: We find no evidence for an increasing spillover impact from sovereign bond spreads to private sector borrowing rates after 2009.[8] As we are interested in the maximum spillover effect, in the following, we focus on estimates for the entire sample.

We robustify the results of Models 3 and 4 by changing the estimation methodology (Models 5 and 6). Using dynamic fixed effects and the mean-group estimator, the coefficients of EONIA and TERM decline somewhat, but the GS5Y-coefficient remains in the range of 10 to 20 basis points. Although the Hausman test does not reject PMG as the preferred model over MG, the p-value of 10.6% is rather low. It is slightly higher once the unemployment rate is reintroduced as an additional credit risk proxy (Model 7). Nonetheless, we interpret the overall low p-value as an indication of cross-country heterogeneity. This reading is confirmed when comparing the results

[8]Country-specific results for the reduced sample 2009M1 to 2017M12 are available on request. While the number of observations decreases, the finding of our first conclusion does not change. If there is any time-variation, then it runs in the other direction. The coefficient of U increases slightly, while that of GS5Y decreases somewhat. The fact that the coefficient of GS5Y does not increase, does not imply that the corresponding credit risk, i.e. the coefficient times the data, has not increased since 2009.

of PMG Models 8 and 9 with MG models 10 and 11. Model 8 is based on a sample of European countries less affected by the euro crisis (Belgium, Finland, Germany, Netherlands) and the coefficient of sovereign bond spreads is again within the range of 10 to 20 basis points. Model 9, on the other hand, samples countries severely affected by the euro crisis (Ireland, Italy, Spain) and the coefficient of sovereign bond spreads is now 35 basis points and hence significantly above the range. Compared to the MG models 10 and 11, the p-value of the Hausman test is higher. Therefore, these panel data models are more homogeneous than those including all euro countries. The low R^2 and the low long-term coefficient of the term premium in Model 9 (and 11 respectively) indicate that the structural interest rate model is not optimally specified leading us to estimate country-specific models. Nevertheless, we can already draw our second conclusion: The average impact of sovereign bond spreads on private sector borrowing rates is between 10 and 20 basis points, while we suspect that the impact differs significantly across countries.[9]

Turning to the country-specific results in Table 4, note that the selected estimation design is based, first, on optimizing the co-integration test results and, second, on selecting lags according to Schwarz information criterion. This is why we first discuss the test results. The bounds test suggests for all countries that our variables are bounded together in the long run, as the test statistic exceeds the critical values as provided by Pesaran et al. (2001). Merely in the case of Ireland and an assumed I(1) bound, the null hypothesis ('No long-run relationship exists') can only be rejected at a 5 percent instead of a 1 percent significance level. By contrast, the Hansen instability test is more selective. Based on a 5 percent significance level, for Belgium, France, Italy and Netherlands 'a stable co-integration relation' can be rejected. The situation is the same when looking at the results of the Phillips-Ouliaris test. Again, based on a 5 percent significance level, it reports 'no stable co-integration relation' for Belgium, France, Italy and Netherlands.

Overall, the co-integration test results may be unsurprising in the case of France as, by construction, we can use only U as a credit risk proxy to avoid singularity in the estimated equation. For the other three countries experiencing problems with the stability of the co-integration relation, instability of the co-integration relation in the Belgian and Dutch cases is clearly supported by the p-values of both the Hansen instability and the Phillips-Ouliaris test results.[10] The Italian case is not as clear-cut. Indeed, if the sample size is reduced, the tests actually confirm a stable co-integration relation for the Italian specification, albeit not for the Belgian and Dutch specifications. Overall, we therefore accept the Italian model as well-specified.

[9] Strong heterogeneity among euro countries has also been found for other interest rate relations. For instance, see Bernhofer and van Treeck (2013) and Arnold and van Ewijk (2014) for the bank interest pass-through.

[10] In case of Phillips-Ouiliaris test, the null states that series are not co-integrated, in case of Hansen instability the null states that they are.

Discarding Belgium, France and the Netherlands due to the tested instability of the co-integration relation, we look at the long-run coefficients of the structural interest components in the remaining countries. For the long-run coefficient of EONIA, the values range from 50 to 120 basis points, the TERM premium from 30 to 120 basis points with the exception of the models for Ireland and Spain. Surprisingly, the Spanish coefficient for the term premium is insignificant and the Irish one negative. The long-run coefficients of U for Austria, Germany and Finland are also negative. In all other cases, where a change in the unemployment rate can be interpreted as a change in the credit risk premium, the values of the long-term coefficient are within a range of 10 to 25 basis points and hence quite similar to the average credit risk impact we obtained from the panel estimates.

