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SUMMARY

The thesis presents three essays on credit risk modelling, which respond to the constant needs of financial institutions and their supervisors to update models on credit risk according to last macro-financial events and regulatory innovations.

Subject of the first and second essay is the credit risk of sovereign entities in Eurozone. Subject of the third essay is the credit risk of loans owned by European banks.

The first essay (Giorgione et al., 2016) investigates the price discovery process of credit risk of Eurozone countries from 2008 to 2014 within the two interconnected Bond and Credit Default Swap markets.

Subject of the analysis are the short- and long-term dynamic relationships between the two time-series of credit spreads obtained from the Bond and Credit Default Swap markets. Through a Vector Error Correction (VEC) model, the analysis allows determining which credit spread anticipates the other in pricing a certain Sovereign entity, because it is more efficient and timely to incorporate new credit information.

The empirical evidences obtained in periods characterized by different levels of credit risk, and then associated to different volatility and liquidity conditions, suggest implications for practitioners who want to take advantage of potential market inefficiencies and for policymakers interested in maintaining the stability of financial systems through timely and correct responses to possible price shocks.

From these evidences, arises that limited attention has been devoted in literature to use price discovery models for policy analysis. In particular, for analysing the effects of monetary policies, of bailout policies and of regulatory innovations on the microstructure of financial markets as well as in measuring the effectiveness of policies in reducing the financial instability.

Then, the second essay provides an econometric model able for investigating these economic interests. The model adjusts current price discovery measures, so far mostly used in finance theory for studying the microstructure dynamics of financial markets, in a way that are suitable for policy analysis.

Leveraging on the VEC approach, the model extends time-invariant price discovery measures to time-varying measures, this means to leave the model free to estimate when structural changes happen in the price discovery dynamics.

The empirical application of the model, on a topic currently investigated in literature “the sovereign credit risk in Eurozone”, shows the impact of regulation, of monetary policies and of crisis events on price discovery, such that policy implications can be derived.

Such a model could contribute on the literature of both financial market microstructure and policy impact analysis.

The topic of the third essay is the Stress Test exercise requested by the European Central Bank to the majority of European Banks. In the contest of credit risk stress testing, the essay presents a methodological approach, compliant with regulatory requirements, to build a satellite model for stress testing the probability of default of a loan portfolio.

The approach models the default rates on a loan portfolio as the sum of a systematic risk factor and idiosyncratic risk factor. The systematic risk depends on the economic conditions of a country. The unobservable idiosyncratic risk is assumed to depend on the strategic decisions of bank managers as well as on the specific features of the borrowers.

The essay shows empirical projections of the baseline and stressed probability of default on corporate, small business and retail loan portfolios of the Italian banking system, through the simulation of European Banking Authority macroeconomic scenarios.

The model has implications for supervisors and credit risk managers of European banks. For supervisors interested in a common methodology to compare the effects of an economic scenario on different institutions. For credit risk managers interested in an internal model, compliant with ECB stress testing framework, to obtain forward-looking probability of default. At the same time, this satellite model can be used for other regulatory exercises, such as the ICAAP report and “the expected credit losses impairment” modelling under the new accounting standard IFRS 9. Moreover, the model can be adopted not only for risk management activities, but also for budget and planning operations.

A Model to Test the Price Discovery of Sovereign Credit Risk in the Eurozone¹

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ABSTRACT

This paper investigates the price discovery process of credit risk of Eurozone countries from 2008 to 2014 within the two interconnected Bond and Credit Default Swap markets.

Subject of the analysis are the short- and long-term dynamic relationships between the two time-series of credit spreads obtained from the Bond and Credit Default Swap markets. The analysis allows determining which credit spread anticipates the other in pricing a certain Sovereign entity, because it is more efficient and timely to incorporate new credit information.

The empirical evidences obtained in periods characterized by different levels of credit risk, and then associated to different volatility and liquidity conditions, suggest implications for practitioners who want to take advantage of potential market inefficiencies and for policymakers interested in maintaining the stability of financial systems through timely and correct responses to possible price shocks.

JEL classification: G01, G12, G14, G20

Keywords: credit risk, Eurozone countries, price discovery, CDS, bond spreads

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1. INTRODUCTION AND OBJECTIVES

The object of this work is to study the relationships that are established over time between the market of credit default swaps (CDS) and that of bonds. The analysis of these relationships allows us to investigate the dynamic effects of the process by which the pricing of credit risk is determined in the Eurozone countries. This risk typology is, in fact, incorporated in the credit spreads, or rather in a series of indicators suited to quantify the return to an investor for assuming the credit risk that is intrinsic in a bond.

The most commonly used indicators are

- the bond spread (BS), or the difference between the bond return and the risk-free rate (AAA rating) or between the bond return and the interest rate swap (IRS) rate, which synthesises the conditions of the interbank market (AA rating);
- the CDS spread (CDSS), whose quotation depends on the probability of default of the entity underlying the CDS contract.

Under normal conditions, whether in the long or the short term, the prices of the BS and CDSS for the same period and referencing entity are sufficiently related and thus leave no room for arbitrage opportunities. The same is not true, especially in the short term, under conditions of financial instability such as those that have been prevalent in the market in the last few years. This observation, in the context of the sovereign market of the Eurozone, offers us a way to verify whether, in the adjustment process between the various spreads, there emerges a *leading* market and possible *following* markets. Should this be the case, the practical implications are obvious. Knowledge of the mechanism and timing of any adjustment process between the markets should, in fact, be particularly interesting for market operators, policy makers and academics.

For operators, the phenomenon of mean reversion over the long term offers the opportunity for arbitrages or margins for making profits taking on zero or low risks.

For policy makers, a deeper familiarity with the dynamics by which the markets evolve creates a starting point for planning timely and calibrated monetary policies.

For academics, analysing the dynamics between the prices of a given asset traded in several different markets is fundamental to perfecting the study of the microstructure and efficiency of the financial markets.

A critical step to fully understand the objectives of this work is to understand that, if the “asset” credit risk were traded in a single market, then the whole process of price discovery would be concentrated in that market. In reality, the asset in question is traded in more than one market, CDS and bond, and the complex flow of information is fragmented. From this, the importance emerges of determining which market has the greater informative power, or which market succeeds in incorporating new information into its own price first.

To this end, the work sets out two intermediate objectives and one final objective. The intermediate objectives aim to test whether:

1. the action of the market forces reduces or annuls the divergences observed in the short term between the CDSSs and BSs, such that those spreads converge towards an equilibrium in the long term;
2. between the bond market and the CDS market it is possible to identify a market that is more efficient than the other, in terms of speed with which it absorbs new credit risk information into its prices.

Verifying the intermediate objectives constitutes the methodological premise for realising the final objective of this study, that is, adequately representing the changes in the actual dynamics of price discovery in the following three periods:

1. from the second half of September 2008 to the first half of September 2010, approximately equivalent to the period of greatest intensity of the financial crisis;
2. from the second half of September 2010 to the first half of September 2012, approximately equivalent to the period of greatest intensity of the sovereign debt crisis;
3. from the second half of September 2012 to the first half of September 2014, approximately equivalent to the period marked by a trend of progressive downsizing of credit risk in the Eurozone².

² The second period of analysis finishes with the start of the program of Outright Monetary Transactions (OMT) that has successfully limited the credit spreads of the Eurozone countries.

The final objective will be reached by verifying the intermediate objectives for each Eurozone country and for each of the three abovementioned time periods. The results will be synthesised by period, both at the country level and the aggregate level, distinguished into the group of Core countries (Austria, Belgium, Finland, France and Holland) and the group of Peripheral countries (Greece, Ireland, Italy, Portugal and Spain) of the Eurozone.

The paper is organised as follows. Subsection 1.1, discusses in detail the related literature. Section 2, describes the dataset. Section 3, introduces the econometric methodology to be used and develops in such a way that the reader will have at their disposal a conceptual map that will facilitate comprehension of the empirical results, which are reported in Section 4. In that section the results are presented in three tables that refer to the three time periods examined. With the aim to offering greater consistency for the reader, the results in each table are configured in such a way as to show the succession of the methodological phases, distinguishing between the Core and Peripheral countries. Section 5, summarises and interprets the results of this work. Subsection 5.1, concludes the paper with practical implications of the findings.

2. REVIEW OF THE LITERATURE

This work positions itself within the more generic thread of study that examines the relationship of theoretical parity between CDSS and BS, considered inviolable prior to the crisis. Specifically, it relates to a series of scientific works that empirically investigate the propositions of Duffie (1999). That author was the first to argue that, theoretically, there is a perfect correspondence, or no probability of arbitrage, between a risky obligation, a risk-free security and a CDS contract with equivalent maturity on a notional of equal nominal value. In truth, a portfolio composed of a risky obligation and a CDS on the same security should replicate synthetically a position on a risk-free security. The difference between the BS and the CDS spread is referred to, both by market operators and academics, as the CDS-Bond Basis (or Basis). Its theoretical value is equal to zero, but in practice, even under normal market conditions, because of multiple technical factors and other variables, the basis takes positive values, albeit close to zero (Choudhry 2006). In phases of market turbulence, the basis can take values that are negative or even highly positive. These configurations of the basis offer,

therefore, opportunities to apply arbitrage strategies, or operations that are theoretically risk free but produce positive returns and do not absorb capital (Amadei et al, 2011)³.

These considerations have driven academics to investigate the link, at times contradictory, between the theoretical efficiency of the microstructure of the markets and the above described evidence. The key research interests can be divided into two threads.

The first thread has aimed at determining the drivers of the Basis that cause the theory of the absence of arbitrage opportunities between BSs and CDSSs to be violated. Bai & Dufresne (2013) and Fontana (2011) provide evidence that, during periods of financial turbulence, deviations from the parity of the Basis do not allow for profitable arbitrage. This happens because of multiple factors that can be attributable to a rise in the cost of liquidity and in the return, to frictions implicit in the functioning of the markets, or the availability of counterparties and quality standards for the necessary guarantees.

The other thread of research, within which the current work sits, is directed towards investigating the dynamic relations established over time between the bond and CDS markets. More precisely, the focus of this second thread is the analysis of the *price discovery* of credit risk, defined by Lehman (2002) as “the efficient and timely incorporation of the information implicit in investor trading into market prices”.

³ From among the first authors who dealt with the topic of arbitrage limits, we cite Shleifer and Vishny (1997). In principle, and from the theoretical viewpoint, arbitrage does not require capital and is not risky. In reality, however, the majority of arbitrage opportunities do require capital and are risky. Along these lines, Amadei et al (2011) hold that there exists counterparty risk tied to the use of CDS that prevents the arbitrage from being perfectly risk free. More generally, it has been observed that arbitrage strategies are not always attractive to intermediaries or professional market operators. Especially in extreme market conditions (such as those examined in this work), it could be difficult to find resources and allocate them in a timely manner to the most opportune strategies, with inevitable consequences for the economic effectiveness of the strategies themselves. Because of possible frictions in the functioning of markets, there is, moreover, the risk of having to close out positions before their natural maturity and, therefore, of incurring penalties on the implemented strategies. Li, Zhang and Kim (2013), following on from Shleifer and Vishny (1997), reach similar conclusions about the limits of arbitrage. In their 2013 work, they argue that a non-zero Basis, although from a theoretical point of view represents the operating assumption for setting up an arbitrage, in fact creates a wide range of risks. In the first place, a non-zero Basis could have been caused by differences in contract types between bonds and CDSs and might not necessarily represent the theoretical premise for realizing an arbitrage strategy. Secondly, arbitrageurs could lose even in potentially profitable exchanges. They could, for example, incur liquidity risks relating to both the bond market and the CDS market, just as they could incur deleveraging risks from other investors. Therefore Li, Zhang & Kim (2013) argue that a non-zero Basis in fact exposes operators and arbitrageurs to risky activity. In the European sovereign context, the main cause of obstacles to the realization of arbitrage has been the difficulty of short selling government bonds, in turn due to the elevated costs and the difficulty of finding bonds to borrow.

The fact that the price discovery process for credit risk does not occur within a single market has led academics to pose two important research questions:

1. Does the divergence in the short term between CDSS and BS present a return to the equilibrium in the long term?
2. Between the bond market and the CDS market, does one anticipate the other in pricing credit risk? In other words, which is the leading market and which the lagging market?

In the face of these questions, the first contribution is the study of Blanco et al (2005) that analyses a sample of 33 US and European investment-grade companies. Blanco et al. (2005), for 26 of the 33 entities examined, show the presence of a stable cointegrating relationship, or the tendency for the deviations in the short term between the CDSSs and BSs to return to equilibrium in the long term. Other important works that investigate the corporate sector in the pre-crisis period are those of Zhu (2006) and Norden & Weber (2009). The two verify the presence of cointegration for 14 out of 24 and 36 out of 58 entities analysed respectively. These studies show that in the corporate sector the Basis is usually positive and narrow and that, in the majority of cases and predominantly for the US companies, the CDS market anticipates the bond market. The authors attribute this evidence to the greater liquidity of the CDS market and to the different types of intermediaries that operate in it .

Among the more recent contributions examining the impact of the financial crisis on the price discovery of credit risk, we find Giannikos et al (2013). In this study, the authors, using a sample of 10 US financial companies, investigate the short-term and long-term dynamic relationships between the stock, bond and CDS markets. In the quest to identify the possible existence of a leading market, this work offers an important contribution by showing how the role of the stock market changes with the start of the financial crisis. Specifically, the authors observe that before the financial crisis the stock market played a dominant role in price discovery. Then, with the beginning of the crisis, its informative power diminished and the CDSSs assumed the dominant role. Within this line of research, Avino et al (2013) provide a wider examination of price discovery in the corporate market, taking into account a sample of 30 non-financial European and US companies. The empirical evidence shows that the leading role was taken by both the CDS market and the stock market before the default of Lehman Brothers (2006-2007) and during the sovereign debt crisis (2009-2012). In contrast, during the period of the sub-prime crisis (2007-2009), and only in the European sample, the BSs took

leadership of the price discovery process. In contrast, for the US sample, during the sub-prime crisis, the CDSSs took on the dominant role, as also evidenced in Giannikos et al. (2013). From the evidence that can be inferred from the literature, it is not however possible to draw unequivocal conclusions because the process of price discovery among credit spreads changes significantly depending on the periods, geographical areas and specific reference entities examined.

The studies on price discovery are concentrated predominantly on the corporate sector. Little attention has been given to the sovereign sector. However, the recent European sovereign debt crisis has offered further opportunities to study the dynamics of price discovery in relation to European government entities, and possibly to compare the results to those obtained for the corporate sector.

Among the first to study the price discovery of credit risk in the Eurozone countries, we refer to Fontana and Sheicher (2010), who investigate the relationship between CDSSs and BSs in 10 European countries from 2006 to 2010. These authors show, before the financial crisis, a general absence of cointegration between the two credit spreads, as a result of the scarcity of trading in CDS contracts. Then, from the default of Lehman Brothers onwards, all of the countries prove to be cointegrated. This fact allows an investigation of the dynamics in the adjustment process of the two series of spreads towards equilibrium in the long term. For Germany, France, Holland, Austria and Belgium, the bond market guides the price discovery. In contrast, for Italy, Ireland, Spain, Portugal and Greece, the CDS market moves first.

The most recent contributions on the price discovery of sovereign credit risk are those from Coudert and Gex (2013) and Alper et al. (2013). Coudert and Gex (2013) analyse the links between CDSSs and BSs in order to determine which market is the leader in the price discovery process. The analysis is carried out on a sample made up of CDSSs and BSs for 18 government entities, for the period 2007-2010. The authors show significant differences between the Peripheral and Core countries of the euro area. Specifically, for the countries with high credit ratings, the bond market leads the CDS market. For the countries with low ratings, the CDS market leads the bond market. Moreover, the CDS prices see their informative role increase in the period with the highest perception of credit risk. Similar conclusions are reached by Alper et al (2013), using a sample of 10 advanced countries, limited to the period 2008-2010.

In any case, it is necessary to interpret the results with due caution, given that all the studies cited are based on rather short intervals and none take into consideration the period of instability identified by the sovereign debt crisis in Eurozone. While the current paper is based on a sufficiently long period of time. The chosen extension of the time period allows us to represent particularly accurately the actual dynamics of the price discovery of credit risk in market contexts characterised by different levels of volatility and liquidity.

3. DATASET

To carry out the present study, a dataset was extracted from the Bloomberg platform. The dataset includes, for the time period 2008-2014, daily quotations, from the close of trading, of the CDSSs and rates of return of government bonds with five-year maturities from 11 euro-area countries⁴.

For each country, the series of BS is calculated as the difference between the bond rate of return for that country and the rate of return on a five-year German government bond. The choice to use the German government bonds is justified for at least three reasons. First, for the interval considered, the rate of return on German government bonds was recognised by the market as a valid proxy for the risk-free rate. Second, with this approach, in contrast to that using interest rate swaps (IRS) as a proxy for the risk-free rate, one obtains bond spreads that are always positive⁵. The third reason is directly linked to the second. Having only positive bond spreads allows us to run our analysis on the logarithms of the two series of credit spreads.

