

“Determinants of credit default swaps implied ratings during the crisis: was sovereign risk mispriced?”

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ARTICLE INFO

Maria Alberta Oliveira and Carlos Santos (2018). Determinants of credit default swaps implied ratings during the crisis: was sovereign risk mispriced?. *Investment Management and Financial Innovations*, 15(3), 1-14. doi:[10.21511/imfi.15\(3\).2018.01](https://doi.org/10.21511/imfi.15(3).2018.01)

DOI

[http://dx.doi.org/10.21511/imfi.15\(3\).2018.01](http://dx.doi.org/10.21511/imfi.15(3).2018.01)

RELEASED ON

Wednesday, 04 July 2018

RECEIVED ON

Monday, 09 April 2018

ACCEPTED ON

Monday, 25 June 2018

LICENSE



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JOURNAL

"Investment Management and Financial Innovations"

ISSN PRINT

1810-4967

ISSN ONLINE

1812-9358

PUBLISHER

LLC "Consulting Publishing Company "Business Perspectives"

FOUNDER

LLC "Consulting Publishing Company "Business Perspectives"



NUMBER OF REFERENCES

35



NUMBER OF FIGURES

0



NUMBER OF TABLES

3

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BUSINESS PERSPECTIVES



LLC "CPC "Business Perspectives"
Hryhorii Skovoroda lane, 10, Sumy,
40022, Ukraine

www.businessperspectives.org

Received on: 9th of April, 2018

Accepted on: 25th of June, 2018

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DETERMINANTS OF CREDIT DEFAULT SWAPS IMPLIED RATINGS DURING THE CRISIS: WAS SOVEREIGN RISK MISPRICED?

Abstract

This paper addresses the question of whether sovereign risk pricing was related to macroeconomic fundamentals, between 2007 and 2015, in a sample of OECD countries. The authors argue that the conflicting evidence in the literature is due to poor methodology options. The researchers innovate by modelling sovereign credit default swaps implied ratings as our sovereign risk proxy, instead of spreads, avoiding common pitfalls. Furthermore, the authors improve the variable selection, model specification and the econometric procedures used. A panel ordered probit model is chosen, assuring robust inference. The authors relax the parallel lines assumption, allowing for rating-varying coefficients of explanatory variables. The result is the first congruent model of sovereign risk during the years of the financial crisis and of the Euro Area crisis. Fiscal space variables, economic activity indicators, variables pertaining to external imbalances, and contagion proxies are relevant, with effects matching theory priors. The scientists clarify conundrums in the previous literature, posed by lack of significance of some macro fundamentals and by puzzling signs of some estimated coefficients. Moreover, this is the first paper to estimate not only the global risk premium, but also the impact of changing risk aversion. The authors find no support for claims of sovereign risk mispricing during the sample period. The results allow relevant policy conclusions, namely concerning the validity of different fiscal consolidation paths in financially distressed countries.

Keywords

credit default swaps, sovereign risk, ratings, macro-financial fundamentals, panel ordered probits

JEL Classification

C23, E44, F34, G01, G12, G15

INTRODUCTION

Understanding the fluctuations of credit risk is of the utmost relevance both for policy-makers and for investment managers, particularly since the onset of the financial crisis in 2007. The Euro Area (EA) debt crisis has further increased this relevance. If earlier academic research on credit risk mispricing was focused on US corporate bonds, and on sovereign credit default swaps (SCDSs) for emerging economies, Heynderickx et al. (2016) argue the EA is now at the core of the credit spreads puzzle. Feldhüter and Schaefer (2018) restrict the dimension of the puzzle to sovereign securities. As pointed out, inter alia, by Bannier et al. (2014), the surge of SCDSs in recent years is mainly due to the European debt crisis. In this journal, Oliveira and Santos (2015) have extensively discussed these OTC derivatives, while providing evidence of the vast related research propelled by the EA crisis.

The relevance of this topic for policy-makers is easy to understand. During the EA crisis, fluctuations in sovereign risk premia account for 30 to 50% of the forecast errors in unemployment, and for 20 to 40% of the increase in private borrowing costs (Bahaj, 2014). The transmission mechanism from SCDSs spreads to the real economy

rests on their leading role on the price discovery process in bond markets (Delatte et al., 2012), and on the upper bounds sovereign ratings still pose on corporate ratings (Borensztein et al., 2013)¹. Policy-makers are aware of the role of country ratings in financial stability (Klinger & Lando, 2018), since the linkage between sovereign and banking risk is well documented (Bruneau et al., 2014). Furthermore, access to the ECB Asset Purchasing Program is contingent on the existence of at least one major rating agency grading such debt securities above a minimum threshold².

From the perspective of an investments manager, the ability to assess credit risk is also of paramount importance, since portfolios usually contain corporate and / or sovereign bonds. A correct hedging strategy entails assessing the likelihood of default, namely to decide the amount of protection to buy in the form of Credit Default Swaps (CDSs). Moreover, if the investment manager wishes to buy CDSs for trading, the ability to anticipate changes in spreads allows significant profit opportunities³.

Irrespective of whether the motivation is policy design or building trading and hedging strategies, understanding the relationship between sovereign risk and economic fundamentals has become a part of the research agenda in finance. Notwithstanding, the literature is far from reaching sound conclusions. As we shall debate in section 1, this is largely due to poor methodology options. Afonso et al. (2007) had argued that modelling sovereign spreads is difficult, advising the usage of ratings as latent dependent variables in ordered probit models. A panel approach was also strongly recommended to increase the robustness of statistical inference. However, the empirical literature has neglected proxies for sovereign risk other than spreads and has paid no attention to the authors' recommendation on econometric methods.

