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## Ratings matter: Announcements in times of crisis and the dynamics of stock markets

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### ABSTRACT

In this paper we propose a novel approach in analysing the impact of changes in sovereign credit ratings on stock markets. We study the evolution of a segmented form of the stock market index for several crisis-hit countries, including both European and Asian markets. Such evolution is modelled by a homogeneous Markov chain, where the transition probabilities from one starting level of the index to a new (lower or higher) level in the next period depend on some explanatory variables, namely the country's rating, GDP and interest rate, through a generalised ordered probit model. The credit ratings turn out to be determinant in the dynamics of the stock markets for all three European countries considered - Portugal, Spain and Greece, while not all considered Asian countries show evidence of correlation of market indices with the ratings.

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## 1. Introduction

Thomas Friedman acknowledges in 1995 from the columns of the New York Times that “In the 1990s the most important visitor a developing country can have is from Moody’s Investors Service Inc.” (see [Friedman, 1995](#)). He continues warning that “Moody’s is the credit rating agency that signals the electronic herd of global investors where to plunk down their money, by telling them which countries’ bonds are blue-chip and which are junk”. He concludes with the famous often cited statement “That makes Moody’s one powerful agency. In fact, you could almost say that we live again in a two-superpower world. There is the U.S. and there is Moody’s. The U.S. can destroy a country by leveling it with bombs; Moody’s can destroy a country by downgrading its bonds.” More recently, since the sub-prime crisis and the bankruptcy of Lehman Brothers on the

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15th of September 2008, the role of these agencies has come to the spotlight in the debates of politicians, economists and the general public (Sangiorgi and Spatt, 2017).

Credit rating agencies play an important role in reducing information asymmetries in capital markets, providing significant information to (institutional) investors to evaluate risks. This makes them fundamental players in a globalised financial world (Levich et al., 2012), and the economic consequences of their activity should be analysed. The increasing number of private and public issuers and debt products, and the growing requirements applied to investors, banks and other financial institutions about risks, portfolios' compositions and capital adequacy, explain the growing importance of their role (Al-Sakka and Gwilym, 2010).

Rating agencies gained a particular negative reputation following downgrades of crisis-hit countries in Europe, becoming the synonym of austerity measures imposed in order to balance public budgets. All this social cost has been undertaken to regain the confidence of the markets in the ability to pay public debt. This confidence was, and still is, summarized by a rating of a few letters assigned by the main rating agencies - Standard & Poor's (S&P), Moody's and Fitch - that determine whether a country is trustworthy or not. The credit ratings expressed by these companies have become so entrenched and common in the public knowledge that even the way we talk and think about credit risk has absorbed the rating scale of the famous 'AAA' and the infamous 'junk'.

Rating agencies were initially seen as private companies with a decentralized opinion, immune to governmental interference, and acting in a competitive market where their credibility was their strongest asset (White, 2010). As a matter of fact, by the end of the 90's the market was dominated by the 'Big Two': S&P and Moody's represented 90% of the market share (Bruner and Abdelal, 2005), with Fitch as a distant third. Moody's has been publishing ratings since 1909 and S&P since 1923; according to Mora (2006), their long life and the fact that investors pay to subscribe to their reports implies that their activity and function - providing an estimation of default probability of borrowers - must be of some value. In recent years, the importance of rating agencies increased due to the globalization of financial markets and the widespread attention to their ratings in banking and financial regulation. Consequently, the debate about the behavior of rating agencies-focused on issues like lack of competition and transparency, or the conflict of interests with issuers - and their role has gained increasing interest (Stolper, 2009; Mathis et al., 2009; Hunt, 2009; Opp et al., 2013; Fuchs and Gehring, 2017).

The greatest impact of these agencies is, by far, the way they influence the access to institutional investors all over the world, acting as *de facto* gatekeepers to huge sums of capital. Indeed, credit ratings have been crucial in assigning default probabilities to debt issuers and, consequently, the access of governments to capital inflows. As Ferri et al. (1999) outline: "(...) besides affecting the cost at which issuers can borrow, ratings determine the extent of potential investors. Specifically, regulations and statutes either forbid institutional investors to invest in assets carrying ratings below a certain level or they require extra capital to be posted. These investments are referred to as 'below-investment-grade' or 'speculative assets.'" Bruner and Abdelal (2005) provide an interesting discussion about the new status enjoyed by these agencies, the power that derives from it and the quasi-symbiotic relationship established between the public sphere of governments and these private companies.

One of the motivations for this work is the fact that the credit rating of a sovereign country heavily influences private-sector issuers of debt, what Bruner and Abdelal (2005) call the 'sovereign ceiling'. Since only rare exceptions of private companies' foreign currency debt is issued with a higher credit rating than their sovereign's (Borensztein et al., 2013), the rating assigned to sovereigns affects all the other debt issues in the world. In fact, a sovereign debt rating downgrade is followed by a rising in the cost of debt for private issuers, with consequences both in financial and real markets (Almeida et al., 2017; Bedendo and Colla, 2015).

In the present study we explore data on the European and Asian crises to investigate whether there is evidence of an impact of the changes in sovereign credit ratings on the stock market indices of several countries, and in particular whether and how the credit ratings affect the probabilities of the indices increasing or decreasing in subsequent periods, as modelled by an homogeneous Markov chain.

The next section presents some background on research in this field, while Section 3 proposes a model for assessing the impact of ratings changes on stock market dynamics. Section 4 presents the application to the European and Asian data, while the last section concludes.

## 2. Background

### 2.1. Determinants of credit ratings

The sovereign credit rating literature is extensive. Its start can be traced back to the work of Cantor and Packer (1996) who describe and quantify the determinants and impact of sovereign credit ratings given by the major U.S. based rating agencies, Moody's and S&P. They find that the ordering of risks that credit ratings imply is broadly consistent with specific macroeconomic fundamentals, particularly per capita income, GDP growth, inflation, external debt, level of economic development and default history.

Later studies examine the quantitative determinants of S&P and Moody's sovereign ratings before and during the global financial crisis, finding evidence that the most relevant economic variables are the Gross National Product (GNP) per capita and inflation (Bissoondoyal-Bheenick, 2005); GDP (per capita and growth rate), inflation and unemployment rate,

government and external debt, fiscal balance and political and historical factors (Afonso et al., 2009); and finally GDP growth and volatility, GNI, Import/Exports, Inflation, Rule of Law and Government Debt (Basu et al., 2013).

## 2.2. Performance of credit rating agencies

In recent years the role of credit rating agencies in international markets has been the subject of deep analysis, debate and also criticism, and a large body of work assessing their performance has been published. In fact, the agencies' role in reducing information asymmetries among financial agents - by condensing deep debt issuer's analysis into one grade that provides a global risk evaluation - is crucial in the current complex financial world. Ferri et al. (1999), in line with Radelet and Sachs (1998), demonstrate that the pro-cyclical nature of rating agencies' sovereign rating may have contributed to aggravate the East Asian financial crisis. However, later Mora (2006) contradicts the previous findings, concluding that rating agencies did not aggravate the Asian crisis.

Altman and Rijken (2004) address the investor's perception of rating agencies being very slow to adjust their ratings. An explanation of this perception is the "through-the-cycle" rating methodology that agencies apply in their assessments, while investors have a "point-in-time" perception of creditworthiness. The authors conclude that rating agencies' methodology suppresses the sensitivity of the ratings to short-term fluctuations in credit quality.

Several authors investigate the agencies' influence and the independence of their evaluations. Bruner and Abdelal (2005) focus on the relationship between sovereign governments and the rating agencies' authority in the bond markets, discussing possible alternatives in order to improve their accountability, given the immense power they have gathered, especially with institutional investors. Tennant and Tracey (2013) find a negative bias in S&P assignment of rating upgrades and downgrades to low and lower-middle income countries, while Fuchs and Gehring (2017) find evidence of home bias in the production of the sovereign ratings, as a result of political and cultural factors.