⁴ In detail, the countries subjected to the analysis are Austria, Belgium, Finland, France, Greece, Ireland, Italy, Holland, Portugal, Slovakia and Spain. Germany is not included, as it is used as the reference country for the calculation of the BSs. The selection of the time period and countries is based on data availability. The choice of a five-year maturity is based on the greater liquidity at this maturity for the financial instruments considered. It is scarcely necessary to note that there is little interest in investigating the period preceding the financial crisis for two reasons: i) the CDSs and BSs for sovereign entities were particularly contained and stable at that time, since, for the majority of issuers, creditworthiness was perceived to be high; ii) the orders on government CDS showed trading volumes particularly contained.

⁵ This approach avoids incurring negative credit risk quotations, implicit in the bond market, for those government entities with high creditworthiness. Precisely, the IRS rate synthesizes the conditions of the interbank market and is typically associated with an AA rating. Using this rate as a proxy for the risk-free rate produces negative bond spreads for countries to which the market attributes creditworthiness superior to the interbank creditworthiness. Since, during the period of analysis, Germany benefited from an interest rate on debt that was lower than those for all other Eurozone countries, using this rate of return as a proxy for the risk-free rate allows us to obtain positive BSs for all of the countries under study.

The data sample is subdivided into three periods: (1) 2008-2010, (2) 2010-2012 and (3) 2012-2014.

4. ECONOMETRIC METHODOLOGIES

In order to verify the changes from one period to another of the country-dependent dynamics of the price discovery process in the Eurozone, a specific econometric model is developed, based on the methodology proposed by Gonzalo and Granger (1995) that is widely used throughout the literature on this subject. In detail, the price discovery analysis is divided into two phases. This subdivision is justified by the fact that the first and second phases are aimed at addressing, respectively, the two intermediate objectives of this work.

In the first phase, in order to verify whether the short-term deviations in the CDSSs and BSs converge, following the action of market forces, towards parity in the long term (first intermediate objective), the concept of “cointegration” is used. The verification of the presence of cointegration between the CDSS and BS series offers an adequate response to the first research question, since two series are said to be cointegrated if, despite deviating from each other in the short term, they exhibit a process of adjustment towards equilibrium in the long term. In other words, and in technical terms, when a linear combination of two historical non-stationary series generates a stationary process, one can say that the two series are cointegrated.

Initially, then, the hypothesis of non-stationarity (or unit root) of the CDSS and BS series is verified using the Augmented Dickey-Fuller (ADF) test, for each of the 10 Eurozone countries. Next, using Johansen’s (1988) trace statistic, the existence of a cointegrating relationship between the CDSSs and BSs, or of a stationary linear combination of the two non-stationary series, is verified. This econometric technique allows us to test the null hypothesis of the absence of cointegration between the CDSSs and BSs for each of the countries studied.

The results of this first step of the analysis serve as preparation for the second phase of the research⁶. In this phase, in order to check whether there exists a leading and a following market, or something in between, the lead-lag relationship between the BSs and CDSSs is analysed. In

⁶ See Section 1 (Introduction), 2nd intermediate objective.

other words, we analyse the adjustment process of the short-term deviation between the two series towards the long-term equilibrium that was confirmed by the test of cointegration (in the first phase). An econometric model that allows us to investigate this adjustment process is the Vector Error Correction Model (VECM) proposed by Engle and Granger (1987). The VECM is a vector autoregressive (VAR) model with an error correction term (ECT). The model is specified in this paper by the following equations:

$$\Delta CDS_t = \alpha_1 + \lambda_1 Z_{t-1} + \sum_{i=1}^p \beta_{1i} \Delta CDS_{t-i} + \sum_{i=1}^p \gamma_{1i} \Delta BOND_{t-i} + \varepsilon_{1t} \quad (1)$$

$$\Delta BOND_t = \alpha_2 + \lambda_2 Z_{t-1} + \sum_{i=1}^p \beta_{2i} \Delta CDS_{t-i} + \sum_{i=1}^p \gamma_{2i} \Delta BOND_{t-i} + \varepsilon_{2t} \quad (2)$$

Where:

- a) ΔCDS_t and $\Delta BOND_t$ are the CDSS and BS series expressed in first differences and represent, respectively, the dependent variables in equations 1 and 2;
- b) ΔCDS_{t-i} and $\Delta BOND_{t-i}$ are the lagged first differences of the CDSS and BS series; they are included, in equations 1 and 2, as model regressors that explain the deviations in the short term in the first differences of the CDSSs and BSs;
- c) p =the number of lags in the regressors;
- d) ε_{1t} and ε_{2t} are error terms;
- e) $Z_{t-1} = CDS_{t-1} - c - \beta_0 BOND_{t-1}$ (the ECT) describes the deviations of the CDSSs and BSs, lagged by one day, from their theoretical relationship of parity in the long term⁷.

⁷ In theory, the intercept and the beta should be equal to 0 and -1 respectively, but this condition is not imposed because, as noted earlier, equality between CDSSs and BSs, for equal maturities and reference entities, occurs only approximately, due to technical and market factors (Choudhry, 2006). In other words, if the quotations of the two credit spreads were exactly equal in the long term, they would be cointegrated with a vector (1, -1, 0), while in this study the vector (1,-b,c) is used, in line with the relevant literature. These two parameters are estimated in the first phase of the analysis using Johansen's method.

- f) λ_1 and λ_2 are termed “adjustment coefficients” in view of the fact that they correct the deviations in the CDS and bond spreads towards equilibrium in the long term, as verified by the cointegration analysis.

In practice, equations 1 and 2 express the deviations in the short term between the BSs and CDSSs in first differences, with the respective addition of “ $\lambda_1 Z_{t-1}$ ” and “ $\lambda_2 Z_{t-1}$ ” that capture the relationships between the two spreads in the long term.

For the purpose of our analysis, the significance and the signs of the adjustment coefficients (λ_1 and λ_2) allows us to verify the degree of informative power of one market with respect to the other. Specifically, we can consider three cases:

1. If λ_1 is negative and significant, the bond market incorporates new information more quickly than the CDS market, and consequently the latter moves in the direction of restoring some kind of equilibrium in the long run.
2. If λ_2 is positive and significant, the CDS market incorporates new information more quickly than the bond market, and consequently the latter moves in the direction of restoring some kind of equilibrium in the long run⁸.
3. When both coefficients are significant and have the correct sign (λ_1 negative and λ_2 positive), both markets participate in the adjustment process towards equilibrium in the long term.

⁸ A more detailed explanation of the theoretical interpretation of the adjustment coefficients (λ_1 and λ_2) is the following. If, in equation 1 the ECT of the equilibrium relationship in the long term predicts significant changes in the CDSSs, then that signifies that the BSs generally move first and the CDSSs conform to them afterwards. Precisely, if the ECT is positive (negative), it signifies that the CDSS is greater (smaller) in numerical terms than the BS and is outside of the long-term equilibrium verified by the cointegration test. In other words, it is higher (lower) than the theoretical value expected in equilibrium, which signifies that one would expect a readjustment downwards (upwards). In order to find such a dynamic empirically, λ_1 should take a negative sign. If, in equation 2 the ECT of the equilibrium relationship in the long term predicts significant changes in the BSs, then that signifies that the CDSSs generally move first and the BSs conform to them afterwards. Precisely, if the ECT is positive (negative), it signifies that the BS is smaller (greater), in numerical terms, than the CDSS and is outside of the long-term equilibrium verified by the cointegration test. In other words, it is lower (higher) than the theoretical value expected in equilibrium, which signifies that one would expect a readjustment upwards (downwards). In order to find such a dynamic empirically, λ_2 should take a positive sign. In other terms, the lower is the speed of adjustment (the smaller is λ) of a certain market (dependent variable), the greater is its contribution to price discovery. Ultimately, the more efficient is a market (dependent variable), the more quickly it incorporates new information (λ decreases) and the more informative power it has in the price discovery process.

In the first two cases, when only one of the coefficients is statistically different from zero (significant) and has the correct sign (negative for λ_1 and positive for λ_2), a single market contributes to the price discovery (*Market Share* = 100%). This is termed the leading or dominant market. In the third case, when both coefficients are significant, in order to determine how much each market contributes to the price discovery, the following formula of Gonzalo and Granger (1995) is applied:

$$\text{Market Share: } \frac{\lambda_2}{\lambda_2 - \lambda_1} \quad (3)$$

If the CDS market is dominant, *Market Share* will be close to 1. On the other hand, if the bond market is dominant, it will be close to 0. If the two markets contribute similar amounts to the price discovery, the ratio will be close to 0.5.

When the CDSSs and BSs turn out not to be cointegrated, the VECM is not valid. In this case, in order to study the lead-lag relationship between the historical series, it is necessary to apply the Granger (1969) causality test, in line with Forte (2009). In detail, given two series X and Y, one says that X “Granger causes” Y if the past values of X contain useful information, other than that contained in the past values of Y, for explaining the current value of Y.

In order to determine whether, in the price discovery process, there exists a leading market between the CDS and bond markets, or whether both play a significant role, Granger’s causality test, as just described, is applied to the following regressions 4 and 5, in which the terminology is the same as that in equations 1 and 2:

$$\Delta CDS_t = \alpha_1 + \sum_{i=1}^p \beta_{1i} \Delta CDS_{t-i} + \sum_{i=1}^p \gamma_{1i} \Delta BOND_{t-i} + \varepsilon_{1t} \quad (4)$$

$$\Delta BOND_t = \alpha_2 + \sum_{i=1}^p \beta_{2i} \Delta CDS_{t-i} + \sum_{i=1}^p \gamma_{2i} \Delta BOND_{t-i} + \varepsilon_{2t} \quad (5)$$

In detail, equations 4 and 5 represent a VAR model without ECT and are obtained by eliminating from equations 1 and 2 the terms “ $\lambda_1 Z_{t-1}$ ” and “ $\lambda_2 Z_{t-1}$ ” due to the absence of

cointegration. Specifically, the Granger causality test of the coefficients of the VAR model considers the following three cases⁹:

1. If, in equation 4, the null hypothesis that the BSs do not Granger cause the CDSSs, that is $\gamma_{1i} = 0$ for $i=1, \dots, p$, is rejected, one may conclude that the bond market contributes to the price discovery.
2. If, in equation 5, the null hypothesis that the CDSSs do not Granger cause the BSs, that is $\beta_{2i} = 0$ for $i=1, \dots, p$, is rejected, one may conclude that the CDS market contributes to the price discovery.
3. If the tests reject the null hypothesis in both directions of causality, one may conclude that the price discovery is manifested in both markets.

In the first two cases, when only one of the two directions of causality is significant, a single market contributes (*Market Share* = 100%) to the price discovery. This is termed the leading market, in the same way as in the approach adopted in the VECM. It should be observed that, in the third case, or when both directions of causality are significant, it is not possible to apply the *Market Share* formula, unlike in the case of the VECM.

In a nutshell, the first phase of the analysis verifies the possible presence of cointegration between the CDSSs and BSs. The second phase of the analysis investigates the lead-lag relationship between the two markets using either the VECM or the VAR, depending on whether there is a presence or lack of cointegration¹⁰. The two methodological phases are repeated for each Eurozone country and for each of the three periods indicated earlier.

⁹ It is appropriate to underline some limitations of the Granger causality test when applied in this context. If the variables are linked by a cointegrating relationship, the omission of the ECT, or the process of adjustment towards the long-term equilibrium, will produce a specification error in the model that could distort the estimations of the coefficients and of the statistical test in an unpredictable direction. At the same time, it is not possible to include the ECT in cases where the hypothesis of cointegration has been rejected. The results should therefore be interpreted with caution. However, given that the interest of this study lies predominantly in investigating the lead-lag relationship and not the magnitude of the parameters, the test utilized provides sufficiently robust empirical evidences. See Forte (2009).

¹⁰ One resorts to the VAR model in two cases: (1) if the Johansen trace test accepts the null hypothesis of an absence of cointegration; (2) if the ADF test for a unit root rejects the hypothesis of non-stationarity on the levels for at least one of the two series of credit spreads, as a consequence of which it is not possible to carry out the trace test of cointegration.

5. DATA ANALYSIS

The empirical results of this paper are presented in the following Tables 1, 2 and 3, in each one the Core countries of the Eurozone are grouped on the left, and the Peripheral on the right. The tables refer respectively to the following time periods:

- the period of the financial crisis (15/9/2008-14/9/2010), Table 1;
- the period of the sovereign debt crisis (15/9/2010-14/9/2012), Table 2;
- the period of contraction of credit spread quotations (15/9/2012-14/9/2014), Table 3.

In the tables, each column, for each country, repeats the methodological process to identify the dominant market in the credit risk price discovery process. This market is explicitly indicated, for each country and for each of the three intervals studied, in the penultimate row of each table.

As one can deduce from the tables, the price discovery analysis can be summarised as follows:

- In the first phase rows, the results of the unit root tests on the levels and on the first differences of the two series of credit spreads are presented, and then the estimates of the trace statistics, to verify the presence or lack of cointegration.
- In the second phase rows, the estimates of the VECM or VAR model are presented. The VECM is used when the two series are cointegrated (trace test). The VAR model is used in one of two cases:
 - when the unit root test has rejected the hypothesis of non-stationarity on the levels for at least one of the two series, such that it is not possible to estimate the trace statistic, or
 - when the trace statistic has led to the acceptance of the hypothesis of the absence of cointegration.
- In the penultimate row, the conclusions about price discovery are indicated, inferred from the estimates of the VECM or VAR model. The last row contains, where identifiable, the proportion of informative power (*Market Share*) of the dominant market in the process of price discovery.

6. DISCUSSION OF DATA ANALYSIS

In Table 1 (15/9/2008-14/9/2010, CORE COUNTRIES), the estimates from the first phase of analysis show that the CDSSs and BSs for Belgium and France follow a process of returning to equilibrium in the long term. This emerges from the trace statistic that rejects the hypothesis of the absence of cointegration. For Austria, Finland and Holland, the rejection of the preliminary hypothesis of non-stationarity in the levels of one of the two series of credit spreads prevented us from using the trace statistic. In detail, for the Austrian and Dutch CDSSs and the Finnish BSs the hypothesis of non-stationarity was rejected. This evidence explicitly obliged us to apply the second phase of the analysis using the VECM for Belgium and France, and the VAR model for the other Core countries. The results underline that, for Belgium and France, the only market that contributes to price discovery is the CDS market, given that only coefficient of adjustment λ_2 is significant. The Granger causality test applied to the VAR models for Austria and Finland confirms that even in these cases only the CDS market contributes significantly to the price discovery. In contrast, for Holland, the Granger causality test shows that the process is sustained solely by the bond market.

In Table 1 (15/9/2008-14/9/2010, PERIPHERAL COUNTRIES), the estimates from the first phase of analysis show that the CDSSs and BSs from all the Peripheral countries, except for Slovakia, follow a cointegrated process. The results of the second phase of analysis show that, for Greece, Ireland and Spain, it is exclusively the CDS market that contributes to the price discovery, given that only λ_2 (coefficient of adjustment) is significant. For Slovakia the same result is achieved, but from the Granger causality test that is applied to the VAR model. In Italy, the significance levels of both coefficients of adjustment, λ_1 and λ_2 , show that both the CDS and the bond market contribute to the process of price discovery. For the Italian market, the significance levels of both coefficients (λ_1 and λ_2) allow us to calculate *Market Share*, from which we learn that the informative power of the CDS market is equal to 66.7%.

In Table 2 (15/9/2010-14/9/2012, CORE COUNTRIES) the estimates from the first phase of analysis show that the CDSSs and BSs of Austria, Finland and France follow a cointegrating process. For Belgium the hypothesis of the absence of cointegration can be accepted. For Holland, the rejection of the preliminary hypothesis of non-stationarity of the levels of the BSs meant that it was not possible to run the trace statistic. The results of the second phase of analysis show that, for Belgium, Finland and Holland, the CDS market alone contributes to the

price discovery. In Austria and France, both the CDS and the bond market contribute. However, for both countries, the calculation of *Market Share* reveals that the CDS market has more informative power, at 69% and 71.3% respectively.

In Table 2 (15/9/2010-14/9/2012, PERIPHERAL COUNTRIES) the estimation of the first phase of analysis shows that the CDSSs and BSs of Greece, Italy, Portugal and Slovakia follow a cointegrating process. For Spain, the hypothesis of the absence of cointegration can be accepted. For Ireland, the rejection of the preliminary hypothesis of non-stationarity of the levels of the BSs and CDSSs meant that it was not possible to run the trace statistic. The results of the second phase show that for Greece, Italy and Portugal only the bond market contributes to the price discovery, in view of the fact that only λ_1 (adjustment coefficient) is significant. It is hardly necessary to underline that, for Greece, the effect of the adjustment of the CDSSs towards the BSs is more intense ($\lambda_1=0.14$). In Slovakia, both the CDS and the bond market contribute to the price discovery process. However, the calculation of *Market Share* reveals that the bond market has greater informative power, at 66.1%. For Ireland and Spain, the Granger causality test applied to the VAR model shows that both the CDS and bond markets contribute significantly to the price discovery.