The research question motivating this paper is whether there is a role for economic fundamentals in sovereign risk pricing. We improve on the literature in different ways. Firstly, this is the first paper to use SCDs implied ratings, constructed using spread implied credit default probabilities (CDPs)⁴ as a proxy of sovereign risk. Secondly, we use a panel ordered probit model, as recommended by Afonso et al. (2007). Thirdly, we explore the implications of relaxing the parallel lines assumption, which had never been discussed in this literature. As such, instead of simply including time-varying coefficients, we also allow rating-varying coefficients. We do not assume that different crisis stages have similar cross-sectional impacts on sovereign risk, irrespective of each country's rating at the time. Rather, the no parallel lines hypothesis allows to assess the impact of different crisis periods per rating class. We also innovate by including in our model a proxy for time-varying risk aversion. Our estimation results improve on the literature: the first congruent model relating macro-financial variables to sovereign risk is achieved. We have tested the inclusion of proxies for all the relevant dimensions suggested in earlier papers. They proved to be significant and their estimated impact on sovereign risk matches theory priors. Thus, previous empirical conundrums were solved. Finally, this paper also innovates by exploring a novel data set on SCDs.

The paper is organized as follows. Section 1 provides a critical overview of the relevant literature. Section 2 explains our methodology. Section 3 discusses the research hypotheses. Section 4 describes the data. Section 5 discusses estimation results. Final section concludes the paper.

1 Although rating agencies claim to have moved away from the "sovereign ceilings" policy, evidence shows that sovereign ratings are still a major determinant of corporate ratings (Borensztein et al., 2013).

2 This is of special concern in the financially distressed EA periphery (De Santis, 2016). In the Portuguese case, until early 2017, financial stability was dependent on the Canadian rating agency DBRS, the only one classifying the country's sovereign debt just above "speculative grade" (as BBB).

3 Oliveira and Santos (2015) discuss profitable trading strategies for investors holding SCDs.

4 Research on CDPs and implied ratings has been conducted in Finance, but not with respect to the macroeconomic determinants of sovereign risk. For details on filtering CDPs from CDSs spreads see, inter alia, Elkamhi et al. (2014).

1. LITERATURE REVIEW

The conclusions emerging from the empirical literature on sovereign risk and fundamentals are largely contradictory. Indeed, not only do candidate explanatory variables vary across papers, but also even papers including similar covariates reach opposite conclusions regarding their significance and impact. Furthermore, poor choices of the proxy for sovereign risk are common. Finally, the econometric methods chosen are misleading. We will divide our critical overview of the literature in 3 subsections, each pertaining to one of the referred problems.

1.1. Candidate covariates

It is possible to arrange the variables assessed in the literature into five main groups: fiscal space; external imbalances; risk and contagion; economic activity; less common covariates.

1.1.1. Fiscal space

The literature has used several fiscal space indicators. Indeed, from the most relevant empirical studies, only Boffelli et al. (2017) and Arghyrou and Kontonikas (2012) neglect this dimension.

The ratio of public debt to GDP is used in Caceres et al. (2010), Santos (2011), Beirne and Fratzscher (2013), and Afonso et al. (2015). Paniagua et al. (2017) use the difference between this ratio, for each country, and its value for Germany. The government balance to GDP ratio is used in Caceres et al. (2010), Beirne and Fratzscher (2013), Afonso et al. (2015), Yuan and Pongsiri (2015), and Kriz et al. (2015). Yuan and Pongsiri (2015) also include a measure of the expected government balance to GDP ratio. Aizenman et al. (2013) use the ratios of public debt and of the government budget balance to total taxes. The ratio of total government revenue to GDP is included in Oliveira et al. (2012) and in Kriz et al. (2015). Oliveira et al. (2012) also include subclasses of government expenditure.

Fiscal space variables are significant in Caceres et al. (2010), Aizenman et al. (2013), Kriz et al. (2015) and Paniagua et al. (2017). Santos (2011) has concluded that the public debt to GDP ratio was only significant for the top 10% of the CDPs distri-

bution. Oliveira et al. (2012) find their classes of government expenditure to be relevant, but only before the EA crisis. Differently, the taxes to GDP ratio is relevant throughout their entire sample period. Afonso et al. (2015) conclude against the significance of the public debt to GDP ratio. This is unexpected, since their sample contains only the EA countries. Notwithstanding, the authors find the government balance to be relevant. Beirne and Fratzscher (2013) achieve even more puzzling results: the public debt to GDP ratio is always irrelevant for EA countries, and the government balance ratio matters only before the crisis period. Yuan and Pongsiri (2015) conclude that the government balance is irrelevant, although the expected government balance is significant. The authors acknowledge that multicollinearity is affecting their results.

1.1.2. External imbalances

Cantor and Packer's (1996) seminal work provides the expectation of a positive correlation between a country's current account balance and its sovereign spreads. However, Beirne and Fratzscher (2013) find that the current account to GDP ratio is not significant for the EA countries during their sample period, although it is relevant for other advanced economies. Furthermore, Aizenman et al. (2013) conclude that the external debt to GDP ratio is not significant. Albeit concluding against the relevance of the current account, Yuan and Pongsiri (2015) find that the ratio of external debt to GDP increases sovereign spreads. Oliveira et al. (2012), Santos (2011) and Kriz et al. (2015) also find the external debt ratio to be significant.