Note that we only combine country-specific estimates in first stage of our analysis. This allows us to examine the significance of spill-over effects with a higher number of observations and obtain a rough estimate of the average effect. However, as our line of argument focuses on financial market valuation variability, these estimates are the starting point, not the end of our analysis.

Given our country-specific estimation design, the sovereign bond spread only plays a role for some countries, specifically Finland, Italy, Portugal and Spain. A 100 basis point increase in the government bond spread causes a 78 basis point increase in private sector financing conditions in Finland, a 73 basis point increase in Italy, but only an 8 basis point increase in Portugal and a 34 basis point increase in Spain. Even though the results should be read with caution in the Spanish case, where the long-run coefficient of the term premium is rather low, values of the adjusted R^2 being above 0.8 for each country estimation guarantee a minimum of statistical power. Our third conclusion therefore is that the maximum impact of a 100 basis bond increase in the sovereign bond spread on private sector borrowing rates is about 70 basis points. This is more than three times the average impact we previously found.

Our analysis reveals how much the coefficients differ across countries. The considerable range shows that there is ambiguous valuation of financial markets. Indirectly, this indicates that financial markets react too strongly in some cases implying that spill-over effects entail contagion costs. In summary, our empirical analysis points towards a sub-optimal policy design of the euro area.

Finally, we turn to the results of the non-linear ARDL estimates (Table 5). Overall, the results for the structural interest rate components (EONIA and TERM), for which we do not allow asymmetric effects, are similar to those previously discussed. With the exception of Finland, the long-run coefficients of both positive and negative changes in sovereign bond spreads are significant for all countries for which the spread has been identified as part of the co-integration relationship. The coefficients differ numerically in all cases with the one belonging to the positive spreads being larger for Finland, Italy and Spain. This means that a positive change of 100 basis points has a greater impact on private sector financing conditions than a negative one. For Portugal, surprisingly, hints of asymmetry go in the other direction.[11] In order to find out whether the differences in the long-run coefficients, including the short-term dynamics, produce significant effects, it is useful to consider the dynamic multiplier (Figure 2).

As can be seen from Figure 2, there is no evidence for asymmetries in countries with only the unemployment rate as a 'fundamental' credit risk proxy in the co-integration relation. For Portugal, the difference between positive and negative spread changes is significant, but the overall impact is small. For Spain, the impact of a changing positive spread level is at the upper end of the range we observed for the euro area average (+24 basis points as measured by the long-term coefficient). There is evidence of (temporary) asymmetry, but after 12 periods the difference between the effects of positive and negative spread changes is no longer significant. For Finland and Italy we observe the strongest effects. While in the case of Finland it is still questionable whether the difference between the impact of positive and negative spread changes is significant, the existence of an asymmetric influence in the case of Italy is clearly identified. Based on the non-linear estimates, the maximum effect of a 100 basis point increase in the government bond spread is about 50 basis points in the private sector borrowing rate.

At first glance, it is vexing that we do not find higher coefficients for Portugal and Spain. However, a more detailed analysis reveals other factors at work that are in line with our interpretation but cannot be accounted for within the present econometric framework. For example, Portugal experienced a decline in the volume of outstanding loans to nonfinancial corporations of 32 % from July 2011 to July 2016 which can be interpreted as an alternative reaction to the higher yield spread weakening the impact on borrowing rates. In addition, the duration of loans underlying the ECB's composite interest rate shifted markedly as the share of loans with a duration of over 5 years increased by almost 8 percentage points. This further mitigated the rise in the composite interest rate because interest rates on longer-term loans were on average 1.9 and 1.2 percentage points lower than those on short- and medium-term loans, respectively. Similar factors affect the econometric outcome for Spain.

Overall, our panel estimates show a significant impact of euro area sovereign bond spreads on private-sector lending. Furthermore, our linear country-specific estimates show substantial

[11] Note, however, that both long-run coefficients of the sovereign bond spread are close to 0 so that the difference between the coefficients, albeit significant, is too small to allow for insightful interpretation.

cross-country heterogeneity indicating the presence of financial valuation effects in excess of changes in economic fundamentals. This result is underscored by the lack of heterogeneity in the impact of unemployment, used as second risk proxy to capture fundamental credit risk. The results imply, firstly, that the effect of monetary easing was muted in the countries most affected by the crisis and secondly, that the higher financing costs and their depressing effect on private investment was more pronounced than may be justified by country-specific fundamental data. Our findings correspond to recent findings of analyses of interest-rate pass through in the euro area (Arnold and van Ewijk, 2014; Wolski, 2018). The spillover-effects of sovereign risk to private lending rates therefore exacerbated the economic downturn and contributed to the financial fragmentation within the euro area since the financial crisis.