In Table 3 (15/9/2012-14/9/2014, CORE COUNTRIES) the estimates from the first phase of analysis show that only for France do the CDSSs and BSs follow a cointegrating process. For Belgium and Holland, the hypothesis of the absence of cointegration is accepted. For Austria, the rejection of the preliminary hypothesis of non-stationarity of the levels of the BSs meant that it was not possible to run the trace statistic. The results of the second phase show that, for France, only the CDS market contributes to the price discovery, given that λ_2 (adjustment coefficient) is significant. For Austria and Holland, the Granger causality test applied to the VAR model evidences a significant contribution to the price discovery from the bond market alone. For Belgium and Finland, neither the CDS nor the bond market turns out to be significant for price discovery.

In Table 3 (15/9/2012-14/9/2014, PERIPHERAL COUNTRIES) the estimations of the first phase of analysis evidence that the CDSSs and BSs of Ireland, Italy and Portugal follow a cointegrating process. For Slovakia and Spain, the hypothesis of the absence of cointegration can be accepted. It is scarcely necessary to note that, for Greece, the small number of

observations makes it inadvisable to run the analysis¹¹. The results of the second phase of analysis show that, for Italy and Portugal, only the bond market contributes to the price discovery. The only coefficient of adjustment that is significant is λ_1 , in fact. In Ireland, both the CDS and the bond market contribute towards determining the price. However, the *Market Share* calculation shows that the bond market has slightly more informative power, at 58.5%. The Granger causality test applied in the case of Spain demonstrates that solely the bond market provides a significant contribution. For Slovakia, both markets contribute significantly to the price discovery.

In the light of what has emerged from the above detailed evaluation of the results, it can be confirmed that during the financial crisis the CDSSs dominate price discovery. In the second period the CDS market's share of informative power diminishes and the bond market assumes greater relevance. In the third period, price discovery is dominated by the BSs. This pattern is more marked for the high-yield (Peripheral) countries and more restrained for the low-yield (Core) countries.

The robustness of the theory and the correct application of econometric model is inferable from the discord in the signs of the adjustment coefficients. For the countries that exhibit cointegrated CDSSs and BSs, the estimated values of λ_1 are negative, and those of λ_2 positive. In other words, the signs of the coefficients are consistent with mean reversion of the CDS-bond basis in the long term. This evidence confirms the theoretical expectation of a diffuse pattern of adjustment in the long term between the CDS and bond markets. It is hardly necessary to observe that the levels of significance and speed of adjustment of the coefficients, even if generalizable to the aggregate level, assume specific connotations at the level of the single country and with reference to each of the three time periods investigated¹².

¹¹ This is due to the lack of issuances of Greek government bonds for three or four years.

¹² An approach that allows the reader to comprehend the abovementioned specific connotations at the level of the country and the time period is to analyse the results in Tables 1, 2 and 3 according to two interpretations:

- a) the heterogeneity between the 11 countries within the same time period (in terms of level of significance and speed of adjustment coefficients);
- b) the heterogeneity between the 3 time periods for each single country (in terms of the level of significance and speed of adjustment coefficients).

For example, regarding the heterogeneity between countries, during the sovereign debt crisis (second period), in France, Austria and Slovakia, the p-values (or t-statistics) and the signs of the coefficients of adjustment (λ_1 and λ_2) reveal that the CDS market contributes just as significantly as the bond market to the phenomenon of adjustment towards equilibrium). On the contrary, again in the second period, in Italy, Greece and Finland, even though the signs of both coefficients of adjustment are consistent with the phenomenon of mean reversion of the CDS-bond basis, the process of adjustment is realized significantly (p-values below 10%) in just a single market: in Finland in the bond market, in Italy and Greece in the CDS market, with significance levels of 1%, 10% and

7. CONCLUSIONS

This study has analysed the dynamics of the relationships between BSs and CDSSs in the price discovery process of credit risk in the Core and Peripheral countries of the Eurozone, from 2008 to 2014. In order to fulfil the two intermediate objectives, the analysis of price discovery was carried out in two phases. In the first phase it was investigated whether the misalignments of the CDSSs and BSs exhibited a mean-reverting effect, or a return towards the equilibrium in the long term. In the second phase a lead-lag analysis was carried out, looking at which market anticipated the other in pricing credit risk. Both methodological phases were carried out in three time periods: 2008-2010, 2010-2012 and 2012-2014.

This methodological approach has allowed us to present the changes in the actual price discovery dynamics, for every country in the Eurozone.

From the first phase of analysis, the following can be observed:

- During the periods with greater volatility (2008-2010 and 2010-2012), it is possible to verify the existence of an equilibrium relationship in the long term between the two credit spreads.
- During the period with less financial turbulence (2012-2014), the two series turn out to be less strongly related in the long term¹³.

These results can be interpreted as follows:

The action of the market forces is able to close the divergence between the two credit spreads towards an equilibrium in the long term, more frequently in periods of high volatility. From

1% respectively. As well as the significance level, the speed of adjustment varies according to the country considered. For example, again in the second period, the Greek CDSSs exhibit a speed of adjustment (0.145) towards the bond market far greater than the Italian CDS market (0.0029), signifying that during the sovereign debt crisis the bond market had greater informative power in Greece than in Italy.

Similar findings emerge for the case of heterogeneity between time periods. For example, in France, during the financial crisis and the sovereign debt crisis (first and second periods), the CDSSs were more efficient than the BSs in the price discovery process. However, from the observation of the significance levels and speeds of adjustment of the adjustment coefficients there emerge differences between the first and second periods. The CDS Market Share goes from 100% in the first period to 71.26% in the second. This change reflects the lack of significance of λ_1 in the first period and its significance (at 5%) in the second. In other words, only in the second period there is significant adjustment towards equilibrium from the CDS market as well, and therefore statistically significant informative power for the bond market (28.74%), albeit inferior to that of the CDS market. The process of adjustment towards equilibrium on the part of the bond market is marked, in the two periods of analysis, by the same level of significance of λ_2 (5%) but different speeds of adjustment (0.028 in the first period and 0.059 in the second).

¹³ The low presence of cointegration in the period of low volatility is in line with Fontana and Sheicher (2010), who find an absence of cointegration for the Eurozone countries in the period preceding the financial crisis.

the operating point of view, that is equivalent to a situation in which the market forces come into play when the misalignment between the two credit spreads is sufficiently wide for strategies to be implemented that are profitable net of transaction costs.

- The effect of mean reversion in the long term is less evident for the low-yield (Core) countries in all three periods.

This result can be interpreted as follows:

Especially in periods of turbulence, market participants look to rebalance their positions by investing in those bond markets that are considered risk free, such as those of the Core countries. This generates, for the low-yield countries, an increasing volume of trading in the bond market and lower liquidity in the CDS market. This causes the absence of a stable relationship between the two credit spreads in the long term.

From the second phase of analysis, the following can be observed:

- During the financial crisis (2008-2010), the CDS market assumes the dominant role in the price discovery of credit risk in the Core and Peripheral countries.

This result can be interpreted as follows:

The period of financial crisis triggers a strong demand for protection and consequently CDS are more convenient and simpler instruments for trading credit risk. The principal reasons for positions on credit risk being taken up through the CDS market rather than the bond market are the following:

- *the greater ease of buying and selling large quantities of credit risk with CDS, given the unfunded nature of CDS and the funded nature of bonds;*
- *the difficulty of short selling bonds for those who wish to engage the protection of the bond market, in particular in periods of stress and high default risk.*

These circumstances shift the liquidity towards the CDS market, making this market de facto more efficient at pricing credit risk.

- During the period of the sovereign debt crisis (2010-2012), both the CDS and the bond market contribute to the price discovery of credit risk. As a consequence, relative to the preceding period, the share of informative power of the CDS market falls in favour of the bond market. This tendency is sharply more evident for the Peripheral countries, particularly Greece, in comparison to the Core countries. It is hardly necessary to

underline that this result is particularly interesting in view of the related works that have investigated the 2008-2010 period of financial crisis. Such works conclude that, during the periods of greatest perception of credit risk in the euro area, the CDS market dominated price discovery for the high-yield countries. In contrast, the present work demonstrates the significant role played by the bond market in the Peripheral countries, during the sovereign debt crisis.

This result can be interpreted as follows:

Following the downgrading of sovereign debt by the rating agencies, there is a more objective evaluation of credit risk with respect to the period of 2008-2010. The investors begin to differentiate their positions within the European market between Core and Peripheral countries. The speculation about the disintegration of the euro area triggers market tensions concerning the Peripheral bonds. This causes large volumes of capital to be moved towards the bond market, increasing the informative power of the bond spreads.

- During the period marked by a progressive downsizing of credit risk in Eurozone (2012-2014), the informative power of the bond market grows still further with respect to that of the CDS market.

This result can be interpreted as follows:

The lower trading activity in the CDS market and the greater liquidity in the bond market contribute towards the bond spreads' active role in price discovery.

Clearly, the concepts of volatility and liquidity assume a key role in the interpretation of the results of this study. Volatility represents the fundamental premise for market forces having adequate margins to exploit the misalignment between the bond and CDS spreads. Liquidity represents the fundamental requirement of the leading market in terms of price discovery, which occurs in the market that is more informed and liquid.

8. PRACTICAL IMPLICATIONS

The results that emerge from the empirical analysis support the scientific studies that investigate the microstructure of the financial markets. Meanwhile, they are particularly useful for market operators and policy makers, too. For market operators, the ability to identify the leading and following instruments of price discovery offers them the opportunity to profit from

the implementation of strategies that exploit the process of adjustment between the two markets. For policy makers, in their capacity as guarantors of price stability and in their function as controllers of any systemic crises, it would seem appropriate for them to have a deep knowledge of the dynamics of price discovery. The identification of leading instruments, related to specific phases of the market (of volatility, liquidity, credit risk etc.), in fact facilitates the assessment of adequate and efficient monetary and fiscal policies. Nevertheless, the policy makers should take due notice of the fact that their own manoeuvres will in turn affect the future dynamics of price discovery.

9. TABLES

Table 1: Price Discovery of Credit risk in Eurozone - First Period: 15/9/2008-14/9/2010 - Core Countries (Austria, Belgium, Finland, France, Netherlands) and Peripheral Countries (Greece, Ireland, Italy, Portugal, Slovakia, Spain).

	CORE COUNTRIES					PERIPHERAL COUNTRIES						
	AUSTRIA	BELGIUM	FINLAND	FRANCE	NETHERLANDS	GREECE	IRELAND	ITALY	PORTUGAL	SLOVAKIA	SPAIN	
I P H A S E	FIRST PERIOD											
	UNIT ROOT Analysis											
	CDS Level	-4.64***	-2.02	-2.34	-2.25	-2.85*	-1.23	-1.48	-2.46	-1.19	-1.46	-1.83
	BOND Level	-2.35	-2.18	-3.26**	-2.31	-2.55	-0.73	-1.09	-1.87	-1.34	-1.63	-1.69
	CDS Difference	-6.60***	-13.00***	-26.53***	-20.85***	-20.17***	-20.53***	-11.84***	-18.68***	-18.71***	-15.47***	-13.36***
	BOND Difference	-18.94***	-19.97***	-11.18***	-22.11***	-21.33***	-19.42***	-12.18***	-16.78***	-18.33***	-15.96***	-15.72***
	COINTEGRATION Analysis											
	TRACE Test	-	13.54*	-	15.37*	-	14.99*	16.61**	21.80***	16.25**	11.40	16.76**
	VECM Analysis											
	Lambda 1	-	-0.012436	-	-0.004474	-	-0.021992	-0.015	-0.019365	-0.012656	-	-0.022454
[t-stat]	-	-1.44	-	-0.73	-	-1.38	-0.68	-1.83*	-0.91	-	-1.54	
Lambda 2	-	0.017017	-	0.028639	-	0.027807	0.092	0.038743	0.046238	-	0.039072	
[t-stat]	-	1.98**	-	2.56**	-	1.94*	3.36***	2.63***	2.84***	-	2.26**	
VAR Granger causality												
BOND	2.53	-	2.94	-	7.89**	-	-	-	-	2.12	-	
CDS	8.66**	-	9.05**	-	0.06	-	-	-	-	6.41**	-	
PRICE DISCOVERY	CDS	CDS	CDS	CDS	BOND	CDS	CDS	CDS>BOND	CDS	CDS	CDS	
Market Share	100%	100%	100%	100%	100%	100%	100%	66.7%	100%	100%	100%	

***, ** and * denote:

- in UNIT ROOT Analysis, rejection at the 1%, 5% and 10% levels, respectively, of the null hypothesis of non-stationarity;
- in COINTEGRATION Analysis, rejection at the 1%, 5% and 10% levels, respectively of the null hypothesis of absence of cointegration;
- in VECM Analysis, significance of adjustment coefficients at the 1%, 5% and 10% levels, respectively;
- in VAR Analysis, significance of Granger causality at the 1%, 5% and 10% levels, respectively.

Table 2: Price Discovery of Credit risk in Eurozone - Second Period: 15/9/2010-14/9/2012 - Core Countries (Austria, Belgium, Finland, France, Netherlands) and Peripheral Countries (Greece, Ireland, Italy, Portugal, Slovakia, Spain).

	CORE COUNTRIES					PERIPHERAL COUNTRIES							
	AUSTRIA	BELGIUM	FINLAND	FRANCE	NETHERLANDS	GREECE	IRELAND	ITALY	PORTUGAL	SLOVAKIA	SPAIN		
I P H A S E	SECOND PERIOD												
	UNIT ROOT Analysis												
	CDS Level	-0.87	-1.59	-1.58	-1.46	-1.47	-0.41	-6.34***	-1.25	-1.71	-1.46	-1.99	
	BOND Level	-1.63	-1.35	-2.29	-1.73	-2.63*	0.53	-6.63***	-1.55	-1.67	-2.30	-1.96	
	CDS Difference	-18.83***	-17.73***	-29.62***	-19.75***	-17.65***	-18.82***	-19.57***	-15.96***	-17.73***	-17.55***	-16.36***	
	BOND Difference	-22.58***	-19.45***	-21.14***	-22.27***	-23.49***	-13.05***	-21.77***	-19.15***	-17.23***	-20.75***	-18.63***	
	COINTEGRATION Analysis												
	Trace test	14.64*	11.72	17.41**	21.11***	-	22.01***	-	13.47*	13.69*	16.33**	12.86	
	II P H A S E	VECM Analysis											
		Lambda 1	-0.010069	-	-0.013251	-0.024135	-	-0.145080	-	-0.029933	-0.043162	-0.012495	-
[t-stat]		-2.00**	-	-1.59	-2.41**	-	-3.82***	-	-1.93*	-3.05***	-1.85*	-	
Lambda 2		0.022462	-	0.049295	0.059852	-	0.013849	-	0.006809	-0.013976	0.024399	-	
[t-stat]		1.97**	-	3.27***	2.39**	-	0.50	-	0.34	-0.84	2.42**	-	
VAR Granger causality													
BOND		-	3.12	-	-	2.97	-	39.77***	-	-	-	6.30**	
CDS		-	10.91***	-	-	5.02*	-	30.93***	-	-	-	4.78*	
PRICE DISCOVERY													
Market Share		CDS>BOND	CDS	CDS	CDS>BOND	CDS	BOND	CDS&BOND	BOND	BOND	CDS>BOND	CDS&BOND	
	69%	100%	100%	71.26%	100%	100%	n.d.	100%	100%	66.1%	n.d.		

***, ** and * denote:

- in UNIT ROOT Analysis, rejection at the 1%, 5% and 10% levels, respectively, of the null hypothesis of non-stationarity;
- in COINTEGRATION Analysis, rejection at the 1%, 5% and 10% levels, respectively of the null hypothesis of absence of cointegration;
- in VECM Analysis, significance of adjustment coefficients at the 1%, 5% and 10% levels, respectively;
- in VAR Analysis, significance of Granger causality at the 1%, 5% and 10% levels, respectively.

Table 3: Price Discovery of Credit risk in Eurozone - Third Period: 15/9/2012-14/9/2014 - Core Countries (Austria, Belgium, Finland, France, Netherlands) and Peripheral Countries (Ireland, Italy, Portugal, Slovakia, Spain).