When taking the space of external imbalances to be represented by other variables, results are also ambiguous. Aizenman et al. (2013) conclude against the relevance of trade openness. Arghyrou and Kontonikas (2012) find that country's competitiveness (proxied by the logarithm of the real weighted exchange rate) is relevant in explaining sovereign risk. Neither Afonso et al. (2015) nor Yuan and Pongsiri (2015) find the exchange rates to be relevant.

Boffelli et al. (2017), Caceres et al. (2010) and Paniagua et al. (2017) do not assess the role of external imbalances.

1.1.3. Risk and contagion

The VIX index⁵ is the standard proxy for global risk in this literature. It is significant in explaining country-specific sovereign risk in Paniagua et al. (2017) and Afonso et al. (2015). However, Arghyrou and Kontonikas (2012) do not find a role for global risk in their first sample period (2001–2007). The authors claim that relevance of global risk occurs during the financial crisis alone. Surprisingly, during the EA crisis, they don't find the VIX to be significant.

Beirne and Fratzscher (2013) use the first difference of the VIX as a proxy of global risk⁶. They reach the unexpected conclusion that an increase in global risk decreases the EA sovereign spreads. For emerging economies, they find the expected positive impact. Kriz et al. (2015) use year-specific dummies as proxies for global risk. These are significant, and the estimated coefficients are positive, as expected. Differently, Caceres et al. (2010) use the Index of Global Risk Aversion (IGRA). They find the IGRA to be significant and positively correlated with sovereign risk, with the surprising exception of the EA periphery.

The possible relevance of global risk is not accounted for in Santos (2011), Oliveira et al. (2012), Yuan and Pongsiri (2015) and Boffelli et al. (2017).

Beirne and Fratzscher (2013) use regional dummies to test for shift-contagion, concluding in favor of that hypothesis. Caceres et al. (2010) had also found evidence of contagion within the EA periphery.

1.1.4. Economic activity

The real growth rate of GDP is relevant in the models of Kriz et al. (2015) and Yuan and Pongsiri (2015). Beirne and Fratzscher (2013) find that, for the EA periphery, real GDP growth does not explain sovereign risk. Adding to the authors' surprise, they find that, for other advanced economies, increases in real GDP growth are estimated to significantly increase sovereign spreads. Arghyrou and Kontonikas (2012) use output growth differ-

entials between countries as a proxy for the space of growth related variables, but only in single equation models (not in their panel setting). They conclude that the variable is significant in only one of the sample countries.

The growth rate of industrial production has been used by Oliveira et al. (2012). A negative relationship between industrial production and sovereign spreads is found, although only for the period before the EA crisis. The differential between each country's industrial production growth and the German one is significant in Boffelli et al. (2017), but not retained in the final congruent model of Afonso et al. (2015). Other papers referred to in this section neglect the possible relevance of real growth.

Paniagua et al. (2017) and Boffelli et al. (2017) have concluded that the unemployment rate matters for sovereign risk. However, Kriz et al. (2015) did not find it to be significant. No other study has tested the inclusion of this variable.

The role of inflation has been controversial, since the seminal work of Cantor and Packer (1996). Aizenman et al. (2013), Oliveira et al. (2012) and Yuan and Pongsiri (2015) found inflation to be irrelevant. Differently, Kriz et al. (2015) found the variable to be significant. Notwithstanding, the authors do not clarify whether they are referring to inflation or to changes in inflation. None of the other papers has discussed the relevance of inflation in sovereign risk models.

1.1.5. Other variables

Yuan and Pongsiri (2015) test the relevance of expectations concerning GDP growth, but fail to reach a conclusion. Differently, Boffelli et al. (2017) conclude that expectations, measured by the difference between each country's Eurostat index of business confidence and the corresponding value for Germany, are relevant in explaining sovereign risk. However, this is likely to be capturing the effect of correlated omitted variables, given the authors' extremely parsimonious representation of macroeconomic fundamentals. Liquidity of the

5 Obtained using call and put implied volatilities from the S&P 500 index.

6 In section 4, we shall provide a different interpretation for this difference.

sovereign bonds market is assessed in Afonso et al. (2015) and Paniagua et al. (2017). Both conclude against its relevance, except for some countries in the EA periphery. Notwithstanding, Arghyrou and Kontonikas (2012) found bond markets liquidity to be relevant, but with an unexpected positive estimated coefficient, implying lower spreads for less liquid markets.

1.2. Explained variable

The most common sovereign risk proxy is the spread between the interest rate on a country's government bonds and that on bonds of some other reference country for the same maturity. In the recent literature, the reference is usually the 10 years maturity German bonds. This is the case in Paniagua et al. (2017), Arghyrou and Kontonikas (2012), Oliveira et al. (2012), Afonso et al. (2015) and Boffelli et al. (2017). Caceres et al. (2010) use the spread of a 10 years maturity EuroSwap. Beirne and Fratzscher (2013) use the spread between government bonds and 3 months money market rates.

Notwithstanding, it is widely acknowledged that SCDSs spreads provide a better proxy of sovereign risk (e.g. Delatte et al., 2012). Paniagua et al. (2017) explain they would rather have used SCDSs spreads, but failed to obtain data prior to 2007. Albeit using bond spreads, Beirne and Fratzscher (2013) conduct robustness tests, checking their conclusions with SCDSs spreads. Aizenman et al. (2013) and Yuan and Pongsiri (2015) use SCDSs spreads as their dependent variable.