Finally, Sangiorgi and Spatt (2017) analyse from an economic and regulatory perspective several limitations of credit rating agencies as risk evaluators, comparing with other gatekeeper providers, as the external auditors. In spite of concluding that ratings play an important role as information provider for capital markets players, the authors identify several weaknesses associated with the conflict of interest by paying for information.

## 2.3. Impact of rating changes on the economy and the financial markets

Various scholars have investigated the relationship between changes in credit rating quality and economic and financial national indicators, using different datasets and methodologies. Kräussl (2005) confirms that changes in sovereign rating statistically influence size and volatility of emerging capital markets, but results depend on the type of changes (downgrade/upgrade), while Hooper et al. (2008) examine the impact of rating changes on stock and foreign exchange market returns and volatility, for 42 countries during the period 1995–2003. They conclude that the overall impact is significant, but asymmetric, with amplified effects for downgrades, in emerging markets and during periods of crisis. This conclusion is also confirmed by Brooks et al. (2004), where the effects on stock indices are significant when a downgrade occurs. In an event-study setting, they also find that emerging markets are not sensitive to rating changes.

Kim and Wu (2008) come to the conclusion that long-term ratings are most important for the development of financial sectors in emerging economies. They remark that foreign currency long-term ratings provide the most important contribution for international capital inflows and, consequently, domestic financial market development. This argument is supported by observing that the three forms of capital inflows studied (foreign direct investment, international banking flows and portfolio flows) significantly increased as foreign currency long-term ratings of emerging market sovereigns improved.

Other authors investigate the response of yield spreads of credit default swaps (CDS) to rating announcements, both in emerging markets (Ismailescu and Kazemi, 2010) and in developed countries (Afonso et al., 2012), while De Santis (2014) analyses the developments of euro area sovereign yield spreads, all finding that there is a significant response to changes in both the credit rating notations and in the outlook, with strong spillover effects. More recently, Lee et al. (2018), using a large sample of S&P rating notch and watch changes in the U.S. market, find that the CDS price has predicting power for negative events (downgrading) and contains specific information that contributes to price discovery.

Al-Sakka and Gwilym (2013) examine how the foreign exchange markets reacted to sovereign credit events prior to (2000–2006) and during the crisis (2006–2010). Their results support the view that rating agencies' signals do affect the own-country exchange rate. They also identify strong spillover effects to other countries' exchange rates in the region. The authors also mention that market reactions and spillovers are far stronger during the financial crisis period than pre-crisis.

Baum et al. (2016) study the effects of rating downgrades on the value of the Euro currency and the sovereign bond yields during the crisis. The main findings are that there is little impact on the exchange rate, while downgrades affect significantly the yields. The authors conclude that the common currency showed a stabilising effect, having prevented reduction of foreign investment in the Eurozone in spite of negative rating announcements; they highlight also that the single currency allowed rebalancing of investors' sovereign bond portfolios without changing the currency, by reallocation of investments within the Eurozone towards stronger states.

Chen et al. (2016) analyse the relation between sovereign credit rating changes and economic growth, with a sample of S&P rating events for 103 countries, over the 31-year period between 1982 and 2012, with 271 positive changes and 102

negative rating events. Using a dynamic panel data methodology, they conclude that the growth rate of real per capita GDP exhibits a statistically significant response to changes in sovereign ratings.

[Kaminsky and Schmukler \(1999\)](#) study the effects of news, among which rating announcements, on extreme markets movements. Based on data from nine Asian countries, their analysis concentrates on the twenty largest market changes (jitters) registered in each country between the beginning of 1997 and mid 1998, finding evidence of large spillover and contagion effects, as well as herding behaviour of the investors. Similar results are found in [Baig and Goldfajn \(1999\)](#), based on five Asian countries, reinforcing patterns of contagion during the crisis, as well as large cross-country correlations in the currency and equity markets, even after controlling for own-country news and other fundamentals.

Although the present work can be categorised in this same area of research - since it aims to measure the effect of rating changes on the stock markets - we develop a different approach from what was previously studied, as we explain in the following section.

### 3. A model for stock markets dynamics

The objective of this paper is to study whether the evolution of the sovereign credit ratings effectively acts upon the dynamics of stock markets, and if so, to quantify such influence, after controlling for other relevant variables.

The natural high volatility of the market indices does not match the substantially lower frequency of rating change announcements. Therefore, in order to capture the medium-term impact of the rating levels and the subsequent updates during a crisis period, we propose to focus not on the actual value that the stock market index is attaining in each period of observation, but rather on the qualitative performance of the index, represented by a certain number of classes or levels (ordered from 'very low' to 'very high').

The choice to focus on the level of the stock indices, instead of the changes, has a peculiar reason. In fact, if we assume that the stock index represents the (synthetic) fundamental value of the companies, studying its level is equivalent to study the evolution of the market value of the stocks that belong to the index. If the economic conditions are worsening during a sovereign crisis, most likely also the value of the companies will be affected. More generally, with our analysis we are willing to capture the effect of the rating on the medium-term performance, or fundamental value, of the country stock indices, and to assess the probability of an increase/decrease of the market value of asset given a certain initial level of the index and the rating. In this view, each class or level of the index can be interpreted as a qualitative assessment of the different market value of the (equity) assets in a given country, in the medium term.

The low frequency of the announcement points to the use of a discrete-time setting. The segmented form of the stock index allows to represent its dynamics through a discrete-state process. In order to model the transitions of the stock index levels across different classes over time, a possible choice is therefore a discrete-time Markov chain.

The Markov structure implies that the level of the index in the next time period only depends on its current level, and not on the levels in the previous periods, considering that all needed information about the market is included in the current value of the index. This is a fairly common assumption in the modelling of financial series, which reflects the so-called weak form of market efficiency, where past patterns of prices have no forecasting value (see for instance [Dixit and Pindyck, 1994](#)). As a matter of fact, among the stochastic processes most widely used to model financial time series, many benefit from the Markov property, like for instance the random walk and the Brownian motion (see for example [Taylor and Karlin, 2014](#)). A further discussion of the Markovian properties of financial time series can be found in [Chen and Hong \(2012\)](#).

The calculation of a so-called transition matrix allows to obtain, at each point in time, the probability of the stock market index moving to an upper or lower level (or remaining at the same level) in the following period, conditional on the current position and some other relevant explanatory variables. Given the natural ordering of the index levels, the probabilities of transition can be computed according to an ordered qualitative dependent variable models, and then used to reconstruct the corresponding discrete-time transition matrix.

Similar methodologies have been used in other contexts: for instance, [Hausman et al. \(1992\)](#) use ordered probit models to estimate the conditional distribution of transaction stock price changes; [Nickell et al. \(2000\)](#) and [Hu et al. \(2002\)](#) model the distribution of credit ratings changes by ordered qualitative models, proposing the derivation of ad-hoc transition matrices as an alternative to the rating matrices provided by credit rating agencies. In all previous cases the transition probabilities are constant over time, hypothesis that corresponds to a specification of an homogeneous Markov chain dynamics. [Kiefer and Larson \(2007\)](#) present a test for this restriction, applying the proposed methodology to a set of different data sets, concluding that in most cases "transitions can be usefully modelled as Markovian over periods of useful length". Later, [Feng et al. \(2008\)](#) improve on the previous models for credit rating transitions, proposing two alternatives where transition matrices are allowed to vary over time (time-varying parameter probit and stochastic parameter probit models).

For the modelling of stock markets dynamics, in the present work we follow an homogeneous specification of the Markov chain, which corresponds to assuming that the dynamics of the stock market index in different periods during the debt crisis is comparable, specifically the probability of the index jumping from one level to the other remains the same, whether we are at the beginning, the middle or at the end of the crisis period. This hypothesis is rather widespread, as mentioned earlier, and can easily be tested in a cursory way, for instance, by introducing a series of dummy variables for a certain number of time frames (e.g. beginning, middle, end of crisis), and testing their significance. In case the time effects turn out to be

significant - and according to the objectives of the study - more sophisticated time-varying models such as those proposed by Feng et al. (2008) may be considered.