		CORE COUNTRIES					PERIPHERAL COUNTRIES				
THIRD PERIOD		AUSTRIA	BELGIUM	FINLAND	FRANCE	NETHERLANDS	IRELAND	ITALY	PORTUGAL	SLOVAKIA	SPAIN
I P H A S E	UNIT ROOT Analysis										
	CDS Level	-2.44	-2.09	-2.35	-1.24	-0.80	-0.55	-0.48	-0.45	-0.89	0.26
	BOND Level	-3.90***	-1.77	-3.25**	-2.12	-1.14	-0.02	-0.41	-0.21	-1.13	-0.33
	CDS Difference	-24.61***	-34.27***	-20.80***	-10.88***	-10.79***	-21.40***	-19.77***	-20.08***	-18.56***	-12.57
	BOND Difference	-13.92***	-18.82***	-12.24***	-20.65***	-21.60***	-19.48***	-20.84***	-19.36***	-22.31***	-22.43
	COINTEGRATION Analysis										
	Trace test	-	12.42	-	14.55*	7.14	14.28*	14.21*	13.59*	10.62	11.00
	VECM Analysis										
	Lambda 1	-	-	-	-0.002848	-	-0.020964	-0.031318	-0.048154	-	-
	[t-stat]	-	-	-	-0.74	-	-2.34**	-1.71*	-2.26**	-	-
Lambda 2	-	-	-	-0.036981	-	0.029577	0.022988	0.016346	-	-	
[t-stat]	-	-	-	3.18***	-	1.93*	1.07	0.73	-	-	
VAR Granger causality											
BOND	6.15**	3.46	0.38	-	5.70*	-	-	-	8.76**	21.47***	
CDS	4.28	0.69	0.15	-	1.16	-	-	-	6.39**	0.24	
PRICE DISCOVERY	BOND	n.d.	n.d.	CDS	BOND	CDS>BOND	BOND	BOND	CDS&BOND	BOND	
Market Share	100%	n.d.	n.d.	100%	100%	58.5%	100%	100%	n.d.	100%	

***, ** and * denote:

- in UNIT ROOT Analysis, rejection at the 1%, 5% and 10% levels, respectively, of the null hypothesis of non-stationarity;
- in COINTEGRATION Analysis, rejection at the 1%, 5% and 10% levels, respectively of the null hypothesis of absence of cointegration;
- in VECM Analysis, significance of adjustment coefficients at the 1%, 5% and 10% levels, respectively;
- in VAR Analysis, significance of Granger causality at the 1%, 5% and 10% levels, respectively.

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A Time-Varying Price Discovery Model for Policy Analysis

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ABSTRACT

The paper proposes to exploit the informative content of *price discovery* models for policy analysis. In particular, for analysing the effects of monetary policies, of bailout policies and of regulatory innovations on the microstructure of financial markets as well as for measuring the effectiveness of new policies in reducing the financial instability.

So far, most of the applications of *price discovery* models are in finance to investigate the microstructure dynamics of financial markets. The paper argues that, in order to be suitable for policy analyses, standard *time-invariant price discovery* models should be applied in a *time-varying* framework.

In the first part, a detailed literature review presents the most important price discovery measures. Along with the theoretical framework, the paper provides the economic interpretations of each price discovery measure and compares them.

In the second part, a *time-varying* technique is applied to a topic currently investigated in literature “the impact of Naked CDS Ban on the price discovery of sovereign credit risk in Eurozone”. The empirical evidences show the advantage of *time-varying* models in being informative on the impact of Credit Default Swap regulation on the *price discovery* of sovereign credit risk.

JEL classification: G01, G12, G14, G18

Keywords: price discovery measures, time-varying VECM, policy analysis

1. INTRODUCTION

For policy makers, in their role of guarantee of price stability and in their purpose of avoiding systemic crises, it seems appropriate a thorough understanding of the dynamics of price discovery. The knowledge of these dynamics in relation to specific market phases (as regards credit risk, liquidity, volatility and so on) helps, in fact, the implementation of monetary policies, government policies and financial regulations appropriate and efficient. Nonetheless, policy makers must take into account that the configuration of their own measures in turn affects the future dynamics of price discovery.

Price discovery is one of the central functions of financial markets, defined by Lehmann (2002) to be “the efficient and timely incorporation of the information implicit in investor trading into market prices”. When multiple financial instruments express in their price series the value of a common asset¹⁴, the contribution of a price series to price discovery is typically considered to be the extent to which it is the first to reflect new information about the ‘true’ underlying asset value.

Limited attention has been devoted in literature to use price discovery models for policy analyses. In particular, for analysing the effects of monetary policies, bailout policies by national governments and regulatory innovations on the microstructure of financial markets as well as for measuring the effectiveness of new policies in reducing the financial instability.

Ultimate aim of the paper is to provide an econometric model able for investigating these economic interests. In order to reach this objective, the research idea is to adjust current “price discovery” models, so far mostly used in finance theory for studying the microstructure dynamics of financial markets, in a way that are suitable for policy analysis. The research idea is to develop a model able to estimate time-varying dynamics in the price discovery process. This means to leave the model free to estimate when structural changes happen in the price discovery dynamics. We argue that economic policies play an important role in these changes. Such a model could contribute on the literature of both financial market microstructure and policy analysis.

The paper is organized as follows. In the next section, we describe in detail the most important approaches used in literature to model and measure price discovery. These are Gonzalo and Granger (1995) “permanent–transitory” (PT) decomposition and Hasbrouck (1995) “information share” (IS). These two approaches are directly related and the results of both models are primarily derived from the Vector Error Correction Model (VECM). Moreover, we compare PT and IS and we highlight the

¹⁴ Some of the most investigated cases in literature are: a) derivatives and cash securities linked to the same underlying asset, and b) stocks on the same entity traded in multiple venues.

economic importance of a new measure of price discovery “information leadership share” (ILS) introduced by Yan and Zivot (2010) and Putniņš (2013). In section 3, we propose our extension of standard time-invariant price discovery models. It consists in applying a moving window (MW) technique to the VECM and deriving time-varying PT, IS and ILS. We show the economic importance of studying price discovery dynamics in a time-varying framework by applying the MW technique to a topic currently investigated in literature “the impact of Naked Credit Default Swap (CDS) Ban on the price discovery of sovereign credit risk in Eurozone”. This technique is informative on the impact of CDS regulation on price discovery of sovereign credit risk. We describe the empirical results and derive policy implications.

2. THEORETICAL MODELS

In this section, we first show how price dynamics of related assets traded in two different markets¹⁵ can be modelled using a VECM. Subsequently, we show how to derive, from a VECM framework, PT, IS and ILS metrics to measure price discovery. We give the economic interpretation of each price discovery measure and compare them.

2.1. Vector Error Correction Model

Suppose a security trades on two markets ($i=1,2$) at potentially different observed transaction prices p_{1t} and p_{2t} . Market microstructure theory assumes:

- ass.1: $m_t = \lim_{h \rightarrow \infty} E[p_{1,t+h}|I_t] = \lim_{h \rightarrow \infty} E[p_{2,t+h}|I_t]$,

there is a common underlying implicit efficient price (m_t). This unobserved efficient price represents the underlying value of a security conditional on all public information (I) available at time t , that at minimum includes past prices. All investors share the same public information set, and prices are efficient in the sense that the current price reflects future price expectations conditional on the available information set;

- ass. 2: $s_{i,t} = p_{i,t} - m_t$ and $E[\lim_{h \rightarrow \infty} s_{i,t+h}|I_t] = 0$,

at time t , prices in the two markets, potentially, differ because of microstructural noise (e.g. liquidity, asymmetric information). Microstructural noise (or market frictions) should not

¹⁵ All formulae and interpretations can be easily extended for the case of more than two markets.

affect the price of a security in the indefinite future and thus should be a “*transient*” or “*temporary*” component of prices.

Next step is to arrive to the VECM representation starting from the previous assumptions. We need three additional assumptions:

- ass. 1’: Suppose that for the two prices $p_{1,t}$ and $p_{2,t}$ the continuous sequence of the implicit efficient price (m_t) is an integrated process of order one I(1) that can be represented as a random walk:

$$m_t = m_{t-1} + w_t \text{ with } w_t \sim iid, N(0, \sigma_w^2), \quad (1)$$

where w_t is the random informational arrival over the interval $[m_{t-1}, m_t]$. Because m_t exhibits no tendency to mean revert I(1), information arrivals lead to permanent shocks that cumulate over time into a stochastic trend $\sum w_t$;

- ass. 2’: observed trading prices $p_{1,t}$ and $p_{2,t}$ impound the w_t information arrivals but each differ from m_t (implicit common efficient price) by a zero-mean, covariance stationary (weakly stationary) random disturbance $s_{1,t}$ and $s_{2,t}$:

$$p_{1,t} = m_t + s_{1,t}, \quad (2)$$

$$p_{2,t} = m_t + s_{2,t}, \quad (2')$$

where $s_{1,t}$ and $s_{2,t}$ are assumed to be identically distributed through time but may be autocorrelated. The disturbances are intended to model transitory shocks of microstructural noise. Eq. 2 and 2’ can be seen as integrated process of random walk and news innovations plus the market frictions observed at time t;

- ass. 3: it is allowed that lagged information arrivals w_{t-h} can plausibly affect s_{it} but it is assumed that lagged disturbances $s_{i,t-h}$ cannot improve the prediction of w_t upon the univariate forecast of w_t . In other words, it is allowed that price innovation w_t can be correlated with future values of the microstructural noise $s_{i,t}$, but it is assumed that $s_{1,t}$ and $s_{2,t}$ do not Granger-cause w_t .

Rewriting eq. 2 and 2’ in terms of first differences it follows:

$$\Delta p_{1,t} = p_{1,t} - p_{1,t-1} = \Delta m_t + \Delta s_{1,t} = w_t + \Delta s_{1,t}, \quad (3)$$

$$\Delta p_{2,t} = p_{2,t} - p_{2,t-1} = \Delta m_t + \Delta s_{2,t} = w_t + \Delta s_{2,t}, \quad (3')$$

then, rewriting eq. 3 and 3' in terms of price levels it follows:

$$p_{1,t} = p_{1,t-1} + w_t + \Delta s_{1,t} \text{ and } p_{2,t} = p_{2,t-1} + w_t + \Delta s_{2,t}. \quad (4)$$

Note, the unobserved efficient price follows a random walk while observed transaction prices not.¹⁶ Nevertheless, at any realization ($t=T$), both trading price sequences impound the stochastic trend in the implicit efficient price as a common factor:

$$p_{1,T} = p_{1,0} + \sum_{t=1}^T w_t + s_{1,T} \text{ and } p_{2,T} = p_{2,0} + \sum_{t=1}^T w_t + s_{2,T}; \quad (5)$$

Where each observed trading price ($p_{1,T}$ and $p_{2,T}$) level can be seen to depend on:

- a) a non-stochastic initial value with $t=0$ chosen so that $p_{1,0} = p_{2,0}$,
- b) the common stochastic trend of the cumulated random information arrivals $\sum w_t$.
- c) a zero-mean covariance stationary process that is specific to the trading market and to time period T and is therefore an idiosyncratic transitory disturbance.

Since prices in both markets are driven by the same underlying fundamentals, the prices should be cointegrated. More precisely, although both trading price series are non-stationary the difference between their contemporaneous values is itself a stationary time series $I(0)$:

$$p_{1,t} - p_{2,t} = s_{1,t} - s_{2,t}, \quad (6)$$

i.e., the sum of two zero-mean covariance-stationary random disturbances. When linear combinations of integrated $I(1)$ variables like $p_{1,t}$ and $p_{2,t}$ are themselves $I(0)$, the underlying series are cointegrated. Therefore, by the Granger Representation Theorem for cointegrated variables, $\Delta p_{1,t}$ and $\Delta p_{2,t}$ can be estimated as a VECM:

$$\Delta p_{1,t} = \sum_{h=1}^l \beta_{1,t-h} \Delta p_{1,t-h} + \sum_{h=1}^l \gamma_{1,t-h} \Delta p_{2,t-h} + \lambda_1 Z_{t-1} + c_1 + u_{1,t}, \quad (7)$$

$$\Delta p_{2,t} = \sum_{h=1}^l \beta_{2,t-h} \Delta p_{1,t-h} + \sum_{h=1}^l \gamma_{2,t-h} \Delta p_{2,t-h} + \lambda_2 Z_{t-1} + c_2 + u_{2,t}, \quad (7')$$

where:

- Z_{t-1} is an error correction term (i.e. $Z_{t-1} = p_{1,t-1} - p_{2,t-1}$);
- l optimal lag length in the corresponding system of VAR equations;

¹⁶ Only if innovations in microstructural noise are unchanged ($\Delta s_{1,t}$ and $\Delta s_{2,t}=0$) will the trading price sequence of first differences $\Delta p_{1,t}$ and $\Delta p_{2,t}$ exactly mirror the implicit efficient sequence of first differences $\Delta m_t = w_t$.

- $u_{i,t}$ empirical estimation errors which are expected to be contemporaneously correlated across markets 1 and 2, i.e. $\text{Cov}(u_{1,t}, u_{2,t}) \neq 0$;
- c_i constants;
- $\beta_{i,t-h}$ and $\gamma_{i,t-h}$ are the auto-regressive (AR) coefficients;
- λ_1 and λ_2 measure the speed of adjustment to the error correction term of each market i .

The VECM has two parts:

- 1) the AR part composed by the first two expressions on the r.h.s of eq. 7 and 7'. It represents the short-term dynamics;
- 2) Z_{t-1} represents the long-run equilibrium between the price series.

2.2. Price discovery models

The VECM as expressed in eq. 7 and 7' has been used extensively to study the price discovery of a security traded in multiple markets. The most adopted price discovery measures are Gonzalo and Granger (1995) permanent–transitory (PT) decomposition and the Hasbrouck (1995) information share (IS) measures. The results of both approaches are primarily derived from the VECM and, then, they are obviously related.

2.2.1. Gonzalo and Granger (1995) “permanent–transitory” (PT) decomposition

Gonzalo and Granger (1995) propose a method of decomposing a co-integrated price series into a permanent component and a temporary component using the error correction coefficients λ_1 and λ_2 . Recalling previous assumptions and specifications, we can rewrite a more theoretical (not observable) VECM substituting eq 3, 3' and 6 into eq. 7 and 7' (with $i=1,2$ markets):

$$\Delta p_{i,t} = \lambda_{\perp i} \sum_{h=1}^l w_{t-h} + \sum_{h=1}^l \eta_{i,t-h} \Delta s_{i,t-h} + \lambda_i (s_{1,t-1} - s_{2,t-1}) + c_i + u_{i,t}; \quad (8)$$

where:

- $\lambda_{\perp i}$ common factor weights for each market;
- $\sum_{h=1}^l w_{t-h}$ common stochastic trend;
- $s_{i,t}$ idiosyncratic transitory disturbances in each market;

- $(s_{1,t-1} - s_{2,t-1})$ error correction term specified as an intermarket price divergence $(p_{1,t} - p_{2,t})$;
- λ_i error correction parameters that reflect idiosyncratic adjustment of $p_{i,t}$ to disparities in shocks of microstructural noise that causes cointegrated series to diverge.

By virtue of the permanent-transitory decomposition assumed by microstructure theory (see ass. 1, 2, 1', 2' and 3), no such error correction process can have any permanent effect on the common stochastic trend $\sum w_t$ and therefore no permanent effect on $p_{i,t}$ itself. In eq. 5, Gonzalo and Granger (1995) establish the trading prices as a random walk of information arrivals plus a covariance –stationary private value $s_{i,T}$ which does not Granger-cause the $\sum w_t$. Consequently, in eq. 8, Gonzalo and Granger (1995) identify the common factor weights (vector) λ_{\perp} , by imposing the error correction term $(p_{1,t} - p_{2,t} = s_{1,t} - s_{2,t})$ does not Granger-cause the common stochastic trend $\sum w_t$ in the long run. In other words, they impose the orthogonality condition $\lambda'_{\perp} \lambda = 0$. Since the vector of common factor weights λ_{\perp} is orthogonal to the vector of coefficients λ on the error correction term, in eq. 7 and 7' the λ_1 and λ_2 estimate provide a way to identify the permanent weights $\lambda_{\perp 1}$ and $\lambda_{\perp 2}$.

In this way, Gonzalo and Granger (1995) identify each market's contribution to the common factor, where the contribution is defined to be a function of the market's error correction coefficients, i.e. the speed of adjustment coefficients, λ_1 and λ_2 . In practice, PT price discovery measure is obtained by the orthogonal normalization of error correction term coefficients:

$$PT_1 = \lambda_{\perp 1} = \frac{\lambda_2}{\lambda_2 - \lambda_1}; \quad PT_2 = \lambda_{\perp 2} = \frac{\lambda_1}{\lambda_1 - \lambda_2};$$

Intuitively, the lower the estimated error correction parameter λ_i for a market i , the higher the common factor weight $\lambda_{\perp i}$ and the bigger the contribution of that market to the revelation of the permanent innovations underlying the common stochastic trend. When a market i dominates in terms of price discovery, its value of λ_i will be small, indicating that this market does not correct in response to any differences in prices between markets. Conversely, when a market adjusts as opposed to anticipate the other market, its value of λ_i will be large in absolute terms relative to the dominant market, indicating strong adjustment to price differences. This ratio gives an indication of the degree of dominance of one market over the other market. The PT measure is concerned with the permanent shocks that result in a disequilibrium as markets process news at different speeds.

2.2.2. Hasbrouck (1995) “information share” (IS)

Hasbrouck (1995) proposes an alternative measure for price discovery, the information share (IS). Hasbrouck’s model of “information share” assumes that price volatility reflects new information, and thus the market that contributes most to the variance of the innovations to the common factor is also presumed to contribute most to price discovery.