Different sovereign risk proxies, derived from SCDSs, are used in Kriz et al. (2015) and Santos (2011). The former used a panel of CDPs, for a sample of 57 countries, referring to the period between 2009 and 2013. The latter uses CDPs' implied ratings in a cross-sectional sample.

1.3. Econometric methods

Dynamic panel models, occasionally including time-varying parameters, are the most commonly used when the dependent variable is a spread. This is the case in Aizenman et al. (2013), Beirne and Fratzscher (2013), Arghyrou and Kontonikas (2012), Oliveira et al. (2012), and Yuan and Pongsiri (2015). Paniagua et al. (2017) use a dy-

namic panel setting allowing for time-varying parameters, and for random coefficients (with cyclical asymmetries between countries as the source of randomness).

Afonso et al. (2015) use a cross-sectional model instead of a panel approach to assess the results of the general-to-specific model selection embodied in autometrics (Doornik, 2009). Differently, Caceres et al. (2010) modelled the spread with a GARCH. Boffelli et al. (2017) have extended the Mixed Data Sampling (MIDAS) approach to combine high frequency financial information with lower frequency variables. However, the restriction to monthly macro variables leads to a poor picture of the role of fundamentals in their analysis.

The conflicting results referred to in subsection 1.1 are likely an outcome of the difficulty in modelling sovereign spreads, which had been outlined by Afonso et al. (2007). Statistical inference with such a dependent variable is often invalid. The possibility of explosive roots in spreads is one of the reasons for this (e.g. Oliveira & Santos, 2015). Surrogate proxies for sovereign risk should be favored.

Kriz et al. (2015) have used a panel specification to relate sovereign CDPs to fundamentals. Given the flawed results of their preliminary linear probability models, the authors chose a log-normal distribution for the random error. Notwithstanding, they should have assessed standard panel probits and logits. Indeed, even their log-normal model has led to dubious inference conclusions, given the censored and clustered nature of CDPs data. Flannery and Hankins (2013) had extensively discussed the econometrics of panel data when the dependent variable is censored and clustered, suggesting the best estimation procedures for such cases. Kriz et al. (2015) neglected those recommendations.

Santos (2011) built a cross-sectional ordered probit model of implied SCDSs ratings. Afonso et al. (2007) had clarified that ordinal models should be used in panel settings for robust inference on ratings, given their asymptotic properties. Hence, insufficient data for a panel ordered probit sheds doubts over the conclusions in Santos (2011).

2. ANALYSIS METHODS

Credit ratings provide a natural field of application for ordered probit models (e.g. Cantor & Packer, 1996). The nature of ratings data fits in a framework where the difference between the ordinal scores of, say, 4 and 3, is not equivalent to the difference between the cardinals 4 and 3. For the latter, multinomial probits and logits are better suited. Differently, albeit being associated with differences in credit worthiness, variations in rating scores are not quantitatively equal to those.

In Corporate Finance, selection bias is common, when modelling ratings with ordered probits. This is attributed to privileged information managers possess about their firms. When choosing to solicit ratings, managers anticipate good scores. Hence, firms receiving unsolicited ratings are likely to be subject to downward biased rating estimators (e.g. Poon, 2003).

The methodology framework in which this paper studies sovereign risk avoids the selection bias problem, as data provided by rating agencies is avoided. Instead, ratings are derived from market information (SCDSs spreads and their implied CDPs). This is the case for every country in the sample in every period. Self-selection is prevented, since the decision to trade SCDSs on a country's debt depends on market participants alone, without interference from the country's government. All SCDSs implied ratings may be viewed as "unsolicited".

The structure of an ordered probit contemplates S ordered classes. Let $R_{i,t}^*$ be the latent or unobserved credit risk of country i , at time t . Let $R_{i,t}$ be the rating computed from the country's SCDSs spread at time t , for a certain maturity. It is generally assumed that:

$$R_{i,t}^* = x_{i,t}\beta_* + \mu + \varepsilon_{i,t}, \quad (1)$$

$R_{i,t}^*$ and the observed market implied rating, $R_{i,t}$ are related by:

$$R_{i,t} = \begin{cases} 1 < R_{i,t}^* \leq \alpha_0 \\ 2 \leq \alpha_0 < R_{i,t}^* \leq \alpha_1 \\ 3 \leq \alpha_1 < R_{i,t}^* \leq \alpha_2 \\ 4 \leq \alpha_2 < R_{i,t}^* \leq \alpha_3 \\ \dots \\ S \leq R_{i,t}^* < \alpha_{S-2} \end{cases}, \quad (2)$$

$x_{i,t}$ is a vector of explanatory variables, and β_* the vector of coefficients. As expected for a panel data setting, equation (1) contemplates two random errors: μ_i and $\varepsilon_{i,t}$. At the estimation level, we shall cope with this through the assumption that they are both normally distributed. Hence, we shall work with the Random Effects ordered probit model, using Maximum Likelihood (ML) estimation (see, inter alia, Frechette (2001) for implementation in STATA). Furthermore, we allow the coefficients to vary according to a country's risk class. This option is known as relaxing the parallel lines assumption in ordered probit theory (Williams, 2006, 2016; Pfarr et al., 2011). The intuition for the possibility of rating-varying coefficients results from an observation in Afonso et al. (2007), emerging from subsample analysis, where some variables had a different impact for countries with different ratings. Neither Afonso et al. (2007) nor subsequent authors have pursued this research. Thus, we innovate by checking thoroughly if the effect of each variable has a random component, depending on the rating order⁷. α_i are the cut-off unknown parameters, which are also estimated through ML.