The model for market dynamics can be derived from a latent variable model, where the latent variable  $y^*$  represents the unobservable underlying “quality/reliability” of the market - or the confidence of the investors in the market. The observed value of the level of the market index, say  $y$ , is considered to be a “signal” of the unobserved “quality”  $y^*$ .

The standard ordered probit (OP) model<sup>1</sup> treats  $y$  as an ordered response variable taking values  $1, 2, \dots, J$  for some known integer  $J$ ; the values represent the different levels of the stock index, where  $y = 1$  denotes the lowest level and  $y = J$  the highest. The unobservable confidence in the market is determined by a set of explanatory variables  $\mathbf{x}$  as  $y^* = \mathbf{x}\boldsymbol{\beta} + e$ , where the error is assumed to follow a standard Normal distribution, conditional on the set of regressors  $\mathbf{x}$ . The market index level  $y$  is linked to the underlying “quality” of the market  $y^*$  as follows:

$$y = j \quad \text{if} \quad \alpha_{j-1} < y^* \leq \alpha_j \quad \text{for} \quad 1 \leq j \leq J$$

The  $\alpha_j$  are a set of unknown ordered cut-off points, where  $\alpha_{j-1} < \alpha_j$ ,  $\alpha_0 = -\infty$  and  $\alpha_J = +\infty$ .

The conditional distribution of  $y$  given  $\mathbf{x}$  can be derived from the standard normal assumption for  $e$ , obtaining the following response probabilities<sup>2</sup>:

$$\Pr(y = j|\mathbf{x}) = \Phi(\alpha_j - \mathbf{x}\boldsymbol{\beta}) - \Phi(\alpha_{j-1} - \mathbf{x}\boldsymbol{\beta}) \quad 1 \leq j \leq J$$

They represent the probabilities of the stock market index falling into each of the  $J$  possible levels in a certain period, given the set of specified regressors  $\mathbf{x}$ .

By including among the regressors the level of the index in the previous period, the estimation of the ordered probit probabilities corresponds to the estimation of the homogeneous Markov chain transition probabilities from one index level in one period to the other levels in the next period, as a function of the remaining regressors included in vector  $\mathbf{x}$ . For example, let  $\mathbf{x} = [\mathbf{z}, w]$ , where  $\mathbf{z}$  is a set of country-specific characteristics and  $w$  is the level of the index in the current period; then  $\Pr(y = 4|\mathbf{z} = \mathbf{z}_0, w = 3)$  corresponds to the probability of transition of the index from the current level 3 to level 4 in the next period, conditional on the remaining characteristics taking up the value  $\mathbf{z}_0$ .

The parameters  $\alpha$  and  $\beta$  can be identified and estimated by standard maximum likelihood, supposing that transitions for a specific country are independent over time conditional on  $\mathbf{x}$ , similar to the assumptions made in Nickell et al. (2000).

From the parameter estimates, the above probabilities can be obtained, together with the estimated partial effects of the regressors on each of the  $J$  probabilities. Both the probabilities and the partial effects need to be evaluated at some sensible values for the regressors, for example at the mean value of each variable, or at some specific value of interest.

The standard OP model can be generalised, allowing the effects of the regressors to vary with the levels, i.e. specifying the response probabilities as:

$$\Pr(y = j|\mathbf{x}) = \Phi(\tilde{\alpha}_j - \mathbf{x}\boldsymbol{\beta}_j) - \Phi(\tilde{\alpha}_{j-1} - \mathbf{x}\boldsymbol{\beta}_{j-1}) \quad 1 \leq j \leq J$$

giving rise to the generalised ordered probit (GOP) model under the condition that  $\tilde{\alpha}_{j-1} - \mathbf{x}\boldsymbol{\beta}_{j-1} < \tilde{\alpha}_j - \mathbf{x}\boldsymbol{\beta}_j$ . The GOP specification can be tested against the standard OP model under the parallel lines assumption (PLA), i.e. imposing that  $\boldsymbol{\beta}_j = \boldsymbol{\beta} \quad \forall j$ .

The full GOP model - where all regressors show level-dependent effects - contains a very large number of parameters, and can be simplified relaxing the PLA only when necessary. This partial specification<sup>3</sup> still allows the relevant regressors to have level-varying coefficients, but avoids the inclusion of redundant parameters for those variables that do not violate the PLA. The partial PLA for individual regressors can be tested with likelihood-ratio-type tests.

## 4. Application

### 4.1. Data description

With the objective of capturing periods of crisis with many credit rating changes and high instability of stock markets, three southern European countries - Portugal, Spain and Greece between 2009 and 2014 - and two Asian countries - Indonesia and South Korea, in the period 1997–2003 - were selected. This choice of different countries in separate continents and through different periods of time allows us to test the regional differences in the dynamics of the crisis, looking at areas with different institutional characteristics.

The chosen European countries are all part of Southern Europe, somewhat similar, with a recent post-industrialisation process of a specialization in services (tourism, financial services, etc.). In addition, following the 2008 economic and financial crisis, they needed external financial support and adopted global and sectoral adjustment programs (for example restructuring in banking sector).

Among the countries affected by the Asian crisis, South Korea and Indonesia are large unified countries with a complex and integrated economic environment, as opposed to Hong-Kong and Singapore, which resemble city-states with very

<sup>1</sup> See for instance Boes and Winkelmann (2006).

<sup>2</sup> Similar to binary response models, vector  $\mathbf{x}$  does not contain a constant term to allow identification of the cut-off parameters  $\alpha_j$ .

<sup>3</sup> See for example Fuellerton and Xu (2012).

specialized sectors integrated in the global economy, and therefore with a different nature (Radelet and Sachs, 1998; Desai, 2014).

Throughout this analysis we use S&P long-term ratings for Portugal, Spain, Greece, South Korea and Indonesia, extracted from Bloomberg. Sovereign credit ratings of all countries have been converted into a numerical scale from 21 (AAA) to 0 (Default) as can be seen in Table 1. Other authors propose a different credit rating conversion: Afonso et al., 2012 use a linear scale from 1 (from D to CCC+) to 17 (AAA), and Kim and Wu (2008) use a similar conversion of 0 (D or SD) to 20 (AAA). The time series evolution of the ratings for the five countries considered in the study are reported in Fig. 1.

Stock exchange indices were obtained from Bloomberg too. They are expressed as weekly closing values (on Fridays) of PSI20 (Portugal), IBEX (Spain), ASE (Greece), KOSPI (South Korea) and JCI (Indonesia) - the European indices taking values between the 2nd of January 2009 and the 28th of March 2014, while the Asian indices taking values between the 3rd of October 1997 and the 26th of December 2003. The timespan of European indices was chosen to capture the period of the Eurozone crisis, whereas the Asian indices time frame captures the period with greater volatility in credit ratings derived from the East Asian Financial crisis. Both periods are fertile in credit ratings transitions and volatility in the stock market, as observed in Fig. 1.

Besides the ratings, other macroeconomic variables (GDP and interest rate) were chosen in line with previous studies on the determinants of stock market indices (see for example Cheung and Ng, 1998). The Gross Domestic Product (GDP) of all countries comes from the *OECD StatExtracts*<sup>4</sup> from the Organization for Economic Co-Operation and Development, and are quarterly values of national accounts in billions of US dollars - expenditure approach - at current prices, current PPPs, annual levels and seasonally adjusted.

The following interest rates were used: for all European countries the 6-month EURIBOR rate,<sup>5</sup> while for South Korea and Indonesia the OECD country-specific 3-Month Interbank Rates.<sup>6</sup> The time series of the EURIBOR Interest Rate is shown in Fig. 2, while the country-specific Interbank Rates for Asia are reported in Fig. 3.