In practice, the IS measure weighs the error correction terms (λ_1 and λ_2) in eq. 7 and 7’ by the variances and covariance of each market’s VECM innovations. Due to the existence of a non-zero correlation between the VECM residuals in eq. 7 and 7’ ($\text{Cov}(u_{1,t}, u_{2,t}) \neq 0$), Hasbrouck (1995) suggests to use the Cholesky decomposition to remove this contemporaneous correlation. The Cholesky orthogonalization produces two bounds (a and b) of price discovery for each market (1 and 2) which are given by the following relations:

$$IS_{1a} = \frac{\lambda_2 \left(\sigma_1^2 - \frac{\sigma_{12}^2}{\sigma_2^2} \right)}{\lambda_2^2 \sigma_1^2 - 2\lambda_1 \lambda_2 \sigma_{12} + \lambda_1^2 \sigma_2^2}; \quad IS_{1b} = \frac{\left(\lambda_2 \sigma_1 - \lambda_1 \frac{\sigma_{12}}{\sigma_1} \right)^2}{\lambda_2^2 \sigma_1^2 - 2\lambda_1 \lambda_2 \sigma_{12} + \lambda_1^2 \sigma_2^2};$$

and

$$IS_{2a} = 1 - IS_{1b}; \quad IS_{2b} = 1 - IS_{1a};$$

where:

- IS_{ia} and IS_{ib} give the two bounds for market i;
- $\sigma_1^2, \sigma_{12}, \sigma_2^2$ represent elements of the variance-covariance matrix of residuals ($u_{1,t}, u_{2,t}$) in eq. 7 and 7’.

Baillie et al. (2002) argue that the average of these bounds

$$IS_1 = (IS_{1a} + IS_{1b})/2 \quad \text{and} \quad IS_2 = (IS_{2a} + IS_{2b})/2 = 1 - IS_1$$

provides a sensible estimate of the role of each market in the discovery of the efficient price.

2.2.3. Comparison between PT, IS and ILS (Information Leadership Share)

Market microstructure scholars have made substantial progress in reconciling and understanding the two measures PT and IS (e.g., Baillie et al., 2002; Lehmann, 2002). In order to measure the contribution of each market to price discovery, IS decomposes the variance of the common factor innovations. The IS measures the proportion of variance contributed by one market with respect to

the variance of the innovations in the common efficient price. PT decomposes the common factor itself, and, ignoring the correlation between the markets, attributes superior price discovery to the market that adjusts least to price movements in the other market. When price-change innovations are correlated, IS approach can only provide upper and lower bounds on the information shares of each market. Since neither method is considered universally superior, usually both are reported in literature. Even if, IS seems to have more economic intuition.

However, recently, the theoretical work of Yan and Zivot (2010) and the related empirical work of Putniņš (2013) shed further light on what each price discovery approach really measures. In order to correctly interpret the price discovery measures it is essential to distinguish between speed and noise in the process of price discovery. Both are implicit in Lehman's (2002) definition of price discovery provided earlier, "... efficient and timely incorporation of information ...". "Timely" refers to the relative speed with which a price series reflects new information about the fundamental value. "Efficient" implies a relative absence of noise, such as bid-ask bounce, temporary deviations due to imperfect liquidity and so on. Putniņš (2013) simulations show that when the price series differ in the levels of noise, IS and PT both measure a combination of leadership (relative speed) in impounding new information and relative avoidance of noise, to different extents. IS places more emphasis on informational leadership than avoidance of noise compared to PT and is therefore more closely related to the aspect of price discovery that researchers often seek to measure. Hence, IS approach tells us which price "moves first" regardless of how noisy it is contributes to achieving this objective.

Indeed, the vast majority of literature takes the view that a price series dominates price discovery if it is the first to adjust to new information about the fundamental value, and a price series makes a greater contribution to price discovery (is responsible for a larger "share" of price discovery) the more often it is the first to adjust to new information. This is different from measuring which price is the most informative about fundamental value because the price series that is the first to impound new information may also be considerably noisier than other price series potentially rendering it less useful as a measure of the fundamental value.

To address this issue a third price discovery measure "Information Leadership" (ILS) was introduced by Yan and Zivot (2010) and modified by Putniņš (2013) to be comparable with IS and PT. ILS combines IS and PT formulae in such a way that their dependence on the level of noise cancels out, and thus it provides a more robust measure of the contribution of a price series to impounding new information.

In detail, Yan and Zivot (2010) show that both the PT and IS measures capture each price series' level of noise and liquidity shocks in addition to its informational leadership in reflecting innovations in

the fundamental value. They show that combining the PT and IS metrics would eliminate the noise component and give a clean measure of price discovery which only reflects the informational leadership (IL) of a price series:

$$IL_1 = \left| \frac{IS_1}{IS_2} * \frac{PT_2}{PT_1} \right| \quad IL_2 = \left| \frac{IS_2}{1} * \frac{PT_1}{PT_2} \right|$$

IL has the range $[0, \infty)$, whereby values higher (lower) than 1 indicate that prices of market 1 lead (do not lead) the price discovery process. In order to facilitate both the interpretation of the IL metric and the comparison with the IS and PT measures, we follow Putniņš (2013) and define the information leadership share (ILS) of a price series as follows:

$$ILS_1 = \frac{IL_1}{IL_1 + IL_2} \quad ILS_2 = 1 - ILS_1$$

It is worth explaining that ILS is based on 3 key assumptions (i) there are only two price series; (ii) the structural model has only one permanent and one transitory shock; and (iii) the reduced form VECM errors are uncorrelated. Under these assumptions ILS measures informational leadership and is not affected by the relative level of noise in a price series. However, it is not clear how reliably ILS measures informational leadership when these assumptions are violated.

3. A TIME-VARYING PRICE DISCOVERY MODEL: EMPIRICAL ANALYSIS

Price discovery dynamics are not stable and tend to change substantially and rapidly over time. In order to show these features, most of the empirical literature used to split arbitrarily the dataset in more periods and, then, compare them.

This approach has mainly two limits:

1. reduce the robustness of results;
2. reduce the possibility of policy analysis.

As regard limit 1, we show how, just adding and removing one observation, price discovery dynamics vary substantially. This implies that choosing arbitrarily the dates for different periods of analysis arises doubts on the reliability of empirical evidences. We argue that this practice has limited applicability as the sub-periods need to be sufficiently long as to ensure robust estimates. This constrains the frequency with which time variations can be reliably measured. As regard limit 2, we argue that standard time-invariant models are not able to detect and, then, to analyse the sources of some empirical facts such as:

- a) structural changes in causal link between asset prices;
- b) changes in volatility and correlation between assets;
- c) changes between linear and non-linear regimes.

We propose a first extension of price discovery models by applying a moving window (MW) technique to PT, IS and ILS measures. The MW means that we move the start and end dates of the subsample forward by one day for each iteration of VECM, so the sample size is constant. Early observations are gradually replaced by more recent observations. This approach provides a view of the time-varying contributions of each market to price discovery.

We apply this time-varying price discovery model to a topic greatly investigated in literature: the impact of “Naked CDS Ban” on the price discovery of sovereign credit risk in Eurozone within two interconnected financial markets: the Bond and the Credit Default Swap (CDS) markets.

The aim is to show how the extension of standard price discovery models towards a time-varying framework allows gaining more informative content for policy analysis. We apply MW to the VECM and show time-varying empirical results of PT, IS and ILS measures.

3.1. The impact of naked CDS ban on the price discovery of sovereign credit risk in Eurozone

Sub-prime and sovereign debt crises have raised concerns regarding the use of CDSs. CDSs are a derivative financial product used to hedge against the default risk of any entity. It has been suspected that the crises have been exacerbated by a few investors driving up the prices in the CDS market. This issue has attracted much interest in policy circles and has led to European authorities to ban in some periods covered, or “naked,” purchases of Sovereign CDS (SCDS) protection referencing European Economic Area sovereign debt obligations, that is, banning purchases in which there is no off setting position in the underlying debt. The prohibition is based on the view that, in extreme market conditions, such short selling could push sovereign bond prices into a downward spiral, which would lead to disorderly markets and systemic risks, and hence sharply raise the issuance costs of the underlying sovereigns.

Consequently, a current debate among academics and financial regulators is if the EU’s ban purchases of naked SCDS protection, into effect on November 1, 2012, was necessary and if it is moving in the correct direction. The International Monetary Fund (IMF) in the Global Financial Stability Report (April 2013, Ch. 2) writes: “The empirical results presented in this chapter do not support many of the negative perceptions about SCDS. Such bans may reduce SCDS market liquidity to the point that

these instruments are less effective as hedges and less useful as indicators of market-implied credit risk.” In line with the work of Duffie (2010) where he argues against a potential EU’s ban: “regulations that severely restrict speculation in credit default swap markets could have the unintended consequences of reducing market liquidity, raising trading execution costs for investors who are not speculating, and lowering the quality of information provided by credit default swap rates regarding the credit qualities of sovereign issuers”. From another perspective to approach this debate, Palladini and Portes (2011) and Delatte et al. (2012) document an adverse influence of the SCDS on the underlying bond pricing when bearish investors use these instruments to express their views on the sovereign credit. Recently, Delatte, Fouquau and Portes (2016) document the amplifying effects of financial sector CDS on the market’s evaluation of sovereign risk. Their findings suggest the inconsistency of the EU’ s ban to not include also financial CDSs which drive market sentiment and exacerbated shocks to economic fundamentals of Peripheral European countries.

We know much less of the effects of naked SCDS ban on the efficiency of price discovery dynamics of sovereign credit risk in Eurozone. In this contest, we address the concern of the effects of SCDS financial regulation on the dynamic relationship between sovereign CDS spreads and Bond spreads. To shed light, on how the price discovery dynamics change in response to policy announcements, a natural candidate is a time varying model. In this sense, our first contribution is to propose an extension of standard price discovery models towards a time-varying framework to conduct policy analysis. Our second contribution is to disentangle the two following effects: a) which market is the first to adjust to new information, and b) which market is the most informative about fundamental value. Effect a) is measured by IS metric and effect b) by ILS metric.

Recalling the theoretical model of section 2, the fundamental value or permanent component in (eq.1) is the unobservable efficient price of credit risk that is shared between CDS and bond spreads. We model the observed spreads in CDS and Bond markets at each time t equal to this efficient price plus a component containing microstructural noise $s_{CDS,t}$ and $s_{BOND,t}$ exactly as in eq. 2 and 2’. Then, we specify eq.7 and eq.7’ in the context of this empirical study as it follows:

$$\Delta CDS_t = \sum_{h=1}^l \beta_{CDS,t-h} \Delta p_{CDS,t-h} + \sum_{h=1}^l \gamma_{CDS,t-h} \Delta p_{BS,t-h} + \lambda_{CDS} Z_{t-1} + c_{CDS} + u_{CDS,t}; \quad (9)$$

$$\Delta BS_t = \sum_{h=1}^l \beta_{BS,t-h} \Delta p_{CDS,t-h} + \sum_{h=1}^l \gamma_{BS,t-h} \Delta p_{BS,t-h} + \lambda_{BS} Z_{t-1} + c_{BS} + u_{BS,t}; \quad (9')$$

where BS stands for bond spreads which are computed as the difference between a bond yield and interest rate swap (proxy for the risk-free rate) on the same maturity. In summary, we substitute market 1 and 2 in the theoretical model in section 2 with CDS and Bond markets, respectively. All interpretations remain equal except that, following Blanco et al. (2005), we specify the co-integration

vector in the error correction term Z_{t-1} as $[1,-1,a]$ instead of $[1,-1]$. The constant in the co-integrating space, “a”, should be equal zero, but since we know that our proxy for the risk-free rate is imperfect and that the equality between CDS and Bond spreads does not hold strictly owing to technical and market factors (Choudhry, 2006), we do not impose this condition.¹⁷

We apply the VECM specified in eq. 9 and 9’ for the Eurozone countries of Belgium, France, Germany (Core countries) and Italy, Portugal, Spain (Peripheral Countries). Data on 5 years Bond yields and interest rate swaps are downloaded from Bloomberg. Data on 5 years CDS are downloaded from Markit. The exact dates of regulations and monetary policies announcements are collected from ECB and IMF reports. Fig 1 shows the evolution in basis points of CDS and Bond spreads for each country from January 1, 2006 to December 31, 2014. In each graph, we detect four regimes (delimited by three vertical lines): the pre-crisis period, the financial crisis period, the sovereign debt crisis period and the reduction of sovereign credit spreads period. Table 1 reports summary statistics for the whole sample period and for each sub-period.

Unit root tests (table 2) and Cointegration tests (table 3) confirm for all countries the presence of a common component in the long run between CDS spreads and Bond spreads. Hence, we first apply the VECM in eq. 9 and 9’ in a time invariant framework, as usually presented in related empirical papers. Table 4 reports estimates of error correction coefficients (λ_1 and λ_2), their t-statistics and the permanent-transitory (PT), information share (IS) and information leadership share (ILS) measures of price discovery of CDS over the whole sample period (Panel A), the pre-crisis period (Panel B), the financial crisis period (Panel C) the sovereign debt crisis period (Panel D) and the reduction of sovereign credit spreads period (Panel E). For the IS measure, we report its lower bound (LB), upper bound (UB) and the midpoint (MID). Note, IS and ILS are bounded between 0 and 1, so they are easily comparable and interpreted as a percentage share of contribution to price discovery. While PT assume values above 1 or below 0 when the adjustment coefficients are not in line with theory, i.e. λ_1 negative and λ_2 positive. This makes interpretation of PT measure not always simple.

Then, we apply moving window technique to eq. 9 and 9’ to get time varying results. Precisely, we run moving window of 500 observations starting from January 1 2006 and at each iteration we move ahead the start and the end date of one day until the end of the sample on December 31 2014. We get 1787 moving windows, and then 1787 time varying estimates of PT , IS and ILS.

¹⁷ Here we don’t go in further details, because is not aim of the paper to shed light on the literature on the theoretical arbitrage relation of the CDS-Bond basis. On this topic see Giorgione et al. (2016).

Time-varying empirical results are organized in the following way. Fig. 2 shows the time varying results estimates of error correction terms, λ_1 and λ_2 , as monthly average of the 1787 moving window of 500 observations for each country. Fig. 3 aggregates these estimates in two groups: Core (Belgium, France and Germany) and Peripheral (Italy, Portugal and Spain). Fig. 4 aggregates the time varying estimates of IS and ILS metrics, as monthly average of the 1787 moving window of 500 observations for each country, in two groups: Core (Belgium, France and Germany) and Peripheral (Italy, Portugal and Spain). We aggregate the results in Core and Peripheral groups because single countries present similar patterns according to the group from which they belong. Hence, we proceed to analyse aggregate results and when necessary we skip to results for single country.

We want to shed light of the effects of CDS financial regulation policies on the price discovery process of sovereign credit risk. Precisely, the permanent EU' ban on naked short selling was announced on March 24, 2012, went into effect on November 1, 2012 and it followed a similar but temporary ban in Germany from May 19, 2010 to March 31, 2011. Moreover, consider that on June 14, 2010 started the EU commission consultation on short selling. On September 15, 2010, European Commission short selling regulation was proposed about banning naked short sales and SCDS protection sales. And on November 15, 2011 the European Parliament adopted short selling regulation. So that, in Fig. 3 and 4, we distinguish the CDS financial regulation in two waves:

- the first one (darker shaded area) consider all the moving windows which have observations on the dates of the Germany's Ban and EU proposal
- the second one (lighter shaded area) consider all the moving windows which have observation on the dates of the EU's Ban announcement and entered into force. Note the lighter shaded area cover the final part of the darker shaded area.

During the sample period, other important events reasonably affect the dynamics between Bond spreads and CDS spreads. With the attempt to have a broader picture, we report, in fig. 3 and 4, crucial crisis events and monetary policies events. Solid lines represent the month in which events related to financial crisis and to sovereign debt crisis happen. Precisely, they appear in the following order: Bear Sterns collapse march 2008 (BS), Lehman Brothers default September 2008 (LB), Greece aid request April 2010 (GRE), Ireland aid request October 2010 (IRE) and Portugal aid request April 2011 (POR). Dashed lines represent the month in which monetary policies were announced. Precisely, they appear in the following order: announcement of Long Term Refinancing Operations (LTRO) with maturity one year (June, 2009), announcement of Long Term Refinancing Operations

(LTRO) with maturity three years (December 2011), Outright Monetary Transactions (OMT) announcement (August, 2012).

First, we discuss time-varying estimates of causal relations between CDS and Bond spreads in fig.3. Subsequently, we discuss time-varying estimates of IS and ILS measures in fig.4.

In fig.3, keeping in mind the PT model, we can look at the dynamics of CDS spread adjustments (λ_1) and Bond spread adjustments (λ_2) and highlight which market is the leader (dominant market) and which market adjust more (lag market) to restore the long term equilibrium and correct deviations from the common efficient price.

Aggregate results on Core countries (first graph of fig.3) show that with the beginning of the sub-prime crisis the CDS is the dominant market in line with literature that shows the higher price discovery contribution of CDS during market distress. As the estimates start to incorporate information on CDS Ban in Germany and CDS Ban proposal in EU, there is at the same time a reduction of the feedback from Bond spreads to CDS spread and an increase of CDS adjustments. This suggests a leadership of Bond market. These effects are obviously related also to the 1st LTRO program.