The probability that the market assigned the rating s_j , $j = 1, 2, \dots, S$ to country i , in period t , is:

$$P(R_{i,t} = s_j | x_{i,t}) = P(\alpha_{j-1} < x_{i,t}\beta_* + \mu_i + \varepsilon_{i,t} \leq \alpha_j) = \begin{cases} P(x_{i,t}\beta_* + \mu_i + \varepsilon_{i,t} \leq \alpha_j | x_{i,t}) <= j = 1 \\ P(\alpha_{j-1} < x_{i,t}\beta_* + \mu_i + \varepsilon_{i,t} \leq \alpha_j | x_{i,t}) <= 1 < j \leq S-1 \\ P(\alpha_{S-1} < x_{i,t}\beta_* + \mu_i + \varepsilon_{i,t} | x_{i,t}) <= j = S \end{cases} =$$

⁷ Detailed discussion of the autofitting criteria and options of STATA's REGOPROB2 package are provided in Williams (2006) and Pfarr et al. (2011).

$$= \left\{ \begin{array}{l} \Phi(\alpha_1 - x_{i,t}\beta_*) \leq j = 1 \\ \Phi(\alpha_j - x_{i,t}\beta_*) - \Phi(\alpha_{j-1} - x_{i,t}\beta_*) \leq 1 < j \leq S-1 \\ 1 - \Phi(\alpha_{S-1} - x_{i,t}\beta_*) \leq j = S \end{array} \right\} \quad (3)$$

$\Phi(w)$ denotes, as usual, the cumulative distribution for a multivariate normal evaluated at some vector w . We choose to assign 1 to the best rating class and S to the worst. A rating order increase implies a higher sovereign CDP. Furthermore, $x_{i,t}\beta_*$ is

$$x_{i,t}\beta_* = \sum_{t=1}^T \theta_t D_t + \sum_{l=1}^K \beta_l x_{l,i,t} + \sum_{l=1}^K \sum_{j=1}^{S'} \delta_{j,l} \text{Drat}_j x_{l,i,t}, \quad (4)$$

K is the total number of macroeconomic and financial covariates included in the model. Each might have a global effect on $R_{i,t}^*$ and an effect contingent on the credit rating of each country, at each moment in time, $\delta_{j,l} \text{Drat}_j x_{l,i,t}$. Drat_j is an indicator variable, taking the value 1 if the country's rating is $j \geq 2$, and 0 otherwise. Although autfit (see footnote 7) tests the possibility of different coefficients for each variable and rating class, only the statistically relevant cross-product dummies are retained in the final model (in a similar approach to reducing the model by checking congruency at each stage, embodied in general-to-specific model selection (Doornik, 2009). Notice further that, in equation (4) $S' \leq S$, since the number of rating classes acceptable for efficient estimation may need to be reduced, when considering the subset of rating-varying coefficients, in the no parallel lines setting. It is usual that $S' < S$, as discussed in Pfarr et al. (2011). Finally, time specific dummies and the candidate macro covariates are subject to the same congruency and significance criteria for inclusion in the final model (for a detailed discussion of congruency and congruent models, see Bårdsen et al., 2005).

A note should be made on the construction of rating classes. Following the advice in the literature⁸, we were concerned that an excess number of rating classes might result in imprecise pa-

rameter estimates, as some classes would have very few observations. This would also occur if CDP intervals, associated with very high CDPs, had small lengths. As such, we have divided the observed CDPs in 8 classes: [0%-5%]; [5%-10%]; [10%-15%]; [15%-20%]; [20%-30%]; [30%-40%]; [40%-50%]; [50%-100%]. The associated orders are 1, 2, 3, 4, 5, 6, 7 and 8, respectively, representing the rating levels. We use the same mapping strategy as in Afonso et al. (2007). Adding to this, as discussed in the previous paragraph, we assume $S' < S$. The number of classes considered for efficient estimation of the cross-product dummies' coefficients is only $S' = 5$. The first 4 are equivalent to the above, with the 5th being [20%-100%]. The reduction to these 5 classes, when estimating the δ_j parameters, is supported by the 95% ML confidence interval estimates for cut-off CDPs. Results for $\hat{\alpha}_l$ are presented in the Appendix (Table A1). For the first 5 classes, the lower cut-off is included in the estimated interval. However, for the others, intervals nearly overlap. This is due to few observations for very high CDPs. Hence, it supports our choice to estimate the subset of rating-varying coefficients with $S' = 5 \neq S$.

3. HYPOTHESES DEVELOPMENT

Our research question pertains to the possibility of building a congruent model of sovereign risk based on macro fundamentals. The empirical evidence discussed in subsection 1.1 suggests 14 research hypotheses. The first 12 refer to the possible relevance of specific variables in explaining sovereign risk, the 13th hypothesis refers to time-specific effects, and the 14th to regional contagion:

- H1: *Higher real GDP growth improves sovereign ratings.*
- H2: *Higher public debt to GDP ratios worsen sovereign ratings.*
- H3: *Higher government revenue to GDP ratios worsen sovereign ratings.*

8 Cantor and Packer (1996) observe that a bigger number of classes combined with few to none rating observations in some, would induce failures of the ML estimates to converge. Afonso et al. (2007) also discuss the need to reduce the number of rating classes to increase estimation precision.