5 Conclusion

In the wake of the international financial crisis, many euro area countries got caught up in a second crisis as the safe-asset quality of their sovereign bonds eroded. Enormous sovereign yield spreads mirrored the loss of confidence and set in motion a vicious cycle of rising financing costs, fiscal austerity, declining production and investment and strains in the national banking sectors. In our analysis, we show the impact of rising sovereign yield on private lending rates. These spillovers exacerbated the economic downturn which, in turn, fed the loss of confidence at the heart of rising sovereign borrowing costs. Our results show the detrimental effects caused by an erosion of safe-asset quality of sovereign bonds and highlight the importance of finding a solution to this fundamental deficiency in the architecture of the euro area.

Empirically, we find a small, but significant long-run impact of the sovereign bond spread on private lending rates using panel estimates. Moreover, there are large cross-country differences. Using country-specific estimates, the impact is found to be significant in only some countries, but at the same time the size of the maximum effect exceeds the average in linear specifications more than threefold and in nonlinear specifications more than twofold. For one country, we identify an asymmetrical effect so that positive spread changes have a larger impact on private sector borrowing costs than negative ones.

We interpret the existence of cross-country heterogeneity and asymmetry as an indication that the effect of government bond spreads on private credit risk is subject to financial market valuation effects. This interpretation is in line with the fact that heterogeneity of the spread coefficient across countries is not reflected in a corresponding heterogeneity of a more ‘fundamental’ credit risk proxy, i.e. the unemployment rate in our model. Future research should investigate this influence using microdata sets. Overall, the present analysis suggests that the cost of contagion effects in the euro area are substantial and will remain so until an effective form of European safe assets is created.

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Table 1: Country-specific statistics of Composite Costs of Borrowing (CCB), 2003-2017

Country Country Code	Austria OE	Belgium BE	Germany DE	Finland FI	France FR	Irleland IE	Italy IT	Netherlands NL	Portugal PT	Spain SP
Mean	3.33	3.54	3.65	3.49	3.60	4.30	4.08	4.13	5.05	3.67
Standard Deviation	0.81	0.95	1.21	1.07	1.02	0.67	1.01	0.85	1.50	1.04

Notes: All figures in %. Descriptive statistics underline the variety of the composite costs of borrowing within the euro area.

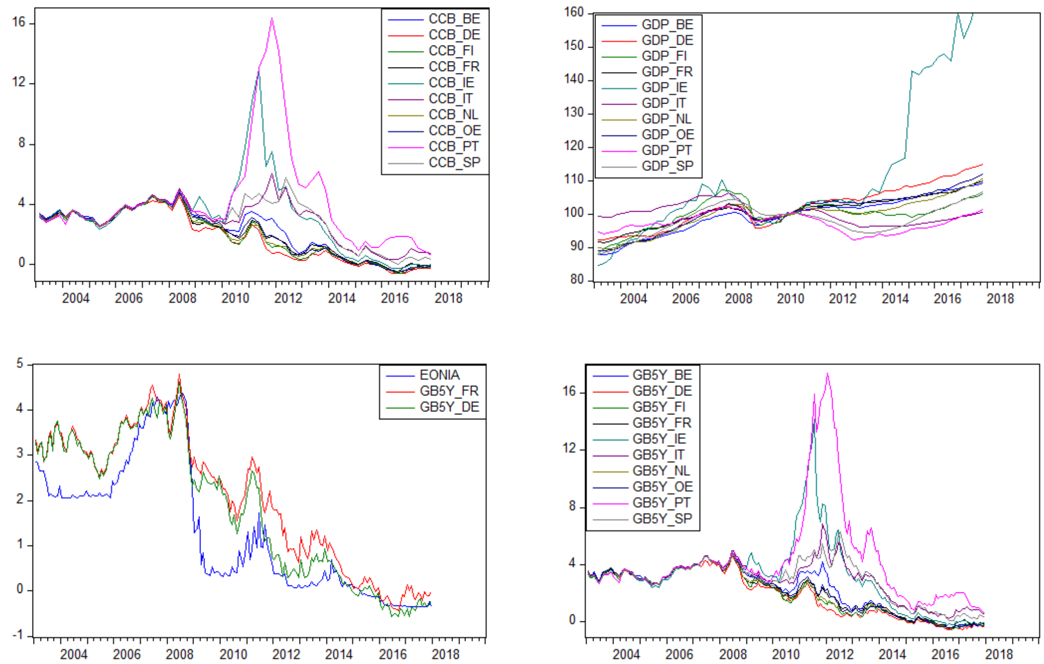


Figure 1: Composite costs of borrowing (top left), real GDP (top right), EONIA and 5 year French bond yield (bottom left) and 5 year bond yields (bottom right) for euro area countries.