#### 4.2. Analysis of variables

The chosen weekly frequency of the index values reflects the objective of measuring the possible effect of rating changes on market dynamics. As noticed earlier, given that the rating changes present a rather low frequency (see Fig. 1), it seems more likely that ratings affect the medium term level of the stock market indices, rather than the much more volatile daily values. Therefore, the weekly or monthly changes of the stock indices seem to be better candidates for the present study. Given that monthly changes would give rise to very few observations in the period of interest, and most likely would not capture the actual dynamics of the market at a time of great instability, we chose weekly observations.

The original series of stock exchange indices were studied before proceeding to their discretisation. Of all countries, only Greece presented a structural break: the series stagnates between April 2010 and March 2011, and then starts again some dynamics on substantially lower levels. During April 2010, the Greek Government bonds have been downgraded from investment grade status to non-investment grade from the major Rating agencies. The non-investment grade status, however, forced the intervention of the European Commission and the International Monetary Fund, launching a bailout intervention to prevent a Greek sovereign default. Various tests were performed (Zivot-Andrews, Gregory-Hansen, Markov switching dynamic regressions), having located the regime shift on different dates in the period comprised between April 2010 and August 2011. In the present study the change in regime has been set at the beginning of April 2011, following the stagnation period. All analyses for Greece are therefore run separately for each period (before and after April 2011).

Some summary descriptive statistics about the indices are reported in Table 3, together with some test results. Tests for stationarity of the series, the presence of structural breaks and the number of lags in the autoregressive structure were performed. All series turned out to be non-stationary; only Greece shows significant trend, negative in the first period of observation and positive in the second; all series present autoregressive structure of order one, in line with the Markov property.

The segmented form of the market index is obtained dividing the range of each index in the period of observation into  $J$  equally spaced intervals (classes), where Class 1 includes the lowest levels of the index attained during the period, and Class  $J$  the highest. In the application, five index levels are considered, the full matrix of empirical transitions between classes for Portugal being shown in Table 2, for the sake of the example. All transitions lie on the three main diagonals; going to a finer definition of the levels, for instance taking  $J = 10$ , would exacerbate even further the sparseness of the matrix, and is therefore not a suitable choice for the present case. Table 3 reports some statistics about the classified series, showing very similar patterns in terms of transitions in all countries.

Finally, the possible correlation of the ratings with time trends is investigated. At the bottom of Table 3, the correlation coefficient of the ratings with a time variable counting the years from the beginning of the crisis is reported for each country.

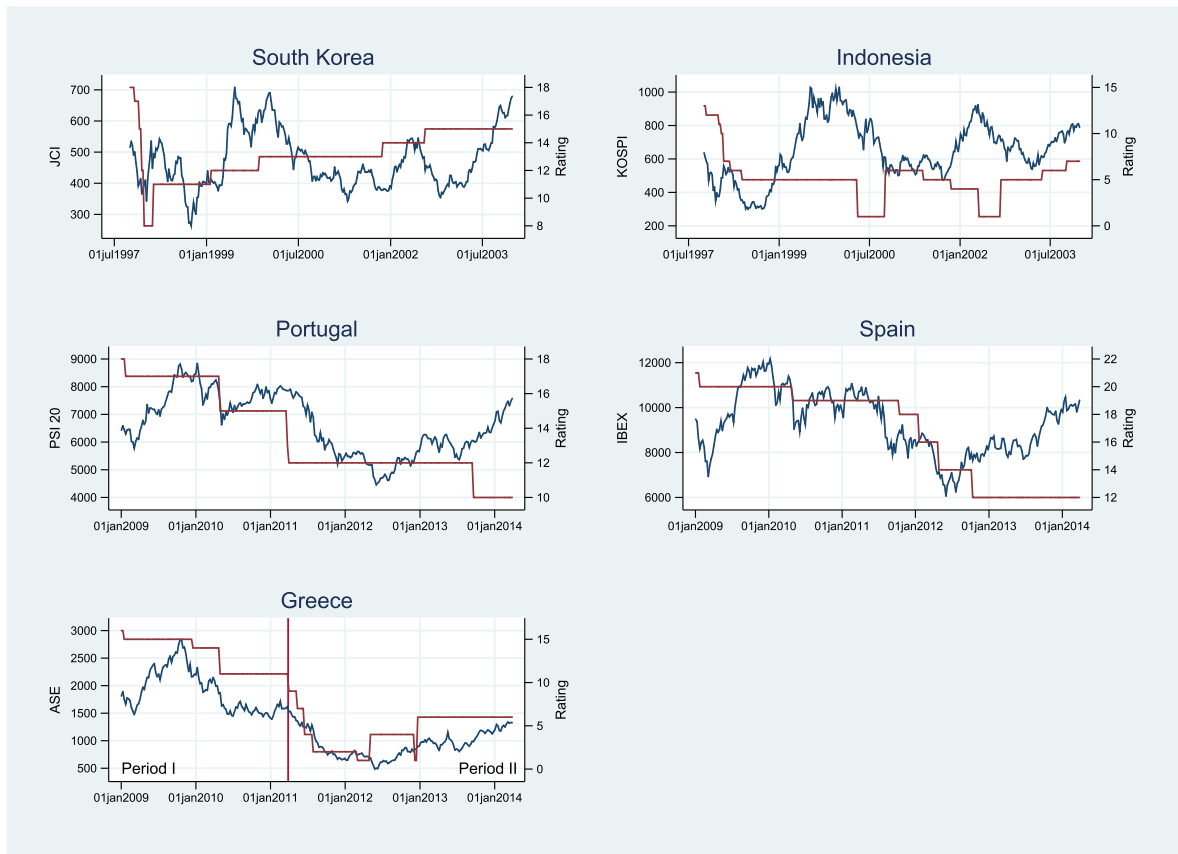
<sup>4</sup> *OECD StatExtracts*: <http://stats.oecd.org/index.aspx?queryid=350#>.

<sup>5</sup> Euro Interbank Offered Rate: unweighted average rate calculated by Reuters since 30 December 1998 according to the act/360 method. Daily quotations downloaded from the Deutsche Bundesbank Eurosystem page [https://www.bundesbank.de/Navigation/EN/Statistics/Money\\_and\\_capital\\_markets/Interest\\_rates\\_and\\_yields/Tables/table\\_zeitreihenliste.html?id=16000](https://www.bundesbank.de/Navigation/EN/Statistics/Money_and_capital_markets/Interest_rates_and_yields/Tables/table_zeitreihenliste.html?id=16000) on 7th August 2018.

<sup>6</sup> Organization for Economic Co-operation and Development, 3-Month or 90-day Rates and Yields: Interbank Rates for the Republic of Korea [IR3TIB01KRM156N] and for Indonesia [IR3TIB01IDM156N], retrieved from FRED, Federal Reserve Bank of St. Louis; <https://fred.stlouisfed.org/series/IR3TIB01KRM156N>, August 7, 2018; monthly averages.

**Table 1**  
Standard & Poor's credit ratings converted into a numerical scale.

AAA	21	BBB+	14	CCC+	5
AA+	20	BBB	13	CCC	4
AA	19	BBB-	12	CCC-	3
AA-	18	BB+	11	CC	2
A+	17	BB	10	SD	1
A	16	BB-	9	D	0
A-	15	B+	8		
		B	7		
		B-	6		



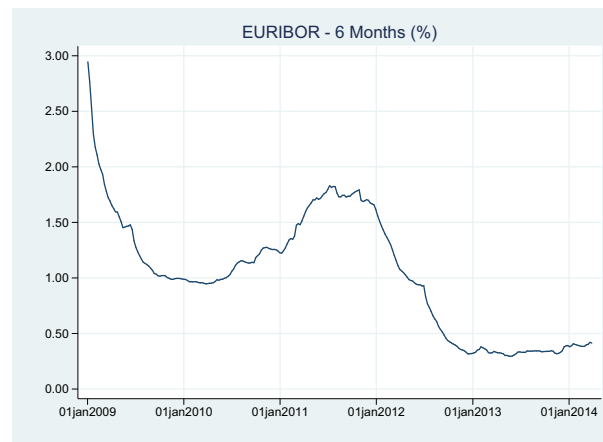
**Fig. 1.** Standard & Poor's credit ratings and Stock Market Indices. Periods of observation: from 01/10/1997 to 26/12/2003 for Asian countries, from 01/01/2009 to 28/03/2014 for European countries. Greece: Period I - until 31/03/2011; Period II - from 01/04/2011.