H4: <i>Higher external debt to GDP ratios worsen sovereign ratings.</i>	Datavision® data, we have chosen end of quarter Markit® CDPs, for each relevant period.
H5: <i>Higher government deficit to GDP ratios worsen sovereign ratings.</i>	With respect to the macro-financial covariates, we have collected data for all the dimensions referred to in subsection 1.1. Hence, we define the following variables:
H6: <i>Higher inflation worsens sovereign ratings.</i>	
H7: <i>Changes in inflation affect sovereign ratings.</i>	• Fiscal Space: government revenue to GDP (X_5); government balance to GDP (X_6); public debt to GDP (X_3);
H8: <i>Higher global risk worsens sovereign ratings.</i>	
H9: <i>Changes in risk aversion affect sovereign ratings.</i>	• External Imbalances: external debt to GDP (X_4); current account to GDP (X_{10}); changes in the current account to GDP (ΔX_{10});
H10: <i>Higher unemployment worsens sovereign ratings.</i>	• Risk and Contagion: VIX index (X_8); changes in the risk proxy C (ΔX_8); D_{GIPSI} , taking the value 1 for observations from Greece, Portugal, Italy, Spain and Ireland;
H11: <i>The current account balance to GDP ratio affects sovereign ratings.</i>	
H12: <i>Changes in the current account balance to GDP ratio affect sovereign ratings.</i>	• Economic Activity: GDP real growth rate (X_2); Consumer Price Index Inflation (X_7); inflation changes (ΔX_7); unemployment rate (X_9);
H13: <i>Year-specific events affect Sovereign ratings.</i>	
H14: <i>There are regional specific effects in the EA periphery.</i>	• Year-Specific Indicators: D_{09} for 2009; D_{10} for 2010; D_{11} for 2011; D_{12} for 2012; D_{13} for 2013; D_{14} for 2014; D_{15} for 2015;
	• Cross-Product Indicators: $Drat_j x_{l,i,t}$, as explained in section 2.

4. DATA

Our panel is strongly balanced. It comprises the periods between the final quarter of 2007 and the 1st quarter of 2015 ($T = 30$ quarters), and 26 OECD countries ($N = 26$)⁹. Hence, the analysis was conducted with 780 observations per variable. As explained in section 2, the rating classes were derived from CDPs. We have built a unique data set for this purpose. From 2007 to 2013, CDPs were collected from Credit Market Limited (CMA) Datavision® quarterly sovereign risk reports. Given that these became unavailable, with the necessary level of detail, from 2013 onwards, we have used, for the subsequent periods, CDPs computed similarly by Markit®¹⁰. To ensure consistency with CMA

It should be noticed that Beirne and Fratzscher (2013) use the change in the VIX as a proxy for global risk, while we use the VIX index for that. Hence, ΔX_8 has a different role in our model: it is a proxy for global risk aversion. Our hypothesis of a time-varying risk aversion is in accordance with the behavior anticipated in expected utility theory, in Financial Economics, matching proposition 6.C.4 in Mas-Colell et al. (1995). Furthermore, Heinz and Sun (2014) had provided empirical evidence of non-constant risk aversion, following the financial crisis¹¹. Ours is the first model in this literature to include both risk and risk aversion as covariates explaining sovereign risk.

⁹ 10 of the original EA countries (except for Finland), Australia, Czech Republic, Denmark, Estonia, Hungary, Iceland, Israel, South Korea, New Zealand, Norway, Poland, Slovak Republik, Slovenia, Sweden, UK and US.

¹⁰ For detailed information on Markit® data see, in this journal, Oliveira and Santos (2015).

¹¹ It should be noticed that, in our sample, the mean of ΔX_8 is negative, further supporting the need to estimate the impact of a changing risk aversion on sovereign ratings (as shall be done in section 5).

Table 1. Estimation results: panel ordered probit

Variable	Coefficient	t-ratio
$X_{2,it}$	-.1406732***	-4.51
$X_{3,it}$.0138477**	2.15
$X_{4,it}$	-.0027528**	-2.07
$X_{5,it}$	-.0965688***	-3.00
$X_{6,it}$	-.117076***	-3.86
$X_{7,it}$.2205815***	4.28
$X_{8,it}$.0595192***	4.30
$X_{9,it}$.1826575***	4.42
$X_{10,it}$.1553171***	3.25
$D_{10,it}$	1.049514***	4.50
$D_{11,it}$	1.818448***	7.42
$D_{12,it}$	1.777177***	7.13
$D_{13,it}$.4974055**	2.03
$Drat_4 X_{4,it}$.0040576**	2.15
$Drat_5 X_{4,it}$.0047933***	4.65
$Drat_3 X_{5,it}$.1803779***	9.32
$Drat_4 X_{5,it}$.3370346***	8.08
$Drat_5 X_{5,it}$.4670705***	10.11
$Drat_5 X_{10,it}$	-.2306103***	-4.98
$\Delta X_{7,it}$	-.054128	-0.63
$\Delta X_{8,it}$	-.023608**	-2.40
$\Delta X_{4,it}$	-.0070837**	-2.03

Note: *** Refers to 1% significance, ** to 5% significance, * to 10% significance.