Table 2: Summary of variables and expected signs in the long run relation

Acronym	Variable (in levels)	Source	Frequency	Expected Sign
CCB	long-term composite costs of borrowing	ECB	monthly	dependent
EONIA	euro area end of period closing	Euribor FBE	monthly	explanatory (+)
TERM	difference between French 5 year sovereign bond yield and EONIA	Macrobond Financial	monthly	explanatory (+)
GS5Y	difference between countries' and French 5 year sovereign bond yield	Macrobond Financial	monthly	explanatory (+)
U	unemployment rate, as percent of labor force, ILO-concept	National Statistics Offices	monthly	explanatory (+)

Notes: All data is loaded in %. If a stable co-integration relation is identified, the ARDL model allows to differentiate between long run (level) and short run relations (differences) by transferring the estimation results into error correction form.

Table 3: Panel ARDL estimates

Model Method Sample	(1) PMG 2009-2017	(2) PMG 2003-2017	(3) PMG 2009-2017	(4) PMG 2003-2017	(5) DFE 2003-2017	(6) MG 2003-2017	(7) MG 2003-2017	(8) PMG 2003-2017	(9) PMG 2003-2017	(10) MG 2003-2017	(11) MG 2003-2017
Regressor	Error Correction Model with Dependent Variable: $\Delta CCB_{i,t}$										
$COINT_{i,t-1}$	-0.235***	-0.142***	-0.162***	-0.154***	-0.155***	-0.244***	-0.313***	-0.139***	-0.203*	-0.244**	-0.308***
	(0.06)	(0.047)	(0.053)	(0.05)	(0.014)	(0.051)	(0.064)	(0.063)	(0.125)	(0.107)	(0.097)
$\Delta CCB_{i,t-1}$	-0.09	-0.106	-0.116	-0.087	-0.217***	-0.068	-0.053	0.061	-0.28***	0.08	-0.245***
	(0.086)	(0.087)	(0.091)	(0.085)	(0.024)	(0.079)	(0.077)	(0.161)	(0.084)	(0.144)	(0.094)
$\Delta CCB_{i,t-2}$									-0.184***		-0.173**
									(0.069)		(0.082)
$\Delta EONIA_{i,t}$	-0.164**	-0.06	-0.112	-0.048	-0.035	-0.112	-0.183**	0	0.035	-0.083*	-0.03
	(0.07)	(0.031)	(0.075)	(0.08)	(0.05)	(0.079)	(0.092)	(0.065)	(0.106)	(0.044)	(0.08)
$\Delta TERM_{i,t}$	-0.175*	-0.115	-0.156	-0.111	-0.086**	-0.13	-0.149	-0.013	-0.001	-0.072*	-0.002
	(0.093)	(0.087)	(0.102)	(0.086)	(0.041)	(0.095)	(0.106)	(0.051)	(0.029)	(0.039)	(0.039)
$\Delta GS5Y_{i,t}$	-0.002	0.009	0	0.01	0.036*	-0.018	-0.022	-0.002	0.04	-0.066	0.022
	(0.041)	(0.038)	(0.038)	(0.038)	(0.02)	(0.053)	(0.06)	(0.062)	(0.049)	(0.135)	(0.035)
$\Delta U_{i,t}$	-0.07	-0.118					-0.082*	-0.061***	-0.147	0.023	-0.171**
	(0.051)	(0.031)					(0.049)	(0.019)	(0.105)	(0.039)	(0.09)
constant	0.284***	0.329**	0.358**	0.335**	0.373***	0.583***	0.524**	0.282**	0.066**	0.688*	0.097
	(0.11)	(0.139)	(0.147)	(0.142)	(0.042)	(0.13)	(0.223)	(0.13)	(0.026)	(0.362)	(0.36)
Regressor	Long Run Coefficients ($COINT_{i,t-1} = CCB_{i,t-1} - \beta_1 EONIA_{i,t-1} - \beta_2 TERM_{i,t-1} - \beta_3 GS5Y_{i,t-1} - \beta_4 U_{i,t-1}$)										
$EONIA_{i,t-1}$	0.751***	0.843***	0.796***	0.805***	0.681***	0.694***	0.814***	0.87***	1.085***	0.722***	1.034***
	(0.038)	(0.019)	(0.042)	(0.012)	(0.038)	(0.057)	(0.097)	(0.021)	(0.073)	(0.063)	(0.257)
$TERM_{i,t-1}$	0.811***	0.946***	0.965***	0.88***	0.51***	0.483***	0.568***	0.977***	0.085	0.744***	0.175
	(0.032)	(0.032)	(0.029)	(0.028)	(0.086)	(0.174)	(0.192)	(0.036)	(0.085)	(0.131)	(0.467)
$GS5Y_{i,t-1}$	0.112***	0.18***	0.149***	0.146***	0.176***	0.103	0.068	0.143*	0.347***	-0.075	0.333*
	(0.031)	(0.035)	(0.033)	(0.029)	(0.036)	(0.195)	(0.19)	(0.089)	(0.074)	(0.429)	(0.175)
$U_{i,t-1}$	0.113***	-0.03***					0.025	-0.047***	0.11***	-0.064***	0.211**
	(0.015)	(0.013)					(0.069)	(0.015)	(0.02)	(0.022)	(0.086)
Test	Cointegration and Homogeneity Tests										
Westerlund	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Hausman						0.106	0.279			0.402	0.514
adjusted R^2	0.841	0.929	0.849	0.929	0.754	0.861	0.822	0.139	0.044	0.545	-0.230
Observations	972	1602	972	1602	1602	1602	1602	712	531	712	531