For the case of European countries, a very strong negative correlation is present, around 90% for Portugal and Spain, about 84% for Greece in the first period.

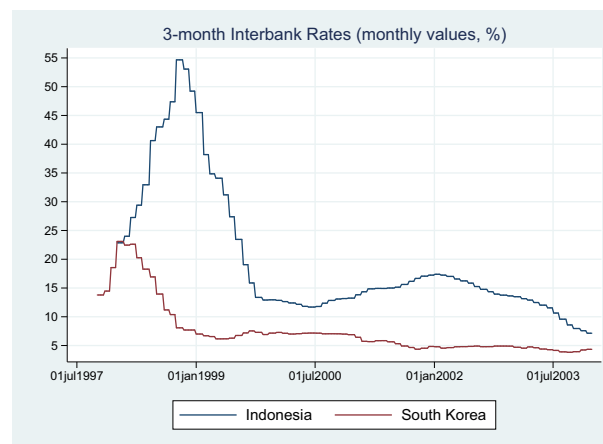
#### 4.3. Model specification and empirical results

Table 4 reports the estimation results of the ordered probit models for the five countries under consideration, where the response variable is the ordered level of the stock market index. The regressors common to all models are the country's credit rating, the GDP and the previous level of the index. In the case of the European countries, the EURIBOR 6-month interest rate is also considered, while country-specific 3-month interbank rates are used for South Korea and Indonesia.

This initial model underwent several robustness checks, both in terms of the Markov transitions and of the ordered response regression, reported in the bottom panel of Table 4. In order to test the homogeneity assumption for the Markov chain, the significance of the time variable (years from crisis onset) in the ordered probit model is tested. Regression results show that the time effects are never significant, i.e. that the dynamics of the stock market does not depend on the phase of the crisis, once accounted for the ratings and the other regressors.



**Fig. 2.** EURIBOR 6-month interest rate. This variable was only used for the estimation of the European Countries' models.



**Fig. 3.** Inter-bank 3-month rates for Asian countries. These variables were only used for the estimation of the Asian Countries' models.

For what concerns the specification of the ordered probit, heteroskedasticity-robust standard errors have been used. The potential endogeneity of the Rating variable has been tested through joint estimation of the main ordered probit regression and a reduced form linear regression for the Ratings,<sup>7</sup> where a time effect - representing the phase of the crisis period - acted as instrument. In all countries the time effect is the years since crisis inception, but in the case of Greece, due to the shorter time span of the two sub-periods of analysis, a Quarter and a Semester effect have been used respectively in Period I and II. The endogeneity test has been performed by testing the significance of the correlation of the errors of the two equations, which in all cases resulted non significant, confirming the exogeneity of the Ratings. While Table 4 reports only the p-value of the test, the full instrumental variable estimation results are reported in Appendix, in Tables 11 and 12.

Whenever regressors were individually not significant at 10%, their joint significance has been tested through exclusion restrictions, allowing to reach a reduced model including only significant variables. In all tests the restrictions were considered valid (see again Table 4). The parallel lines assumption (PLA) of the ordered probit has been tested in the reduced model through an approximate likelihood-ratio test of equality of coefficients across response categories. Results, at the bottom of the Table, show that models for most countries do not violate the hypothesis, however in the case of Greece (period II) and South Korea the Rating variable needs to be modelled allowing for level-varying effects.

After this first exploratory model, a final specification was obtained reducing the set of regressors only to those that were significant, and using a generalised ordered probit (GOP) relaxing the PLA for the Ratings in the cases of Greece period II and South Korea. The estimation results of the final model are reported in Table 5. Notice that in the case of the GOP some of the level-dependent coefficients turned out to be not significant, and were subsequently constrained to be equal to zero when using the model for prediction. The imposed constraints were tested via likelihood-ratio tests, which confirmed the correctness of the restrictions.

<sup>7</sup> For IV estimation in nonlinear models see for example Roodman (2011).

**Table 2**  
Empirical Transitions of Stockmarket Index Levels - Portugal's PSI 20.

Origin	Destination					Total
	Class 1	Class 2	Class 3	Class 4	Class 5	
Class 1	24	5	0	0	0	29
Class 2	5	75	6	0	0	86
Class 3	0	6	32	8	0	47
Class 4	0	0	9	66	5	79
Class 5	0	0	0	5	27	32
Total	29	86	46	80	32	273

Note: Transitions between the class of origin of the index in a given week (rows) and the respective class on the following week (columns). Classes are defined dividing the observed range of the index into 5 equally spaced intervals. Data between 01/01/2009 and 28/03/2014.

As an overall picture, looking at Table 5, the main empirical findings are that the previous level of the index is always significant, similar to the ratings (not significant only in Indonesia). The GDP is significant only for Portugal and Greece, while the interest rate only affects Spain, the earlier period of Greece, and Indonesia.

The fact that the lagged level of the index is significant in all countries is in line with a common hypothesis in the economic literature, that is to assume that stock indices can be well described by a random walk - an autoregressive path whose probability of going up or down only depends on the last value taken. The original series had shown the same pattern, having been found non-stationary and AR(1) in all countries.

The variable GDP, when significant, shows a positive sign, as expected: a higher value of GDP should contribute to a stronger performance of the stock exchange, and contribute to increase the value of the index.

For the European indices considered here, only the Spanish IBEX and the Greek ASE in the first period are significantly influenced by the EURIBOR interest rate, showing a negative effect. A similar behaviour is also observed in Indonesia, where the Interbank Rate has a smaller, but still significant, negative effect on the JCI index. This result is expected. With the increase of interest rates, borrowing for consumers and enterprises becomes more expensive than before and existing mortgages and credits become more burdensome. The decrease in consumption and the expectations of a decrease in future consumption directly affect the price of stocks in the present, materializing in a decrease of the stock indices.

In the analysis of the variable of focus of this research, the S&P sovereign credit rating, we found that it is statistically significant for all three European countries considered - Portugal, Spain and Greece in both periods - while the same does not happen in the case of the Asian countries, where Indonesia fails to show an effect of the ratings on the dynamics of the stock exchange indices. The European crisis and the Asian crisis differed in many aspects: opportunity to devalue its

**Table 3**  
Descriptive statistics and specification tests.

		Europe (02/01/2009–28/03/2014)				Asia (03/10/1997–26/12/2003)	
		Portugal	Spain	Greece		South Korea	Indonesia
				I	II		
<b>Original series</b>							
Index	PSI20	IBEX	ASE	ASE	KOSPI	JCI	
N	274	274	120	154	326	326	
Range	min	4453	6065	1391	485	298	
	max	8839	12163	2838	1435	1028	
ADF test (p-value)							
Unit root	0.808	0.449	0.301	0.291	0.685	0.530	
Trend = 0	0.505	0.270	0.020	0.002	0.437	0.357	
Sign of trend			–	+			
AR(1) vs AR(2) (p-value)							
Lag 1 = 0	0.000	0.000	0.000	0.000	0.000	0.000	
Lag 2 = 0	0.914	0.492	0.952	0.182	0.085	0.712	
<b>Classified series</b>							
Trans. (%)	–1	9.16	9.52	10.93	9.15	8.28	
	0	82.05	81.32	78.99	81.70	83.13	
	+1	8.79	9.16	10.08	9.15	8.59	
<b>Ratings</b>							
$\rho = \text{Corr}(\text{Rating, Year})$ (%)	–90.97	–92.15	–84.35	55.93	65.24	–31.12	

Notes - Top: Stock market indices original series. Augmented Dicky-Fuller (ADF) test for stationarity and trend; autoregressive structure test: AR(1) vs AR (2). Middle: Classified series - 5 equally spaced intervals (levels) within observed range. Transitions (%): –1 = index dropped to immediately lower level; 0 = index remained in same level; +1 = index jumped up to next level. Transitions to more distant levels were not observed. Bottom: Ratings - Correlation of