All macroeconomic and financial variables were obtained from DataStream, except for X_3 and X_4 (retrieved from OECD statistics)¹².

Despite the discussion in subsection 1.1.5, no proxy for liquidity is included in our model, since we are working with SCDSs instead of bonds. For the latter, meaningful liquidity measures are easy to obtain. For the former, the opacity of OTC derivatives renders any liquidity proxy imprecise (e.g. Markit[®] does not have bid-ask spreads neither for all sovereign entities, nor for the entire sample period).

5. RESULTS AND DISCUSSION

Table 1 reports the estimation results. The 1st column lists the explanatory variables, the 2nd the

coefficients' estimates, and the 3rd the t -ratios. Non-significant year-specific indicators, macro-financial covariates and cross-product dummies were omitted from the final model¹³, as outlined in section 2. The likelihood ratio test for global significance rejects the null hypothesis (at 0.1%): $LR_{obs} = -315.96357$. Furthermore, the test of a pooled ordered probit against a panel model rejects the hypothesis of no gains from using a panel at 0.1% ($LR_{obs} = -166.63$). Thus, congruency of our model is assured.

Results in Table 1 translate to relevant improvements over the previous literature. These improvements are robust to multivariate normality of the random errors (results for the panel ordered logit, provided in Table A2 in Appendix, do not change the conclusions of this section).

12 Given the lack of availability of data, for some covariates, at a quarterly frequency, we follow the literature consensus, using standard interpolation (e.g. Beirne & Fratzscher, 2013).

13 $\Delta X_{7,t}$ is maintained in order to draw a conclusion regarding $H7$.

5.1. Fiscal space

Table 1 clarifies the relevance of the fiscal space covariates. X_3 (at 1%) and X_6 (at 5%) are individually relevant, confirming *H2* and *H5*. Both variables have the expected impact: a better government balance to GDP ratio decreases the probability of a worse rating; a greater weight of public debt in GDP increases that probability. Notwithstanding, both X_5 and its associated cross-product dummies (for $S' \geq 3$) are significant at 1%. The impact of an increase in the government revenue to GDP ratio on the probability of a worse sovereign risk differs per rating class. The parallel lines assumption is not imposed. As such:

- for countries in the two best rating classes, the greater the government revenue to GDP, the lower the probability of a worse sovereign risk

$$(\hat{\beta}_5 = -0.0965688 < 0);$$

- for countries in other rating classes, a greater X_5 increases the probability of a worse sovereign risk

$$\hat{\beta}_5 + \hat{\delta}_{3,5} = 0.0830891 > 0,$$

$$\hat{\beta}_5 + \hat{\delta}_{4,5} = 0.2404658 > 0,$$

$$\hat{\beta}_5 + \hat{\delta}_{5,5} = 0.3705017 > 0.$$

Hence, validity of *H3* is rating contingent. This result merits the attention of policy-makers. Firstly, fiscal consolidation on high risk countries should favor lower government spending instead of higher taxes. Secondly, the negative impact of higher taxes worsens as sovereign risk increases.

5.2. External imbalances

Both the external debt to GDP ratio and the associated cross-product dummies for $S' \geq 4$ are significant (at 5%). Conclusions with respect to *H4* are rating contingent, since the parallel lines hypothesis fails. Our results imply that:

- for countries in the 3 best rating classes, a decrease in X_4 augments the probability of a worse credit rating

$$(\hat{\beta}_4 = -0.0027528 < 0);$$

for countries in the other rating classes, a greater X_4 increases the probability of a higher sovereign risk

$$\hat{\beta}_4 + \hat{\delta}_{4,4} = 0.0013048 > 0,$$

$$\hat{\beta}_4 + \hat{\delta}_{5,4} = 0.0020405 > 0.$$

With respect to the current account balance to GDP ratio, the simultaneous significance of X_{10} and $Drat_5 X_{10}$ (both at 1%) needs to be considered. As such:

- for countries in the first four rating classes, an increase in X_{10} augments the probability of worsening sovereign risk

$$(\hat{\beta}_{10} = 0.1553171 > 0);$$

- for a country in the rating class $S' = 5$, an increase in X_{10} diminishes the probability of a greater sovereign risk

$$(\hat{\beta}_{10} + \hat{\delta}_{5,10} = -0.0752932 < 0).$$

In conclusion, *H11* is confirmed. With respect to changes in the current account to GDP ratio, ΔX_{10} is found to be significant at 10%, with an estimated negative coefficient. Research hypothesis *H12* is confirmed.

A final comment should be made on the results of this subsection. Macroeconomic theory supports the seemingly puzzling results for some estimated coefficient signs. Eliasson (2002) argues that external imbalances could serve as an indicator for the willingness of foreigners to cover the current account gap through loans and foreign investment. Hence, a higher current account deficit would be associated with higher creditworthiness or good economic prospects, consequently, a higher sovereign rating.

5.3. Risk and contagion

Table 1 confirms that X_8 is significant (at 1%), and that its estimated coefficient is positive. A higher global risk increases the probability of a country moving to a worse rating class. With respect to risk aversion (ΔX_8), we conclude in favor of its significance at 5%, but with a negative estimated coefficient. Thus, although agents react adversely to risk, they ask for smaller increases in the risk

premium. We estimate that a higher change in risk aversion lowers the probability of moving to a worse rating. This is consistent with the decline in risk aversion over the sample period, referred to in section 4, when discussing the behavioral basis for this variable (and in footnote 11). Furthermore, $H8$ and $H9$ are confirmed.