*Notes: The dependent variable is always the first difference of long-term composite cost of borrowing. Model (1) refers to PMG estimation including all ten euro area countries for a reduced sample; Model (2) refers to the same for the entire sample. Model (3) and (4) are the corresponding estimations including only GS5Y. Model (5) and (6) are dynamic fix effects and MG robustifications including only GS5Y. Model (6) is the MG estimation including both GS5Y and U. Model (8) and (9) are PMG estimation using either a sample of northern European countries (Belgium, Finland, Germany, Netherlands) or peripheral ones (Ireland, Italy, Spain). Model (10) and (11) refer to the corresponding MG robustifications. Robust standard errors are reported in parantheses. *, **, and *** denotes significance at 10%, 5%, and 1% levels, respectively. The long-run coefficient for most of the models is in the range of 10 to 20 basis points irrespectively of the sample size. For statistical tests, p-values are reported. The Westerlund test(s) rejects the null hypothesis of 'no co-integration' for all models. The Hausman tests cannot reject that PMG (Model 4 and 2) relative to MG (Model 6 and 7) is effecient and hence preferred, but p-values are rather small. The p-values increase when comparing estimation based on country-splits (Model 8 and 9 relative to Model 10 and 11).*

Table 4: Country-specific ARDL estimates, 2003-2017

Country	OE	BE	DE	FI	FR	IE	IT	NL	PT	SP
Regressor	Error Correction Model with Dependent Variable: ΔCCB_t									
$COINT$	-0,298***	-0,044***	-0,296***	-0,563***	-0,054***	-0,186***	-0,183***	-0,077***	-0,601***	-0,529***
	(0,042)	(0,008)	(0,021)	(0,061)	(0,01)	(0,046)	(0,027)	(0,012)	(0,065)	(0,066)
ΔCCB_{t-1}	-0,224***		-0,159***		0,08	-0,344***	-0,4***	0,18***		
	(0,065)		(0,058)		(0,069)	(0,07)	(0,066)	(0,068)		
ΔCCB_{t-2}					0,181***	-0,238***	-0,288***			
					(0,066)	(0,067)	(0,066)			
ΔCCB_{t-3}					0,262***					
					(0,066)					
$\Delta EONIA$	-0,124	0,072**	0,084***	0,195*	0,001	0,217**	0,27***	0,037	-0,281	0,369**
	(0,08)	(0,029)	(0,029)	(0,1)	(0,028)	(0,1)	(0,1)	(0,029)	(0,302)	(0,144)
$\Delta EONIA_{t-1}$		0,272***				0,28***				
		(0,029)				(0,087)				