Information on Greek aid request (i.e. the beginning of the sovereign debt crises) push up again the “follower” role of the Bond spreads but not as much as in the previous crisis. Moreover, respect to the sub-prime crises, as the moving windows move ahead and incorporate more information on the application of the EU CDS Ban, the CDS and Bond spreads adjustments behave more related to theory. It’s difficult to say how much this effect is also due to accommodating monetary policies announcement. Challenging are the results in correspondence of OMT which show an increase of CDS leadership due to both a sharp increase of Bond spreads adjustments and reduction of CDS adjustments.

Aggregate results on Peripheral countries (second graph of fig.3) show only partially similar dynamics. Respect to Core group, the reduction of Bond adjustments is sharper in the first shaded area and, at the same time, is stronger during the sovereign debt crises the follower role of the Bond market. These results suggest higher leadership of CDS for Peripheral respect to Core countries during market distress. Again, this result is in line with previous literature. This evidence is also confirmed by looking at the Peripheral CDS adjustments which never seem generate sufficient feedback effects except, moderately, in the first shaded area.

To understand more about the price discovery process is interesting to compare IS and ILS measures. Remember, IS respect to PT and, then, respect to only looking at the adjustment coefficients, takes

into account differences in volatility between the two credit spreads. IS weights adjustment coefficients by the variances and covariance of each market's VECM innovations and thus the market that contributes most to the variance of the innovations to the common factor is also presumed to contribute most to price discovery. However, recently it has discovered that IS approach tells us which price "moves first" regardless of how noisy it contributes to achieving this objective. Hence, is interesting to compare this measure with ILS. ILS formula mixes IS and PT (which gives more emphasis to the less noise market) in order to eliminate the noise effect and have a cleaner measure of the market that is the most informative about the fundamental value regardless of who moves first.

Fig.4, as regard only IS, shows similar dynamics over time of Core (first graph in fig.4) and Peripheral (second graph in fig.4) CDS contribution to price discovery. During the subprime crisis the IS (blue bars) of CDS is high between 80% and 100%, it falls below 60% in the darker shaded area and with the outset of the sovereign debt crises it recovers between 60% and 70%, not at the levels of the financial crises. For Core group the recovering is weaker and again challenging is the behaviour in the last part of the graph around the OMT announcement, where estimates show a sharp increase CDS IS.

Interestingly, is the comparison of IS (blue bars) with ILS (red bars). Until the first LTRO announcement the Core and Peripheral ILS measures are below IS suggesting that noise effects during financial crisis overestimated the contribution to the fundamental value using IS. Same dynamics also in the second part of the darker shaded, where for Core and Peripheral ILS rises slightly above IS. Completely different patterns between the two group of countries in the lighter shaded area. For Core group, ILS remains stable a bit above IS except after OMT announcement. While for Peripheral group there is a stable wide gap, with ILS below IS.

These evidences clearly show the limitations of previous literature in estimating price discovery in a time invariant framework. Our approach is much more informative on the dramatic changes over time of price discovery dynamics, that we rationalize in the following paragraph to derive policy implications.

3.1.1. Policy Implications

In order to gain more economic intuition and policy implications, first we describe the features of financial instability on price discovery during the sub-prime crisis, that is, before the CDS financial regulation. Subsequently, we compare them to the features of financial instability on price discovery during the sovereign debt crisis, which covers all the period of CDS financial regulation.

During the sub-prime crisis, we document two features of financial instability expressed in the price discovery dynamics:

- 1) dominant role of CDS market for both Core and Peripheral groups. This is represented in fig. 3, by the weak CDS adjustments and wide Bond spreads feedbacks. And in fig 4, by the IS close to 100%. This is completely in line with previous literature which shows higher CDS contribution on price discovery during market distress;
- 2) a lower level of ILS respect to IS for Core and Peripheral CDS. We are the first to document that during financial crisis the strong contribution to price discovery of the CDS market is reduced consistently if we limit our attention to the ability of the derivative market to be informative on the unobservable efficient price (fundamental value) rather than to look only on his ability to move first. This is in line with expectation of higher microstructural noise during market turbulence. Hence, one could think to higher liquidity costs as well as other reasons could be that at least one market is pricing risk differently or pricing other factors or containing measurement errors.¹⁸

During the sovereign debt crisis, the price discovery dynamics of Core and Peripheral groups respond to financial instability differently:

- 1) for Core group, contrary to the previous crisis, CDS leadership role is reduced. In fig. 3, this is represented by a decrease of Bond adjustments and by an increase of CDS adjustments. In fig 4, this is represented by the reduction of IS around 60% and after the 2nd LTRO announcement CDS lost his leadership in favour of the Bond market. But with OMT announcement, again the CDS recovers its leadership. More important, respect to the previous crisis the gap between IS and ILS is reduced and inverted.
- 2) for the Peripheral group, respect to the previous crisis, fig 3 shows an increase of CDS leadership due to the increase of Bond adjustments. And, fig. 4 shows a reduction of CDS IS around 60% respect to the higher value of previous crises. But more important, the gap between IS and ILS is even higher than in the previous crisis.

In summary, we could argue that for the Core group there is a positive effect as the MW estimates incorporate information on CDS financial regulation and at the same time lose information on

¹⁸ Consider also that, CDS prices may contain a CTD option related to restructuring. And Bond spreads, as measured, ignore the repo cost of the bond. If these aspects are significant and not purely a short-term phenomenon, we might expect they affect cointegration relation, but also the price discovery measures.

subprime crisis. This effect is reflected in a more efficient price discovery process during the sovereign debt crisis respect to the previous crisis:

1. reduction of excessive CDS informative power during a period of instability;
2. reduction of the difference between IS and ILS during a period of instability.

Overall, these evidences support the argument of a reduction of microstructural noise in the sovereign credit risk market for Core countries with the introduction of CDS regulation.

For the Peripheral group only the reduction of point 1 is satisfied. Instead, in comparison to the previous crisis, there is a wider gap between the ability of CDS market “to move first” and the ability of CDS market to be informative about the unobservable efficient price of credit risk shared with the bond market. This evidence is not temporary but consistent for the whole sovereign debt crisis and CDS EU Ban. This suggests that at least one market (CDS or Bond) is pricing risk differently or pricing other factors. While if it was a temporary effect one could think more on measurement errors.

This evidence could reconcile with the findings of Delatte, Fouquau and Portes (2016). The inefficiency of the EU’s ban was to not include also financial CDSs, which drive market sentiment and exacerbated shocks to economic fundamentals of Peripheral European countries. This in turn could have created distortive effects between government CDS and Bond spreads dynamics only for Peripheral countries as market participants used the financial CDS of these countries to express their view and take positions on the Peripheral sovereign credit risk. Investors have exposures in their portfolio that they need to hedge, so when you restrict one instrument, they are going to look for alternatives. Investors bought protection on European banks on the basis that banks and sovereigns are so intimately linked that any increased risk of a sovereign default will increase the value of a bank CDS in a similar way to a sovereign CDS. This could be one of main reasons for which increased during the sovereign debt crises the microstructural noise and explains our evidence of a wide gap between ILS and IS for Peripheral countries.

Moreover, this interpretation suggests that, for Peripheral countries, sovereign bond spreads became more related to financial CDS. This interpretation may explain our evidence during the Eurozone crises of a higher ILS for the Bond market (around 60%) respect to the CDS market (around 40%). As effect of the sovereign CDS EU ban, fundamentals on Peripherals have been exacerbated by financial CDS and in turns i) the increase of sovereign Bond spreads, and ii) and less liquidity on the

sovereign CDS market, led to the higher informative power of Bond market on the fundamental value of Peripheral credit risk¹⁹.

Our empirical evidences and interpretations suggest, to regulators, to analyse carefully the micro-effects of their macro-oriented policies before implement them. In this case, for micro-effect we intend the change of behaviour of agents in response to a policy. This change could go in a direction that generates distortive effects respect to policy expectations.

It is worth noting the limitations of our empirical evidences. First, it is not trivial to disentangle the effects of monetary policy announcements from the effects of CDS regulation. Second, with moving windows we capture the progressive changes of price discovery dynamics as old information is lost and new information is incorporated. Moreover, is it reasonable to expect that in periods of financial instability the markets relations are exacerbated and could be better modelled with a non-linear model.

¹⁹ Note that when CDS (or Bond) market have the higher informative power to the fundamental value (measured by ILS) that does not mean that credit risk as priced by either the CDS or the bond market reflects correctly “fundamentals”.

4. TABLES²⁰

Table 1: Summary Statistics.										
This table reports for each Core (Belgium, France, Germany) and Peripheral (Italy, Portugal, Spain) Country summary statistics in basis points for the time series of CDS spreads and Bond spreads over the whole sample period (Panel A: 2290 observations for each country), the pre-crisis period (Panel B: 511 observations for each country), the financial crisis period (Panel C: 521 observations for each country) the sovereign debt crisis period (Panel D: 691 observations for each country) and the reduction of sovereign credit spreads period (Panel E: 567 observations for each country).										
	Mean		Median		Minimum		Maximum		St. Deviation	
	CDS	Bond	CDS	Bond	CDS	Bond	CDS	Bond	CDS	Bond
Panel A: Whole sample (January 2006-December 2014)										
Belgium	62.8	16.6	38.1	2.9	1.9	-68.5	339.3	331.6	64.1	49.1
France	42.0	-15.5	35.5	-16.4	1.5	-77.3	198.7	84.9	37.1	18.2
Germany	21.9	-45.6	16.0	-42.5	1.3	-107.0	91.4	-14.1	18.6	20.3
Italy	133.5	98.4	109.8	60.6	5.3	-38.5	501.5	574.4	117.0	128.5
Portugal	282.4	289.0	146.7	122.1	3.9	-61.6	1554	2015.5	336.6	404
Spain	129.1	99.3	97.7	49.6	2.3	-68.5	504.1	645.1	115.1	135.8
Mean Core	42.2	-14.8							39.9	29.2
Mean Peripheral	181.7	162.2							189.6	222.8
Mean Eurozone	111.9	73.7							114.7	126.0
Panel B: Pre-crisis period (January 2006-December 2007)										
Belgium	3.1	-21.4	2.6	-20.4	1.9	-42.0	13.2	-9.8	2.0	6.3
France	2.3	-22.0	2.2	-19.1	1.5	-47.0	7.5	-12.5	1.1	7.1
Germany	2.2	-24.8	2.0	-21.1	1.3	-57.3	6.2	-14.1	0.9	9.5
Italy	9.7	-13.1	9.4	-10.7	5.3	-33.8	21.2	0.2	2.9	6.8
Portugal	7.0	-15.6	7.1	-13.7	3.9	-40.7	19.4	-3.2	2.7	8.2
Spain	4.4	-21.1	3.0	-18.9	2.3	-45.6	19.7	9.9	3.4	7.7
Mean Core	2.5	-22.7							1.3	7.6
Mean Peripheral	7.0	-16.6							3.0	7.6
Mean Eurozone	4.8	-19.7							2.2	7.6
Panel C: Financial crisis period (January 2008-December 2009)										
Belgium	48.4	3.0	37.4	1.6	12.5	-68.5	155.5	84.1	33.1	36.2
France	29.8	-30.0	25.1	-31.4	6.1	-77.3	97.9	3.4	21.0	16.6
Germany	25.6	-52.2	22.3	-49.4	5.0	-107.0	91.4	-14.1	19.8	17.9
Italy	83.3	12.7	73.5	2.7	20.5	-38.5	197.8	86.4	48.4	30.9
Portugal	63.3	6.5	57.3	2.7	17.4	-61.6	158.3	98.1	30.3	34.7
Spain	69.9	-8.1	68.4	-4.0	19.7	-68.5	167.8	54.5	34.7	27.6
Mean Core	34.6	-26.4							24.6	23.6
Mean Peripheral	72.2	3.7							37.8	31.1
Mean Eurozone	53.4	-11.3							31.2	27.3
Panel D: Sovereign debt crisis period (January 2010-August 2012)										
Belgium	142.4	71.5	126.1	63.3	46.0	-2.7	339.3	331.6	57.0	50.2
France	85.8	-7.6	78.9	-14.6	36.0	-46.0	198.7	84.9	32.4	22.2
Germany	40.3	-61.7	39.2	-60.1	17.4	-101.5	80.5	-21.9	13.1	19.6
Italy	243.9	201.3	184.0	121.3	78.7	23.6	501.5	574.4	124.6	147.2
Portugal	661.8	718.2	605.1	785.9	111.1	56.4	1554.0	2015.5	367.4	467.2
Spain	257.6	214.4	252.7	200.4	92.6	19.0	504.1	645.1	92.0	122.7
Mean Core	89.5	0.7							34.2	30.7
Mean Peripheral	387.8	378.0							194.7	245.7
Mean Eurozone	238.6	189.3							114.4	138.2
Panel E: Reduction of sovereign credit spreads period (September 2012-December 2014)										
Belgium	35.9	-1.0	35.3	-1.1	21.2	-26.2	77.1	39.8	8.8	13.2
France	37.4	-5.3	35.7	-5.2	23.3	-19.5	63.1	14.9	7.5	6.3
Germany	14.1	-39.6	13.1	-38.7	7.9	-60.3	26.3	-25.0	4.1	7.8
Italy	161.7	159.0	162.2	158.9	68.4	49.1	340.3	343.8	59.0	76.3
Portugal	289.4	324.5	313.5	364.7	117.7	101.7	578.8	688.0	117.3	148.5
Spain	144.6	173.3	128.2	163.7	44.1	27.5	349.0	393.5	73.3	104.8
Mean Core	29.1	-15.3							6.8	9.1
Mean Peripheral	198.6	218.9							83.2	109.9
Mean Eurozone	113.8	101.8							45.0	59.5

²⁰ In all tables, ***, **, * denote rejection at the 1%, 5% and 10% levels, respectively, of the null hypothesis.

Table 2: Augmented Dickey-Fuller (ADF) and Kwiatkowsky-Phillips-Schmidt-Shin (KPSS) Unit Root Tests.
 ADF and KPSS tests on levels and first differences of CDS and Bond spreads series for Core Countries (Belgium, France, Germany) and Peripheral Countries (Italy, Portugal, Spain). All series are daily for the period from January 1, 2006 to December 31, 2014.

	ADF Test				KPSS Test			
	CDS		Bond		CDS		Bond	
	Level	Difference	Level	Difference	Level	Difference	Level	Difference
Belgium	-1.9	-9.7***	-1.6	-11.1***	1.6***	0.1	1.5***	0.1
France	-1.9	-11.6***	-3.0**	-13.5***	2.3***	0.1	2.3***	0.0
Germany	-2.4	-9.6***	-2.6*	-13.9***	1.4***	0.1	1.1***	0.1
Italy	-1.8	-8.7***	-1.6	-9.2***	2.9***	0.1	3.1***	0.1
Portugal	-1.1	-11.8***	-1.3	-16.2***	2.4***	0.1	2.4***	0.1
Spain	-1.7	-9.0***	-1.3	-9.6***	2.9***	0.2	3.4***	0.1

Table 3: Johansen Cointegration Tests.
 This table reports Johansen Trace test statistic for the null hypothesis of none cointegrating vector $[1, -1, c]$ between CDS spreads and Bond spreads for each Core Country (Belgium, France, Germany) and Peripheral Country (Italy, Portugal, Spain). The number of lags is determined by the Akaike information criterion (AIC). A constant is included in both the cointegrated equation and VECM. All series are daily for the period from January 1, 2006 to December 31, 2014.