Table 1 corroborates $H14$. The regional dummy for the EA periphery is significant at 5%, and $\beta'_{14} = 1.715823 > 0$. The probability of an EA periphery country experiencing a deterioration in credit worthiness is higher than for other advanced economies in our sample. This common effect is interpreted as shift contagion.

5.4. Economic activity space

Research hypothesis $H1$ is confirmed: X_2 is significant at 1%, and its estimated coefficient is negative. An increase in the real growth rate of GDP diminishes the probability of a worse sovereign risk.

Table 1 shows that a higher inflation rate augments the probability of a worse rating. $\hat{\beta}_7 = 0.2205815 > 0$. X_7 is statistically significant (at 1%), confirming $H6$. There is no support for claims regarding the relevance of changes in inflation. ΔX_7 is not significant (at 10%), rejecting $H7$.

Table 1 indicates that the estimated coefficient of the unemployment rate is positive and significant at 1%. A country with a higher unemployment rate has a higher probability of moving to the worse rating class. $H10$ is confirmed.

5.5. Year-specific dummies

Year-specific indicators are included in our model to account for episodes in the sample period, not controlled for by other variables. Table 1 reveals that indicators matching 2010, 2011, 2012 and 2013 are significant, with positive estimated coefficients. Research hypothesis $H13$ is confirmed.

The indicators' estimated coefficients are increasing from 2010 to 2012 and decrease in 2013. This is compatible with the evolution of the EA crisis: the 1st rescue package for Greece occurred early in 2010; 2011 witnessed the Irish and Portuguese bail-outs, as well as the inversion in Italy's yield curve; in early 2012, Greece's 2nd rescue package was implemented, and the EU bailed-out the Spanish banking system. From 2013 onwards, the crisis seems to have been softened, most likely due to regulatory changes and to a more pro-active ECB policy in favor of financial stability (De Santis, 2016).

CONCLUSION

This paper improved on the econometric methods used in the literature on the macro-financial foundations of sovereign risk. A congruent model, including covariates for all dimensions suggested by earlier authors, was achieved. The panel ordered probit, without the parallel lines assumption, using SCDSs implied ratings as the sovereign risk proxy, clarified the puzzles in earlier empirical papers. Thus, SCDSs implied ratings match economic expectations based on fundamentals, showing no evidence (at a quarterly frequency) of credit risk mispricing.

In particular, we have shown that a higher real growth rate of GDP decreases sovereign risk (improving on, e.g., Beirne & Fratzscher, 2013), a lower public debt to GDP ratio decreases sovereign risk (contrary to, e.g., Afonso et al., 2015), a higher government budget surplus benefits ratings (improving on, inter alia, Yuan & Pongsiri, 2015), and that lower inflation and unemployment rates diminish sovereign risk (improving on Kriz et al., 2015) and on Aizenman et al. (2013). Relaxing the parallel lines assumption allowed us to conclude that a worsening of the ratios of the external debt or the current account to GDP only deteriorate ratings for countries in high risk classes. Such risk class-contingent conclusions had not been addressed in the previous literature. Similarly, we improve on Kriz et al. (2015) by concluding the role of the ratio of government revenue to GDP is also varying with a country's rating group. Moreover, we have shown the relevance of controlling for risk aversion and global risk simultaneously, a procedure no other paper had followed before in this literature.

Policy implications of these results were carefully outlined. We plan to develop this research by assessing the impact on ratings of filtered spreads. We also intend to consider other forms of contagion in our analysis, namely volatility contagion.

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APPENDIX

Table A1. Confidence Intervals for α_i

	Lower 95%	Upper 95%
Cut 1	-2,370774	3,291392
Cut 2	2,014583	7,855181
Cut 3	8,927898	14,33164
Cut 4	14,9639	19,03504
Cut 5	19,98533	26,03503
Cut 6	26,73192	33,00836
Cut 7	26,3271	34,0424

Table A2. Estimation results: panel ordered logit

Variable	Coefficient	T-ratio
$X_{2,it}$	-.2404042***	-4.34
$X_{3,it}$.0168277**	2.1
$X_{4,it}$	-.0058903**	-2.55
$X_{5,it}$	-.1616363***	-2.91
$X_{6,it}$	-.2201923***	-3.92
$X_{7,it}$.4043985***	4.4
$X_{8,it}$.1009374***	3.98
$X_{9,it}$.3044485***	4.09
$X_{10,it}$.3245383***	3.75
D_{GIPS}	3.002686***	2.28
$D_{10,it}$	1.860224***	4.52
$D_{11,it}$	3.351913***	7.34
$D_{12,it}$	3.239343***	7.10
$D_{13,it}$.9443354**	2.09
$Drat X_{4,it}$.0085343**	2.57
$Drat X_{5,it}$.0097293***	5.03
$Drat X_{3,5,it}$.3270163***	8.99
$Drat X_{4,5,it}$.6128336***	7.78
$Drat X_{5,5,it}$.8501224***	9.60
$Drat X_{5,10,it}$	-.4498397***	-5.33
$\Delta X_{7,it}$	-.1163345	-0.76
$\Delta X_{8,it}$	-.0409116**	-2.28
$\Delta X_{4,it}$	-.00116338**	-1.89