**Table 4**  
Estimation results of Ordered Probit model.

	Europe			Asia		
	Portugal	Spain	Greece	South Korea	Indonesia	
<b>Coefficients</b>			<b>I</b>	<b>II</b>		
Index( $t - 1$ )	2.3121***	2.2844***	1.6576***	1.7717***	2.5070***	2.3993***
Ratings	0.2624***	0.3562***	0.4123*	0.4594***	0.1361**	-0.0651
GDP	0.0893**	0.0222*	0.0202	0.2140***	-0.0005	-0.0015
EURIBOR 6 M	0.0307	-1.6069***	-1.0322**	-0.8673		
IBR 3 M					-0.0176	-0.0214**
<b>Cut points</b>						
$\hat{\alpha}_1$	31.8571	40.9749	13.1845	63.4359	4.5250	0.3520
$\hat{\alpha}_2$	34.8801	43.7223	15.3838	66.8134	7.5662	3.8027
$\hat{\alpha}_3$	36.9850	46.2010	17.6941	69.4382	10.2045	6.3994
$\hat{\alpha}_4$	40.0954	49.2616	19.6271	71.3729	12.7533	8.7311
<b>Pseudo R<sup>2</sup></b>	0.6438	0.6316	0.6393	0.6170	0.6430	0.6121
<b>Observations</b>	273	273	119	154	325	312
<b>Test for time effects</b>						
$\chi^2$ (p-value)	0.5446	0.3694	0.9998	0.0844	0.5192	0.0582
<b>Test for endogeneity of Ratings</b>						
$\rho = 0$ (p-value)	0.5210	0.6219	0.5319	0.4390	0.7587	0.8480
<b>Test for exclusion restrictions</b>						
F (p-value)	0.8946		0.5616	0.1190	0.7729	0.0583
<b>Test for parallel lines assumption (PLA)</b>						
LR (p-value)	0.7074	0.1938	0.3662	0.0001	0.0075	0.9681
Variable(s) violating PLA				Ratings	Ratings	

Notes - Dependent variable: level of stock market index, expressed in five equally spaced classes. Regressors: previous level of stock market index; sovereign credit ratings; GDP; interest rates: EURIBOR 6-month rate (common rate for all European countries) or country-specific Interbank 3-month rates (for Asian countries). Specification tests for Ordered Probit model: test for significance of time effects; test for endogeneity of Ratings; test for joint exclusion restrictions of all non-significant regressors ( $p > 0.1$ ); test for parallel lines assumption in the reduced model. Greece: Period I - until 31/03/2011; Period II - from 01/04/2011. (\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ ).

own currency in the Asian countries case, as opposed to EU countries where the European common monetary policy and economic stability pact imposed serious constraints to the action of countries; consequent different approaches by governments when addressing the national crisis and different international conditions in the 1990's and in the 2000's. One additional explanation rely on the possibility for large domestic companies to have access to the international capital markets, and thus being less exposed to the rating changes of the home country. [Stulz \(1995\)](#) argues that globalization of firms reduces the cost of capital. [Kaminsky and Schmukler \(2002\)](#) find that the effect of upgrades and downgrades is present also for emerging markets, but it appears not to be large in economic terms.

When a sovereign downgrade happens, rating agencies used to apply the "sovereign ceiling policy": the credit rating of the firm is set at the level or below the sovereign rating of the domicile country. This practice has been recently revised, and for S&P firms can have a rating up to four notches above the sovereign rating.<sup>8</sup>

For all remaining countries, the effects of the Ratings are highly significant; in the case of level-varying effects, we observe for Greece a much larger value for the last threshold, i.e. the ratings influence even more the capacity of the index to reach the top class.

The reduced GOP model results have been used to compute transition matrices for each country, showing the probabilities of upward and downward movements of the stock indices levels. The transition probabilities for all countries are shown in [Tables 6–8](#), evaluated at the mean values for the regressors.

The results for Portugal and Spain are very similar, showing a large persistence, with a probability of remaining at the same index level taking values roughly between 70% and 85%. The Asian countries show a similar slow dynamics at the mean value of the regressors. In the case of Greece, however, the resilience of the index seems to be weaker, showing very high probabilities of falling to lower levels whenever the index is in the top classes. Rather than persisting at the top levels,

<sup>8</sup> A detailed analysis of the sovereign ceiling policy and the effect of changes in sovereign rating is presented by [Almeida et al. \(2017\)](#).

**Table 5**  
Estimation results of Generalised Ordered Probit model.

	Europe			Asia		
	Portugal	Spain	Greece	South Korea	Indonesia	
<b>Coefficients</b>			<b>I</b>	<b>II</b>		
Index( $t - 1$ )	2.3157***	2.2844***	1.6593***	1.9807***	2.4733***	2.3777***
GDP	0.0851***	0.0222*		0.1804***		
EURIBOR 6 M		-1.6069***	-1.1562***			
IBR 3 M						-0.0172**
Ratings	0.2573***	0.3562***	0.5394***			
$\beta_{R,1}$					0.2824***	
$\beta_{R,2}$				0.7885***	0.1966**	
$\beta_{R,3}$				0.8255***		
$\beta_{R,4}$				1.2941***		
<b>Cut points</b>						
$\hat{\alpha}_1$	30.5875	40.9749	8.0538	53.8497	6.7612	2.3610
$\hat{\alpha}_2$	33.6077	43.7223	10.2388	59.9189	8.8575	5.7441
$\hat{\alpha}_3$	35.7113	46.2010	12.5421	63.0466	8.7551	8.2575
$\hat{\alpha}_4$	38.8213	49.2616	14.4835	67.8813	11.3427	10.6430
<b>Pseudo R<sup>2</sup></b>	0.6438	0.6316	0.6381	0.6701	0.6511	0.6063
<b>Observations</b>	273	273	119	154	325	312

Notes - Reduced model specification based only on significant regressors. Dependent variable: level of stock market index, expressed in five equally spaced classes. Regressors: previous level of stock market index; sovereign credit ratings; GDP; interest rates: EURIBOR 6-month rate (common rate for all European countries) or country-specific Interbank 3-month rates (for Asian countries). Specification of Generalised Ordered Probit model includes non-parallel effects for Ratings in Greece Period II and South Korea. Models for all other countries/periods specified according to parallel lines assumption, i.e. standard Ordered Probit model. Greece: Period I - until 31/03/2011; Period II - from 01/04/2011. (\*\*\*)  $p < 0.01$ , (\*\*)  $p < 0.05$ , (\*)  $p < 0.1$ .

**Table 6**  
Transition Probabilities at Means for Iberia.

	Class 1	Class 2	Class 3	Class 4	Class 5
	Portugal				
Class 1	0.6983***	0.3015***	0.0002	0.0000	0.0000
Class 2	0.0362***	0.8533***	0.1100***	0.0004	0.0000
Class 3	0.0000	0.1375***	0.7067***	0.1558***	0.0000
Class 4	0.0000	0.0003	0.0958***	0.8684***	0.0354**
Class 5	0.0000	0.0000	0.0001	0.3051***	0.6947***
	Spain				
Class 1	0.6674***	0.3318***	0.0007	0.0000	0.0000
Class 2	0.0320**	0.7828***	0.1848***	0.0004	0.0000
Class 3	0.0000	0.0825***	0.7797***	0.1378***	0.0000
Class 4	0.0000	0.0001	0.1161***	0.8528***	0.0310**
Class 5	0.0000	0.0000	0.0003	0.3377***	0.6621***

Notes - Probabilities at the mean value of the variables considered in the reduced model, for Portugal and Spain between 01/01/2009 and 28/03/2014. (\*\*\*)  $p < 0.01$ , (\*\*)  $p < 0.05$ , (\*)  $p < 0.1$ .

the index shows probabilities between about 50% and 75% of falling back one class, and almost zero chances to improve the level.