$\Delta EONIA_{t-2}$		0,181***				0,479***				
		(0,021)				(0,077)				
$\Delta EONIA_{t-3}$		0,083***								
		(0,023)								
$\Delta TERM$	0,052	0,002	0,132***	0,172**	0,017	-0,145*	0,035	0,051**	-0,25	0,013
	(0,064)	(0,023)	(0,023)	(0,084)	(0,022)	(0,08)	(0,082)	(0,024)	(0,248)	(0,11)
$\Delta TERM_{t-1}$		0,11***								
		(0,023)								
$\Delta GS5Y$		0,034		-0,037			0,189***	0,049	0,083	0,124*
		(0,037)		(0,149)			(0,051)	(0,07)	(0,053)	(0,069)
$\Delta GS5Y_{t-1}$		-0,113***					-0,165***			-0,071
		(0,038)					(0,052)			(0,068)
$\Delta GS5Y_{t-2}$		-0,187***					-0,142***			-0,279***
		(0,036)					(0,053)			(0,068)
$\Delta GS5Y_{t-3}$		-0,096**								0,087
		(0,038)								(0,072)
$\Delta GS5Y_{t-4}$										-0,182***
										(0,067)
ΔU	-0,15*		-0,029	0,099	-0,11**	-0,087	-0,001		-0,139	0,047
	(0,088)		(0,053)	(0,149)	(0,055)	(0,082)	(0,074)		(0,248)	(0,091)
Regressor	Long Run Coefficients from $COINT_{t-1} = CCB_{t-1} - \beta_1 EONIA_{t-1} - \beta_2 TERM_{t-1} - \sum_{k=3} \beta_k x^k_{t-1} - constant$									
$EONIA$	0,464***	0,799***	0,888***	0,642***	0,834***	0,541***	1,239***	0,663***	0,971***	1,035***
	(0,038)	(0,095)	(0,021)	(0,036)	(0,169)	(0,101)	(0,17)	(0,063)	(0,075)	(0,062)
$TERM$	0,298***	0,336*	0,983***	0,538***	0,938***	-0,53***	0,702***	0,855***	1,208***	0,035
	(0,073)	(0,2)	(0,023)	(0,053)	(0,187)	(0,179)	(0,213)	(0,112)	(0,143)	(0,073)
$GS5Y$		1,453***		0,783***			0,732***	-0,949*	0,079**	0,335***
		(0,542)		(0,243)			(0,118)	(0,594)	(0,031)	(0,076)
U	-0,37***		-0,049***	-0,129**	0,163	0,171***	0,237**		0,177***	0,104***
	(0,101)		(0,013)	(0,06)	(0,264)	(0,039)	(0,113)		(0,046)	(0,02)
$constant$	4,361***	1,908***	2,024***	3,383***	0,152	2,42***	-0,988	2,407***	0,701	0,253
	(0,57)	(0,289)	(0,061)	(0,538)	(2,687)	(0,471)	(1,347)	(0,185)	(0,563)	(0,382)
Test	Cointegration Tests									
Bounds	11,18	5,96	40,95	15,07	5,63	3,98	7,30	9,28	14,08	8,80
Hansen	0,0933	0,0361	> 0,2	> 0,2	< 0,01	0,1606	0,0198	< 0,01	> 0,2	> 0,2
Phillips	0,0101	0,6857	0,0012	0,0000	0,6755	0,0000	0,2492	0,4631	0,0000	0,0000
adj. R^2	0,9529	0,9959	0,9972	0,9577	0,9965	0,8991	0,9538	0,9942	0,8001	0,9238
Observ.	178	176	178	179	176	177	177	178	179	175
#Models	2058	2058	2058	14406	2058	2058	14406	2058	14406	14406