Countries	Trace test: Null hypothesis of none cointegrating vector
Belgium	35.1***
France	36.0***
Germany	25.7***
Italy	28.7***
Portugal	49.4***
Spain	21.7***

Table 4: Time-invariant Price discovery of credit risk in Eurozone for Core Countries (Belgium, France, Germany) and Peripheral Countries (Italy, Portugal, Spain).							
This table reports error correction coefficients (λ_1 and λ_2) estimates, their t-statistics and the permanent-transitory (PT), information share (IS) and information leadership share (ILS) measures of price discovery over the whole sample period (Panel A), the pre-crisis period (Panel B), the financial crisis period (Panel C) the sovereign debt crisis period (Panel D) and the reduction of sovereign credit spreads period (Panel E). For the IS measure, we report its lower bound (LB), upper bound (UB) and the midpoint (MID).							
	λ_1	λ_2	PT_{CDS}	IS_{CDS}			ILS_{CDS}
				IS_{LB}	IS_{UB}	IS_{MID}	
Panel A: Whole sample (January 2006-December 2014)							
Belgium	-0,01*	0,01**	0,56	0,34	0,82	0,58	0,54
France	-0,00	0,01***	0,72	0,78	0,91	0,85	0,82
Germany	-0,00	0,00***	0,81	0,92	0,97	0,94	0,94
Italy	0,00	0,02***	1,05	0,54	1,00	0,77	0,02
Portugal	-0,01	0,01**	0,66	0,67	0,73	0,70	0,59
Spain	0,00	0,01***	1,63	0,84	0,84	0,84	0,80
Mean Core						0,79	0,77
Mean Peripheral						0,77	0,47
Mean Eurozone						0,78	0,62
Panel B: Pre-crisis period (January 2006-December 2007)							
Belgium	0,01***	0,01	2,84	0,04	0,09	0,07	0,00
France	0,00***	0,01	2,12	0,01	0,02	0,01	0,00
Germany	0,00***	0,01	1,66	0,03	0,09	0,06	0,00
Italy	0,01***	0,01	-6,09	0,00	0,02	0,01	0,00
Portugal	0,01***	0,00	-1,99	0,02	0,03	0,03	0,00
Spain	0,01***	0,00	-3,09	0,02	0,38	0,20	0,10
Mean Core						0,05	0,00
Mean Peripheral						0,08	0,03
Mean Eurozone						0,06	0,02
Panel C: Financial crisis period (January 2008-December 2009)							
Belgium	0,00	0,02***	1,11	0,98	1,00	0,99	0,99
France	0,00	0,03***	1,15	0,94	0,94	0,94	0,79
Germany	0,00	0,01**	1,02	0,64	0,99	0,81	0,01
Italy	-0,00	0,02***	0,85	-0,14	1,73	0,80	0,34
Portugal	0,00	0,04***	1,07	0,99	1,00	0,99	0,99
Spain	-0,00	0,04***	0,90	0,53	0,90	0,72	0,07
Mean Core						0,91	0,59
Mean Peripheral						0,84	0,47
Mean Eurozone						0,87	0,53
Panel D: Sovereign debt crisis period (January 2010-August 2012)							
Belgium	-0,03**	0,01	0,25	0,13	0,14	0,14	0,18
France	-0,02*	0,01**	0,45	0,44	0,68	0,56	0,70
Germany	-0,01	0,01*	0,51	0,54	0,70	0,62	0,71
Italy	-0,01	0,02*	0,78	0,30	0,97	0,63	0,19
Portugal	-0,01	0,01	0,63	0,25	0,88	0,57	0,36
Spain	0,00	0,02*	1,09	1,00	0,99	1,00	1,00
Mean Core						0,44	0,53
Mean Peripheral						0,73	0,52
Mean Eurozone						0,58	0,52
Panel E: Reduction of sovereign credit spreads period (September 2012-December 2014)							
Belgium	-0,01	0,02*	0,67	0,54	0,72	0,63	0,42
France	-0,03***	0,03***	0,53	0,41	0,53	0,47	0,37
Germany	-0,01	0,02***	0,78	0,84	0,75	0,79	0,56
Italy	0,00	0,01*	1,03	0,42	1,00	0,71	0,01
Portugal	0,00	0,03	1,03	1,00	1,00	1,00	1,00
Spain	0,00	0,01	1,57	0,80	0,85	0,83	0,74
Mean Core						0,63	0,45
Mean Peripheral						0,84	0,58
Mean Eurozone						0,74	0,52

5. FIGURES

Figure 1: Time series of CDS and Bond spreads.

This figure shows the evolution in basis points of CDS and Bond spreads for each Core Country (Belgium, France, Germany) and Peripheral Country (Italy, Portugal, Spain) from January 1, 2006 to December 31, 2014. In each graph, the pre-crisis period, the financial crisis period, the sovereign debt crisis period and the reduction of sovereign credit spreads period are delimited by three vertical lines.

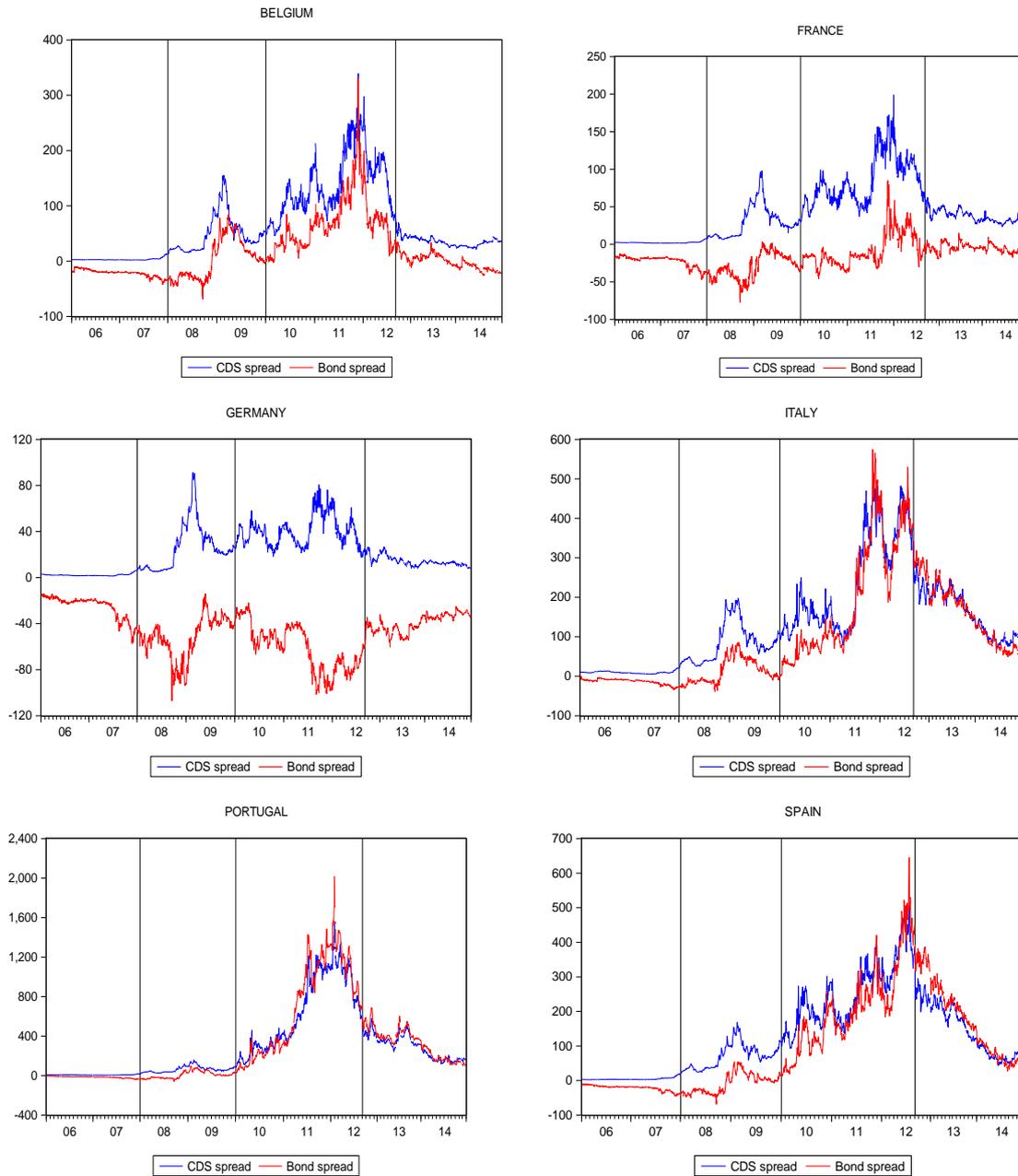


Figure 2: Time-varying causal relations between CDS spreads and Bond spreads for each country

This figure shows the time-varying estimates of lambda 1 (blue line) and lambda 2 (red line) as monthly average of the 1787 moving windows each of 500 observations for the Core countries (Belgium, France, Germany) and Peripheral countries (Italy, Portugal, Spain) from January 1, 2006 to December 31, 2014.

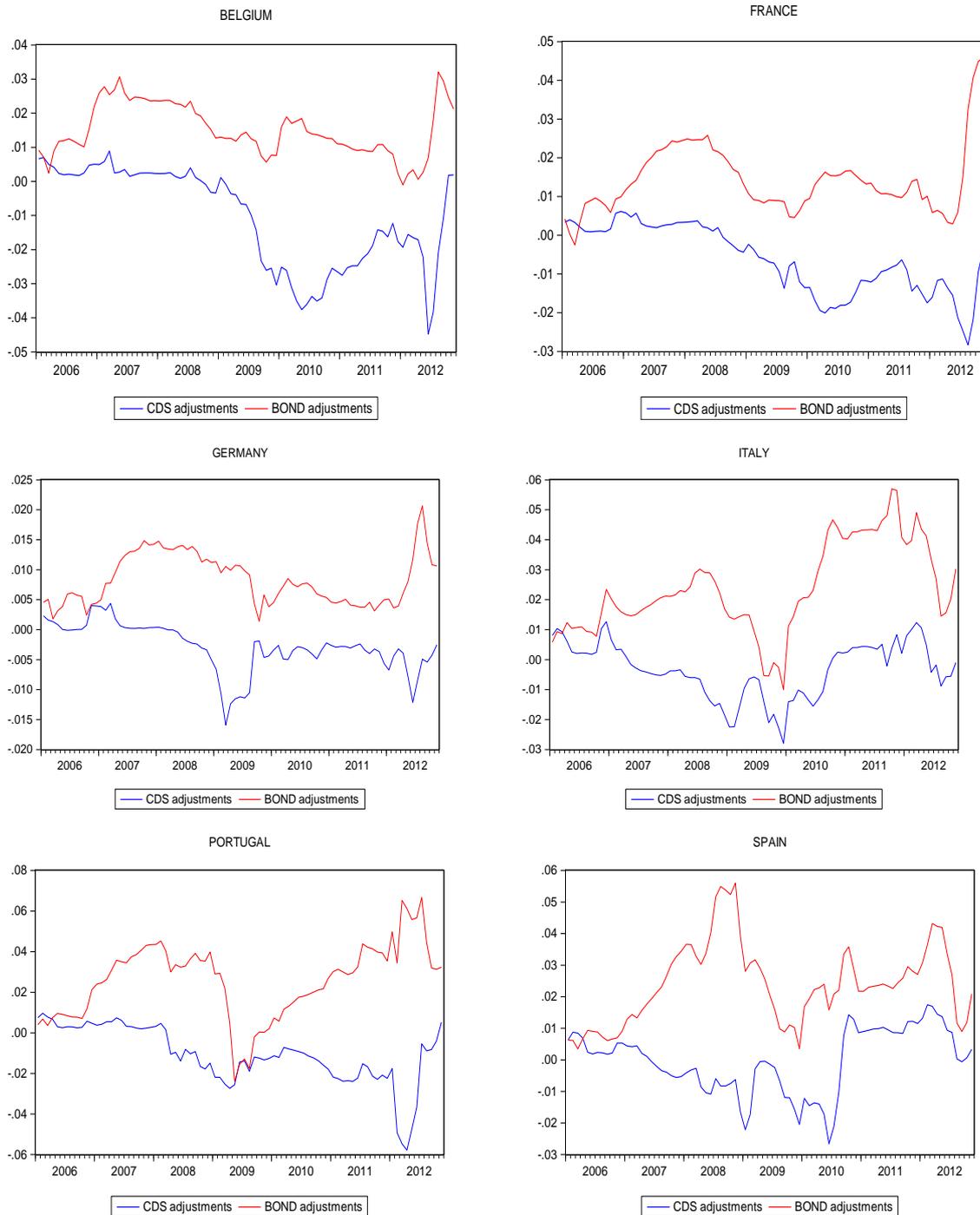


Figure 3: Time-varying causal relations between Core and Peripheral CDS spreads and Bond spreads

This figure shows the estimates of lambda 1 (blue line) and lambda 2 (red line) as monthly average of the 1787 moving windows each of 500 observations for the aggregate Core group and Peripheral group from January 1, 2006 to December 31, 2014. Darker shaded area considers all the moving windows with observations on the dates of the Germany’s Ban and EU proposal. Lighter shaded area considers all the moving windows with observations on the dates of EU’s Ban announcement and went into effect. The latter shaded area covers the final part of the former shaded area. Solid lines represent financial crisis and to sovereign debt crisis events: Bearn Sterns (BS) collapse (March 2008), Lehman Brothers (LB) default (September 2008), Greece (GRE) aid request (April 2010), Ireland (IRE) aid request (October 2010) and Portugal (POR) aid request (April 2011). Dashed lines represent monetary policies announcements: Long Term Refinancing Operations (LTRO) with maturity one year (June, 2009), Long Term Refinancing Operations (LTRO) with maturity three years (December 2011), Outright Monetary Transactions (OMT) announcement (August, 2012).

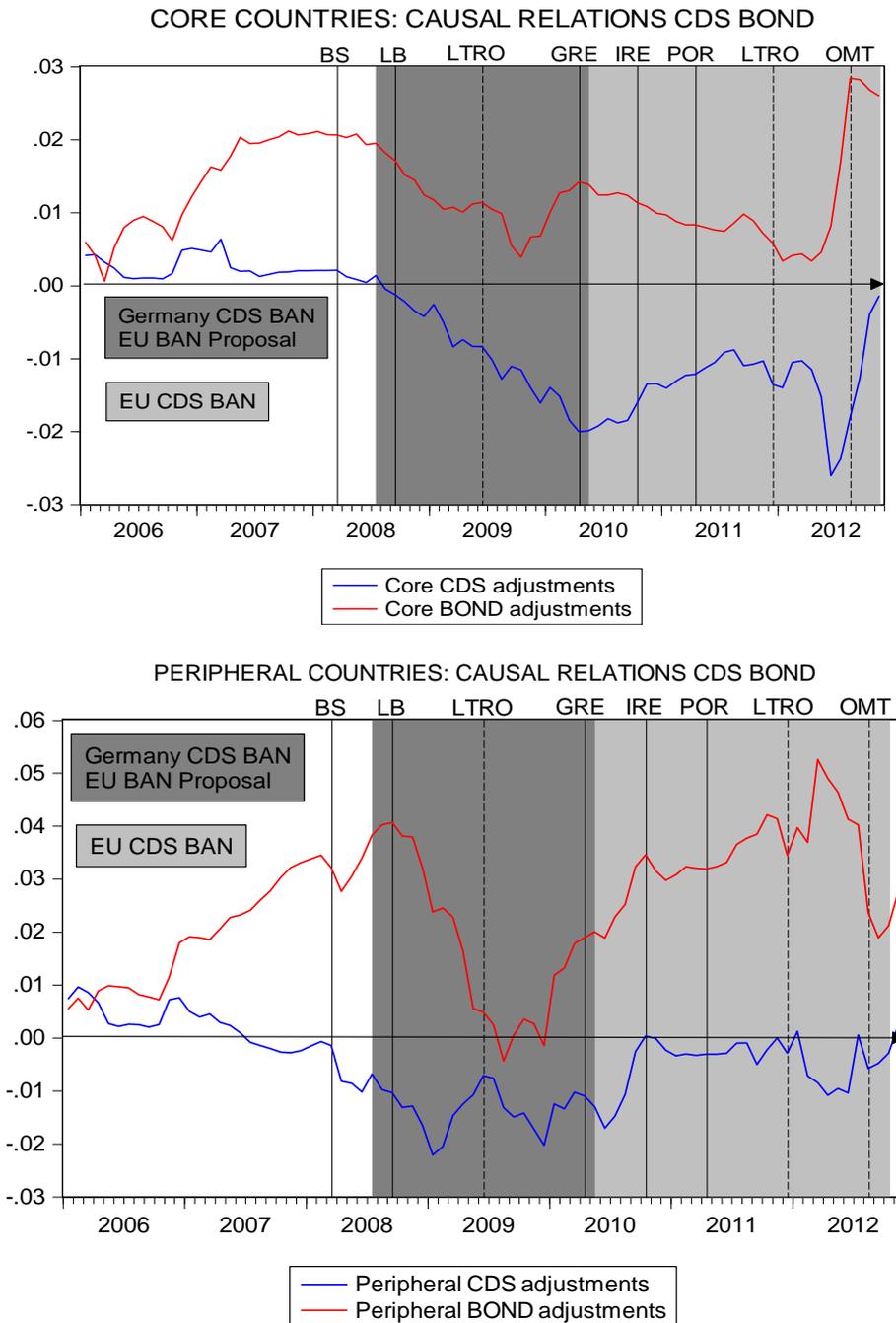
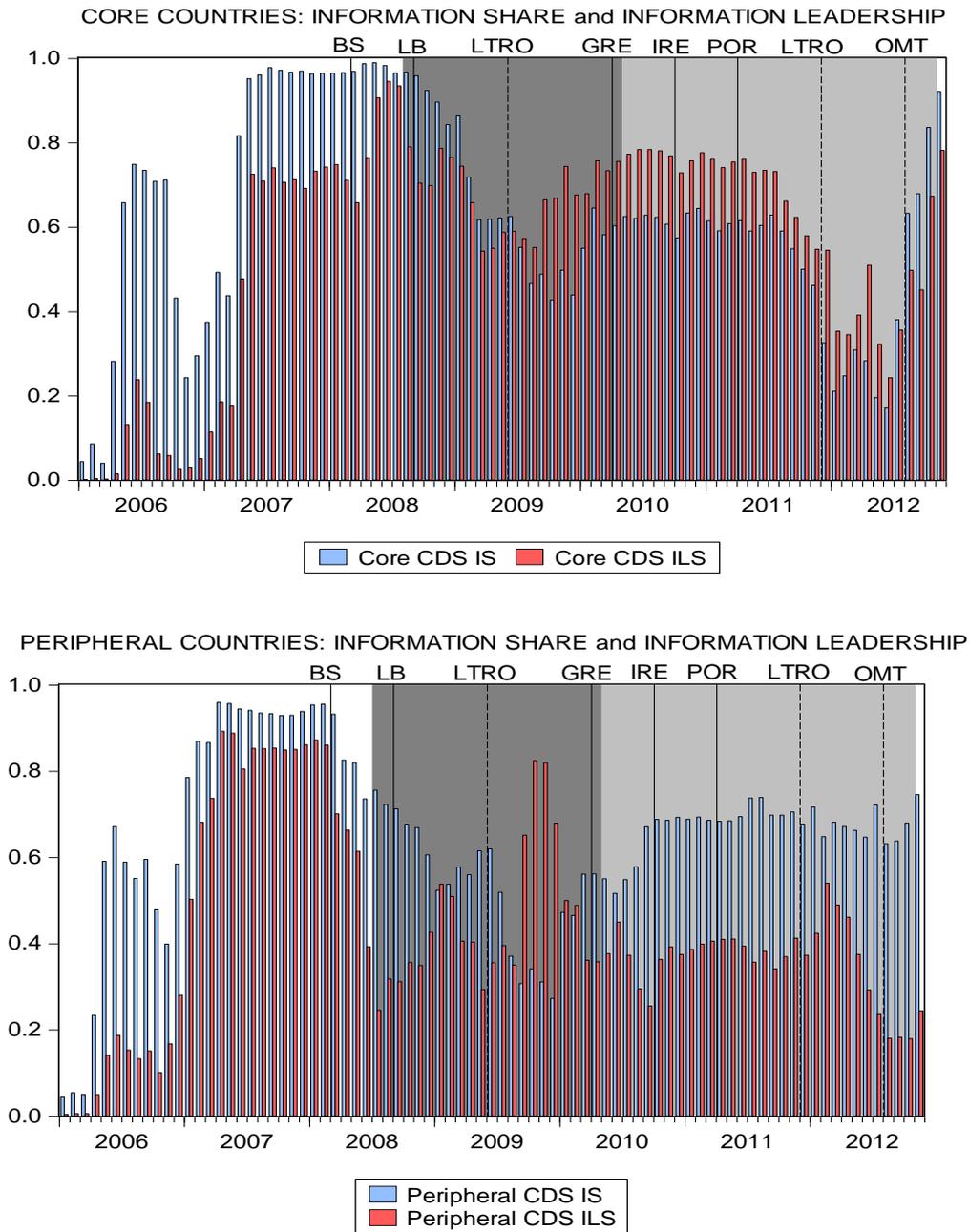