To further explore the effects of the Ratings, the transition matrices are also evaluated at different percentiles of the variable, while keeping the GDP and Interest Rates at their mean values. For the sake of the example, Table 9 shows the probabilities evaluated at the 10th and the 90th percentiles of Ratings, respectively, for Portugal. For this country, the 10th percentile is the credit rating 10 (BB), and the 90th percentile is 17 (A+). The comparison reveals interesting results: for example, all other things equal, the probability of moving upward from Class 1 to Class 2 when in 10th percentile of Credit Rating is 7.52%, rising to 63.81% when in the 90th percentile. This is a very large difference in the probabilities of moving upwards in the stock market classes, being consistent with the expectation that Credit Ratings have a large impact in the stock indices. Likewise, downward movements show similar (symmetrical) results; for example, being in level 3 when the Ratings are lower implies a large probability of falling back to level 2 (43.06%) of the index, while if the Ratings are higher such probability shrinks to 2.41%.

A summary representation of the changes in the transition probabilities when a country moves from a lower to a higher sovereign rating is given in Fig. 4 for all considered countries (except Indonesia, where the variable was not significant). The

**Table 7**  
Transition Probabilities at Means for Greece.

	Class 1	Class 2	Class 3	Class 4	Class 5
Greece I					
Class 1	0.7888***	0.2098***	0.0014	0.0000	0.0000
Class 2	0.1957***	0.7122***	0.0920*	0.0001	0.0000
Class 3	0.0059	0.3642***	0.6056***	0.0243	0.0000
Class 4	0.0000	0.0232	0.5994***	0.3652**	0.0121
Class 5	0.0000	0.0001	0.0889	0.6349***	0.2761*
Greece II					
Class 1	0.6396***	0.3589***	0.0015	0.0000	0.0000
Class 2	0.0523*	0.7850***	0.1627**	0.0000	0.0000
Class 3	0.0002	0.1591**	0.8134***	0.0274	0.0000
Class 4	0.0000	0.0014	0.4748***	0.5202***	0.0036
Class 5	0.0000	0.0000	0.0207	0.7405***	0.2388

Notes - Probabilities at the mean value of the variables considered in the reduced model. Period I (top panel): between 01/01/2009 and 31/03/2011; Period II (bottom panel): between 01/04/2011 and 28/03/2014. (\*\*\*)  $p < 0.01$ , (\*\*)  $p < 0.05$ , (\*)  $p < 0.1$ .

**Table 8**  
Transition Probabilities at Means for Asian countries.

	Class 1	Class 2	Class 3	Class 4	Class 5
South Korea					
Class 1	0.7188***	0.2811***	0.0001	0.0000	0.0000
Class 2	0.0291***	0.8790***	0.0918***	0.0001	0.0000
Class 3	0.0000	0.1263***	0.7828***	0.0909***	0.0000
Class 4	0.0000	0.0001	0.1273***	0.7989***	0.0736***
Class 5	0.0000	0.0000	0.0002	0.1528***	0.8471***
Indonesia					
Class 1	0.6273***	0.3726***	0.0001	0.0000	0.0000
Class 2	0.0200*	0.8882***	0.0917***	0.0001	0.0000
Class 3	0.0000	0.1474***	0.7812***	0.0713***	0.0001
Class 4	0.0000	0.0003	0.1806***	0.7488***	0.0703**
Class 5	0.0000	0.0000	0.0005	0.1824***	0.8171***

Notes - Probabilities at the mean value of the variables considered in the reduced model, for South Korea and Indonesia between 01/10/1997 and 26/12/2003. (\*\*\*)  $p < 0.01$ , (\*\*)  $p < 0.05$ , (\*)  $p < 0.1$ .

**Table 9**  
Transition Probabilities for varying rating in Portugal.

	Class 1	Class 2	Class 3	Class 4	Class 5
Portugal (Rating = BB)					
Class 1	0.9248***	0.0752*	0.0000	0.0000	0.0000
Class 2	0.1901**	0.7939***	0.0161	0.0000	0.0000
Class 3	0.0007	0.4306***	0.5420***	0.0268	0.0000
Class 4	0.0000	0.0064	0.3436***	0.6468***	0.0032
Class 5	0.0000	0.0000	0.0035	0.6554***	0.3412**
Portugal (Rating = A+)					
Class 1	0.3579**	0.6381***	0.0040	0.0000	0.0000
Class 2	0.0037	0.6296***	0.3595***	0.0073	0.0000
Class 3	0.0000	0.0241	0.5269***	0.4483***	0.0006
Class 4	0.0000	0.0000	0.0143	0.8076***	0.1780**
Class 5	0.0000	0.0000	0.0000	0.0818**	0.9182***

Notes - Probabilities at 10th (top panel) and 90th (bottom panel) percentile of Credit Ratings. All other variables are evaluated at the mean value. In this case for Portugal, the 10th percentile corresponds to the value BB, while the 90th percentile corresponds to the value A+. (\*\*\*)  $p < 0.01$ , (\*\*)  $p < 0.05$ , (\*)  $p < 0.1$ .

graphs display the probabilities that the stock market index shows an increase or a decrease in the following period, for specific values of the Ratings. The probabilities are computed at a selection of percentiles of the rating distribution for each country (reported in Table 10), all remaining regressors being evaluated at their mean.

The results are rather striking, showing that when the countries are rated higher the respective stock market index has a much higher probability of increasing, reducing at the same time the chances to fall to lower levels. For the second period of Greece, only when the B-/B threshold is passed we observe some changes in the resilience of the index; below such level the

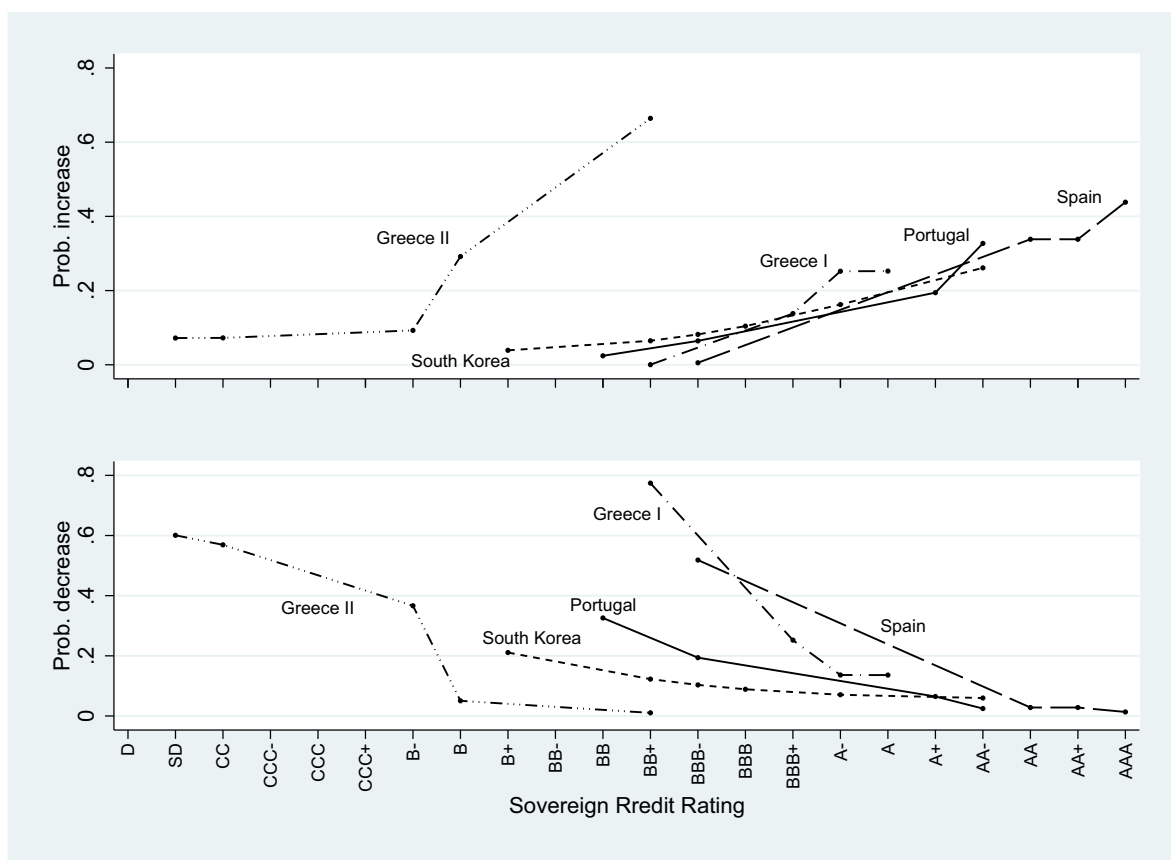


Fig. 4. Probabilities of changes in the index level as a function of observed values of ratings for each country in the respective period of observation. Top: probability of an increase in the index level; bottom: probability of a decrease.