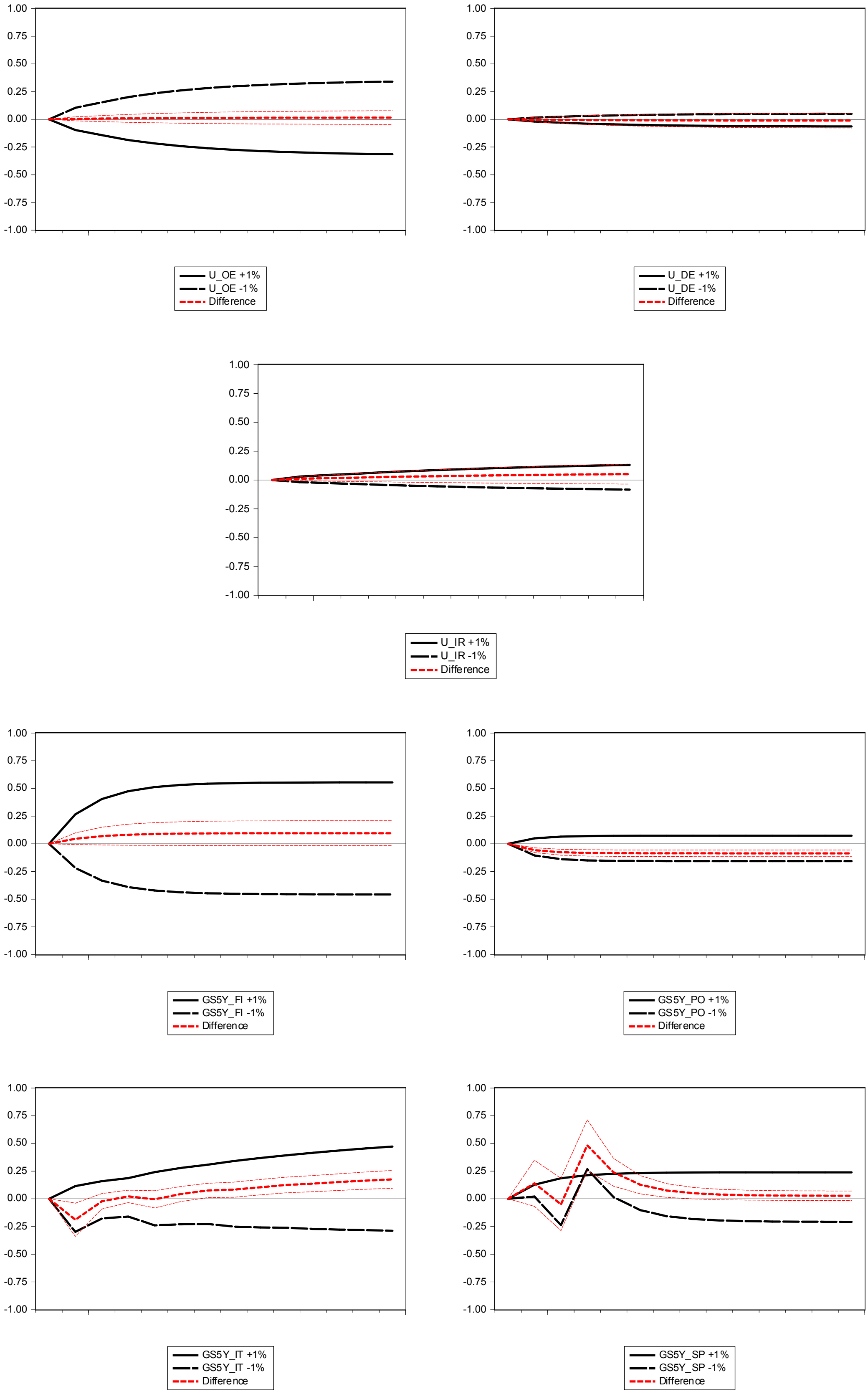
*Notes: The dependent variable is always the first difference of the long-term composite cost of borrowing. Robust standard errors are reported in parantheses. *, **, and *** denotes significance at 10%, 5%, and 1% levels, respectively. In case of the Bounds test, we report the test statistic, whereat values larger 4 indicate a cointegration relation. In case of the other co-integration tests, p-values are reported. In case of Phillips-Ouilliaris test, the null states that series are not co-integrated, in case of Hansen instability the null states that they are. $COINT$ denotes the cointegration relation, the corresponding coefficient is the speed of adjustment and #Models stands for the number of models evaluated using Schwarz information criterium with six as the maximum number of lags. Overall, co-integration test results indicate no stable co-integration relation for Belgium, France and Netherlands, the inclusion of both $GS5Y$ and U for Finland, Italy, Portugal and Spain as well as the exclusion of GS5Y for Austria, Germany and Ireland. Among the well specified models, the largest long run coefficients of GS5Y are above 0.7 (Finland, Italy).*

Table 5: Country-specific NARDL estimates, 2003-2017

Country	OE	DE	FI	IE	IT	PT	SP
Regressor	Error Correction Model with Dependent Variable: ΔCCB_t						
$COINT$	-0,293***	-0,292***	-0,573***	-0,207***	-0,186***	-0,695***	-0,535***
	(0,041)	(0,021)	(0,062)	(0,047)	(0,024)	(0,069)	(0,064)
ΔCCB_{t-1}	-0,228***	-0,159***		-0,342***	-0,43***		
	(0,065)	(0,058)		(0,07)	(0,061)		
ΔCCB_{t-2}				-0,242***	-0,337***		
				(0,066)	(0,062)		
$\Delta EONIA$	-0,117	0,082	0,231**	0,209**	0,313***	-0,384	0,325**
	(0,081)	(0,029)	(0,1)	(0,1)	(0,096)	(0,296)	(0,147)
$\Delta EONIA_{t-1}$				0,244***			
				(0,089)			
$\Delta EONIA_{t-2}$				0,451***			
				(0,078)			
$\Delta TERM$	0,059	0,13***	0,218***	-0,126	0,161**	-0,565**	0,062
	(0,064)	(0,023)	(0,084)	(0,079)	(0,078)	(0,242)	(0,111)
$\Delta GS5Y^{+}$			-0,038		0,09	0,052	0,202**
			(0,181)		(0,068)	(0,074)	(0,098)
$\Delta GS5Y^{-}$			-0,225		0,303***	0,074	-0,023
			(0,252)		(0,075)	(0,075)	(0,109)
$\Delta GS5Y^{-}_{t-1}$							0,141
							(0,112)
$\Delta GS5Y^{-}_{t-2}$							-0,485***
							(0,109)
ΔU^{+}	-0,14	-0,119		-0,17			
	(0,118)	(0,1)		(0,107)			
ΔU^{-}	-0,147	-0,003		-0,029			
	(0,131)	(0)		(0,119)			
ΔU			0,135		0,078	-0,535**	0,011
			(0,209)		(0,074)	(0,25)	(0,093)
ΔU_{t-1}					0,273***		
					(0,074)		
Regressor	Long Run Coefficients from $COINT_{t-1} = CCB_{t-1} - \beta_1 EONIA_{t-1} - \beta_2 TERM_{t-1} - \sum_{k=3} \beta_k x^k_{t-1} - constant$						
$EONIA$	0,502***	0,887***	0,804***	0,702***	1,83***	0,319**	1,108***
	(0,1)	(0,021)	(0,11)	(0,202)	(0,239)	(0,155)	(0,089)
$TERM$	0,337***	0,985***	0,675***	-0,302	1,856***	0,075	0,122
	(0,118)	(0,034)	(0,108)	(0,298)	(0,362)	(0,258)	(0,145)
$GS5Y^{+}$			0,557**		0,624***	0,071**	0,236***
			(0,236)		(0,119)	(0,028)	(0,063)
$GS5Y^{-}$			0,463		0,348***	0,156***	0,212***
			(0,245)		(0,125)	(0,03)	(0,06)
U^{+}	-0,337***	-0,065		0,18***			
	(0,13)	(0,053)		(0,041)			
U^{-}	-0,354***	-0,052***		0,113*			
	(0,11)	(0,016)		(0,064)			
U			-0,035		0,003	0,272***	0,106***
			(0,083)		(0,13)	(0,045)	(0,019)
$constant$	2,504***	1,58***	1,944*	2,511***	-1,943	2,592***	-0,119
	(0,473)	(0,1)	(1,001)	(0,873)	(1,265)	(0,631)	(0,405)

*Notes: The dependent variable is always the first difference of the long-term composite cost of borrowing from the ECB. Robust standard errors are reported in parantheses. *, **, and *** denotes significance at 10%, 5%, and 1% levels, respectively. $COINT$ denotes the cointegration relation and the corresponding coefficient is the speed of adjustment. Model specifications are the same as in the linear case with the exception that for the proxy variables of credit risk a separation is now made between positive and negative values. This increases the number of variables. At the same time, the long run coefficients of $GS5Y^{+}$ decrease slightly relative to the linear case (e.g. Finland, Italy). Asymmetries, as in the case of Italy, Portugal and Spain, are reflected by the significance of both variables with positive ($GS5Y^{+}$) and negative values ($GS5Y^{-}$).*

Figure 2: Dynamic Multiplier of country-specific NARDL estimates, 2003-2017



Notes: The dynamic multiplier captures the overall impact from both short run and long run coefficients. For the three countries (Austria [OE], Germany [DE], Ireland [IR]) at the top, we examine asymmetries of the unemployment rate, as the co-integration relation does not include the sovereign bond spread. For the four countries (Finland [FI], Portugal [PO], Italy [IT], Spain [SP]) at the bottom, we explore asymmetries of the sovereign bond spread. In most cases, we do not find asymmetry. In two cases (Finland, Spain), it is debatable. Here, asymmetric effects are rather small, or temporary. In one case (Italy), there is strong evidence for asymmetry.