Table 10  
Selected percentiles of Credit Ratings.

	p1	p10	p25	p50	p75	p90	p99
Portugal	BB	BB	BBB-	BBB-	A+	A+	AA-
Spain	BBB-	BBB-	BBB-	AA	AA+	AA+	AAA
Greece I	BB-	BB+	BB+	BBB+	A-	A-	A
Greece II	SD	CC	CC	CCC	B-	B-	BB-
South Korea	B+	BB+	BBB-	BBB	A-	A-	AA-
Indonesia	SD	SD	CCC+	CCC+	B-	B	BBB-

Notes - Credit Ratings in each country during the respective period of observation: from 01/10/1997 to 26/12/2003 for Asian countries, from 01/01/2009 to 28/03/2014 for European countries. Greece: Period I - from 01/01/2009 to 31/03/2011; Period II - from 01/04/2011 to 28/03/2014.

probability of improving is extremely low, while the index falls to lower levels with probabilities between 40% and 60%; however, soon after the threshold, the probability of improvement of the index increases steeply, and the chances of falling back are reduced to close to zero.

### 5. Conclusions

In this research we examine the impact of sovereign debt rating changes on security prices in stock exchanges in five countries during times of financial crisis, proposing a new empirical approach based on Markov Chains. Using a generalized ordered probit model (GOP), we compute the transition probabilities for each stock index to move up or down, or remain at the same level, controlling for several macroeconomic factors. Our results show not only that the level of the rating per se matters, but most importantly that a weaker economic condition of the country affects the probability of the stock markets to achieve better performance. In the case of Greece, we identify a significant threshold at the B- rating class. At this level, the

probabilities that the stock market index increases are almost zero, but the chances of a fall are remarkably between 50% and 75%. The likelihood of the country to default and the so-called sovereign ceiling, strongly affect the expectations of the market participants regarding the future performance of the stocks market as a whole. Even if the rating agencies started recently to amend to this rule and, if so, apply a rating above the sovereign level, the effect on the economics of the single country appears to be strong, especially for low-rated countries.

This work is relevant due to the importance of credit rating agencies as a risk assessment provider for portfolio managers, for bankers and regulators (in the context of Basel Committee recommendations and ECB supervision) and also for governments and policymakers. In addition, the prominence of institutional investors in capital markets, the portfolio composition restrictions on securities that have an investment-grade rating, and the impact of stock-market behavior on the cost of funding, both for companies and governments, explain why the effect of sovereign debt ratings on stock markets is particularly important.

The sample used in this study includes five countries with different economic, political, cultural and legal characteristics and two different periods of time, which allows us to test the regional/institutional differences in the dynamics of the crisis. But this diversity recommends that comparisons must be done carefully and further research could be done.

From the policy perspective, our analysis shows that the impact of a credit rating downgrade for less resilient countries might exacerbate the crisis, reinforcing a feedback loop transmission between sovereign and financial markets. An independent assessment of the creditworthiness of the single companies, higher participation of the companies in international activities and supranational financial markets might reduce this strong link and contribute to a quicker recovery.

**Table 11**

Results of Instrumental Variables estimation to test for endogeneity of Ratings. European countries.

Response	Portugal		Spain	
	Index level	Rating	Index level	Rating
<b>Regressors</b>				
Index( $t - 1$ )	2.3121***	0.3841***	2.2844***	0.4961***
Ratings	0.2624*		0.3562**	
GDP	0.0893	-0.1927***	0.0222	-0.0554***
EURIBOR 6 M	0.0307	-1.4451***	-1.6069**	2.9373***
Crisis phase (Year)		-1.0897***		-0.7797***
Constant		72.2330***		99.1424***
<b>Cut points</b>				
$\hat{\alpha}_1$	31.8571		40.9749	
$\hat{\alpha}_2$	34.8801		43.7223	
$\hat{\alpha}_3$	36.9850		46.2010	
$\hat{\alpha}_4$	40.0954		49.2616	
Test for $\rho = 0$ (p-value)		0.5210		0.6219
$\hat{\rho}$		-0.0784		-0.0602
<b>Observations</b>	273	273	273	273
Response	Greece I		Greece II	
	Index level	Rating	Index level	Rating
<b>Regressors</b>				
Index( $t - 1$ )	1.7994***	0.4512***	1.6501***	0.0457
Ratings	-0.2142		0.7051**	
GDP	0.0770	0.0975***	0.1282	0.2986***
EURIBOR 6 M	0.0294	1.5130***	0.2177	-3.9216***
Crisis phase (Quarter)		-0.1534***		
Crisis phase (Semester)				-0.6420***
Constant		-21.4638***		-76.6990***
<b>Cut points</b>				
$\hat{\alpha}_1$	25.3278		40.8033	
$\hat{\alpha}_2$	27.3481		43.9903	
$\hat{\alpha}_3$	29.5218		46.4630	
$\hat{\alpha}_4$	31.3754		48.3230	
Test for $\rho = 0$ (p-value)		0.5319		0.4390
$\hat{\rho}$		0.3646		-0.3131
<b>Observations</b>	119	119	154	154

Notes - Main equation: ordered probit with robust standard errors. Reduced form equation: linear regression with robust standard errors. The instrument is a time effect representing the crisis phase: *Year* for Portugal and Spain; *Quarter* for Greece period I (until 31/03/2011); *Semester* for Greece period II (from 01/04/2011). Coefficient  $\rho$  represents correlation between errors of main equation and reduced form equation. Test of endogeneity performed through test for  $\rho = 0$ . (\*\*\*)  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ ).

**Table 12**  
Results of Instrumental Variables estimation to test for endogeneity of Ratings. Asian countries.

Response	South Korea		Indonesia	
	Index level	Rating	Index level	Rating
<b>Regressors</b>				
Index( $t - 1$ )	2.5070***	-0.0432	2.3993***	-0.3340***
Ratings	0.1361		-0.0651	
GDP	-0.0005	0.0253***	-0.0015	0.0285***
IBR 3 M	-0.0176	-0.0270	-0.0214**	-0.0470***
Crisis phase (Year)		-1.0918***		-1.4176***
Constant		-3.6573		-17.6579***
<b>Cut points</b>				
$\hat{\alpha}_1$	4.5250		0.3520	
$\hat{\alpha}_2$	7.5662		3.8027	
$\hat{\alpha}_3$	10.2045		6.3994	
$\hat{\alpha}_4$	12.7533		8.7311	
Test for $\rho = 0$ (p-value)		0.7587		0.8480
$\hat{\rho}$		-0.0549		0.0311
<b>Observations</b>	312	312	312	312

Notes - Main equation: ordered probit with robust standard errors. Reduced form equation: linear regression with robust standard errors. The instrument is a time effect representing the crisis phase (*Year from crisis onset*). Coefficient  $\rho$  represents correlation between errors of main equation and reduced form equation. Test of endogeneity performed through test for  $\rho = 0$ . (\*\*\*)  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

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## Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

## Disclaimer

The views expressed are purely those of the authors and may not in any circumstances be regarded as stating an official position of the European Commission.

## Appendix A

See Tables 11 and 12.

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