Figure 4: Comparison time-varying IS and ILS of Core and Peripheral CDS market.

This figure shows the time-varying estimates of information share (blue bars) and information leadership share (red bars) as monthly average of the 1787 moving windows each of 500 observations for the aggregate Core group (Belgium, France, Germany) and Peripheral group (Italy, Portugal, Spain) from January 1, 2006 to December 31, 2014. Darker shaded area considers all the moving windows with observations on the dates of the Germany’s Ban and EU proposal. Lighter shaded area considers all the moving windows with observations on the dates of EU’s Ban announcement and went into effect. The latter shaded area covers the final part of the former shaded area. Solid lines represent financial crisis and to sovereign debt crisis events: Bearn Sterns (BS) collapse (March 2008), Lehman Brothers (LB) default (September 2008), Greece (GRE) aid request (April 2010), Ireland (IRE) aid request (October 2010) and Portugal (POR) aid request (April 2011). Dashed lines represent monetary policies announcements: Long Term Refinancing Operations (LTRO) with maturity one year (June, 2009), Long Term Refinancing Operations (LTRO) with maturity three years (December 2011), Outright Monetary Transactions (OMT) announcement (August, 2012).



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Systematic and Idiosyncratic PD of a Loan Portfolio: a Satellite Model for Regulatory Exercises

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ABSTRACT

The paper presents a methodological approach, compliant with European Central Bank stress test requirements, to develop a satellite model and estimate the Probability of Default (PD, i.e. forward-looking default rates) of a loan portfolio according to a macroeconomic scenario.

The approach models the default rates of a loan portfolio as the sum of a systematic risk factor and idiosyncratic risk factor. The systematic risk depends on the economic conditions of a country. The unobservable idiosyncratic risk is assumed to depend on the strategic decisions of bank managers as well as on the specific features of the borrowers.

The paper shows empirical projections of baseline and stressed default rates on corporate, small business and retail loan portfolios of the Italian banking system, through the simulation of European Banking Authority scenarios.

The model has implications for supervisors interested in a common methodology to compare the effects of an economic scenario on different institutions; and for European banks interested in an internal model, compliant with regulatory stress testing exercise, to estimate the PD based on macroeconomic forward-looking information. Moreover, this satellite model can be adopted for other regulatory exercises, such as the ICAAP report and “the expected credit losses impairment” modelling under the new accounting standard IFRS 9.

JEL classification: G17, G21, G28, G32, E44

Keywords: loan portfolio, point-in-time probability of default, systematic and idiosyncratic risk factors, stress test and regulatory exercises

1. INTRODUCTION

Bank for International Settlements (BIS, 2009) defines stress testing as the “evaluation of a bank’s financial position under a severe but plausible scenario to assist in decision making within the bank”.

Since the financial crisis the use of stress tests has considerably increased, both by banks as part of their own risk management and by supervisory authorities as part of their regulatory oversight of the banking sector. In particular, under the Supervisory Review and Evaluation Processes (SREP), the European central Bank (ECB) requests to the majority of European Banks to run stress test exercises, coordinated by the European Banking Authority (EBA). Banks are required to stress test a common set of risks: credit risk, market risk, counterparty credit risk, interest rate risk and operational risk.

In the context of the stress test requirements for credit risk, EBA provides a baseline and an adverse macro-financial scenario and requires banks to estimate their impact on the credit risk parameters in order to translate the impact on the capital available. The projections of capital are required on a future horizon of 3 years. For the estimation of future losses in stress tests, institutions should rely on credit risk parameters different from the ones applied in the calculation of capital requirements, which are usually “through-the-cycle” for PD (probability of default) and under downturn conditions for LGD (loss given default). In particular, institutions should apply estimates based on “point-in-time” parameters in accordance with the severity of the scenario.

This paper, focusing on the stress test of credit risk, attempts to provide a methodological approach, compliant with EBA requirements (“2016 EU-Wide Stress Test”)²¹, to build a satellite model for point-in-time PD (PD-pit) of a loan portfolio.

When it comes to estimate PD-pit, European banks encounter theoretical and practical issues. Among the theoretical ones, first, PD-pit are the forward-looking projections of observed default rates (DRs) and, then, must capture trends in the business cycle²². Second, historical DRs are affected by idiosyncrasies of banks and of borrowers; this makes not trivial to estimate relationship with macro-financial variables consistent from the point of view of economic theory. Among the practical issues, first, banks usually do not have availability of DR time-series over a sufficiently wide time horizon. Second, EBA provides “yearly” forecasts of macroeconomic variables; then a satellite model should allow applying directly EBA “yearly” shocks without transform the unit of measure of variables after the estimates.

²¹ <http://www.eba.europa.eu/documents/10180/1259315/2016+EU-wide+stress+test-Methodological+note.pdf>

²² In contrast to “through-the-cycle” parameters, they should not be business cycle neutral.

The paper presents a theoretical and econometric approach to manage the issues described above. The approach models the default rates on a loan portfolio as the sum of a systematic risk factor and idiosyncratic risk factor. The systematic risk depends on the economic conditions of a country. The idiosyncratic risk depends on the strategic decisions of bank managers as well as on the specific features of the borrowers.

The paper shows empirical projections of baseline and stressed PD-pit on loan portfolios, through the simulation of EBA scenarios on the aggregate default rates of the Italian banking system for corporate, small business and retail sectors.

The approach provided in the paper has implications for supervisors and credit risk managers of European banks. For supervisors interested in a common methodology to compare the effects of an economic scenario on different institutions. For credit risk managers interested in an internal model, compliant with regulatory stress testing framework, to estimate PD-pit, as well as in a tool for credit risk management of loan portfolios. At the same time, this satellite model could be used for other regulatory exercises, such as the ICAAP report and “the expected credit losses impairment” modelling under the new accounting standard IFRS 9. Moreover, the model can be adopted not only for risk management activities, but also for budget and planning operations.

2. STRESS TESTING REGULATORY AND LITERATURE FRAMEWORK

Satellite models can be defined as econometric models used to estimate the relation between the variables quantifying risks bank face (credit risk, market risk, interest rate risk, etc...) and the macroeconomic and financial variables reflecting the state of the economy expected to influence the dynamics of these risks (AIFIRM, 2016).

The main steps involved in the development of a satellite model are:

- i) identification of a target dependent variable (PD, LGD, etc...),
- ii) definition of a long list of macroeconomic independent variables and the expected direction of their impact on the target variable,
- iii) definition of the econometric model and selection, among the estimated models, of the most consistent specification with economic theory,
- iv) forecasts' computation of target variable.

Satellite models are widely used by banks for risk management, budget and planning operations. Within the risk management, the most important applications of satellite models are for the Internal Capital Adequacy and Assessment Process (ICAAP) requirements, for the impairments' calculation under International Financial Reporting Standard (IFRS9) and for the ECB Stress test exercises.

Under ICAAP requirements, a bank needs to have in place internal procedures and processes to ensure that it possesses adequate capital resources in the long term to cover all of its material risks. In particular, the bank will make use of internal models, such as satellite models, to assess, quantify and stress test risk drivers and factors and the amount of capital required to support them (EBA, November 2016: *Guidelines on ICAAP and ILAAP information collected for SREP*).

Under IFRS9 framework, the impairments' calculation of banks must take into account forward-looking information and the effect of macroeconomic scenarios on expected losses of financial instruments (IASB, July 2014: *IFRS9 Financial Instruments*). Then, satellite models are suitable to condition the expected losses to macroeconomic forecasts.

Under ECB stress test of credit risk, banks are required to translate the macroeconomic scenarios, provided by the regulator, into corresponding credit risk impacts on both the capital available (i.e. via impairments and thus the P&L) and the risk exposure amount (REA) for positions exposed to risks stemming from the default of counterparties (EBA, February 2016: *EU Wide Stress Test – Methodological Note*).

For all these applications, banks are requested to make use of their internal satellite models, but are subject to a number of guidelines. Currently, the most important regulatory references on satellite models implementation can be found on the topic of Stress Test (EBA, December 2015: *Guidelines on stress testing and supervisory stress testing*; and EBA, February 2016).

EBA (February 2016) provides a baseline and an adverse macro-financial scenario and requires banks, through satellite models, to estimate their impact on the credit risk parameters in order to translate the impact on the capital available. The projections of capital are required on a future horizon of 3 years. For the estimation of future losses in stress tests, institutions should rely on credit risk parameters different from the ones applied in the calculation of capital requirements, which are usually “through-the-cycle” for PD and under downturn conditions for LGD. In particular, institutions should apply estimates based on “point-in-time” parameters in accordance with the severity of the scenario. Point-in-time are the forward-looking projections of observed DRs and, then, must capture trends in the business cycle.

These are “high-level” guidelines for achieving convergence of practices followed by institutions and competent authorities. However, when it comes to estimate point-in-time risk parameters, banks encounter several issues, due to the absence of “bottom-level” guidelines related to satellite models implementation.

This paper addresses this issue, providing a methodological approach, compliant with EBA requirements (EBA, December 2015 and EBA, February 2016), to build a satellite model for PD-pit of a loan portfolio. And, as satellite models play a fundamental role in many credit risk management activities, the paper aims to provide an approach that banks can adopt consistently and homogeneously across all regulatory exercises (ECB Stress Test, ICAAP, IFRS9). Moreover, this has implications for supervisors interested in a common methodology to compare the effects of an economic scenario on different institutions.

Academic literature on satellite models is currently limited. Relevant papers on this subject are De Bandt et al. (2013) and Assouan (2012). Focusing on the estimation of the PD parameter through its relation with macroeconomic factors, the first work proposes a methodology to implement a satellite model for stress testing “corporate” credit risk, the latter work for the “retail credit risk. Corporate credit risk, also known as “wholesale credit risk” as opposed to “retail credit risk”, is a key component of stress testing for global institutions.

Specifically, both papers are based on a time series approach, in particular, the vector autoregressive (VAR) model in Bandt et al. (2013) and the vector error correction (VEC) model in Assouan (2012)²³. Nonetheless, both approaches are based on Wilson’s (1997a,b) Credit Portfolio View. It develops a credit risk model linking macroeconomic factors (GDP, unemployment, inflation, house prices etc...) and corporate sector default rates. The idea was to model the relationship between default rates²⁴ and macroeconomic factors and, when a model is fitted, to simulate the evolution of default rates over time by applying a stress scenario to the model. The simulated default rates in turn make it possible to obtain estimates of default rates for a defined credit portfolio under a given stress scenario.

In applying such a model, the main issue that practitioners encounter is that default rates are affected also by idiosyncrasies of banks and of borrowers; this makes difficult to estimate relationship with macro-financial variables consistent with economic theory. This difficulty is even greater when it comes to estimate credit risk parameters related to retail portfolios. Indeed, in retail segment the loan

²³ Basically, a VEC model is a VAR model plus a cointegration term. In other words, the VEC model considers not only the short term relations (as in VAR models) but also the long term relations between variables.

²⁴ An important features of Wilson (1997a,b) is to use as dependent variable, not the default rates, but the logistic functional form of default rates.

portfolios present smaller exposure amounts and higher heterogeneity of borrowers respect to corporate portfolios. This should makes PD of retail portfolios more influenced by idiosyncratic factors than the PD of corporate portfolios.

Currently, the satellite models presented in literature and, then, used by practitioners, are biased because do not control for the idiosyncratic risk. Indeed, as pointed out by Souza and Feijò (2011) the causes of the risk of defaulting on bank loans can be divided into two groups: macroeconomic (or structural) factors and microeconomic (or idiosyncratic) factors. While the first group is linked to the general state of the economy, the second group is related to the individual behaviour of each bank and its borrowers.

This paper presents an approach for satellite models' implementation that can separate the effect of macroeconomic and microeconomic risk factors on risk parameters. The approach models the default rates on a loan portfolio as the sum of a systematic risk factor and idiosyncratic risk factor. The systematic risk depends on the economic conditions of a country. The idiosyncratic risk depends on the strategic decisions of bank managers as well as on the specific features of the borrowers. The model can be extended in future research by considering more complex specifications of the idiosyncratic process over time.

Apart from theoretical issues, the paper provides also hints to solve practical issues that practitioners face in building satellite models. In particular, a statistical solution is provided in case banks do not have availability of default rates time-series over a sufficiently wide time horizon. And it is explained in detail how to express the variables in the model in order to apply consistently the economic forecasts to derive forward-looking parameters.

3. A SATELLITE MODEL FOR PROBABILITY OF DEFAULT

We model the $PD_{b,j,c,t}$ of a loan portfolio:

- on segment “j”²⁵
- in country “c”
- owned by bank “b” (with $b \in B$ = banking system in country “c”)
- at time “t”,

²⁵ Main segments on loan portfolios are Corporate, Small Business and Retail.

as a function of a bank-segment-country specific index (Y):

$$PD_{b,j,c,t} = \text{probit}^{-1}(Y_{b,j,c,t}) = \Phi(Y_{b,j,c,t}) \quad (1)$$

Thus, through the probit²⁶ transformation of the PD, a bank-segment-country specific index (Y) is derived, which depends upon a systematic risk factor (S) and an idiosyncratic risk factor (I).

We assume that $Y_{b,j,c,t}$ follows a non-stationary process:

$$Y_{b,j,c,t} = Y_{b,j,c,t-1} + S_{b,j,c,t} + I_{b,j,c,t} \quad (2)$$

where:

- $Y_{b,j,c,t-1}$: last period's value;
- $S_{b,j,c,t}$: stationary process with mean μ_s and variance σ_s^2 .
- $I_{b,j,c,t}$: stationary process with mean 0 and variance σ_i^2 ;

Taking first differences, we get:

$$Y_{b,j,c,t} - Y_{b,j,c,t-1} = \Delta Y_{b,j,c,t} = S_{b,j,c,t} + I_{b,j,c,t} \quad (3)$$

Eq.3 shows that changes in bank-segment-country specific index (Y) are equal to the sum of systematic risk factor (S) shocks and idiosyncratic risk factor (I) shocks.

$S_{b,j,c,t}$ is specified through the following form:

$$S_{b,j,c,t} = \beta_{b,j,c,0} + \sum_{n=1}^{n^{\circ} \text{ of macro-financial variables}} \beta_{b,j,c,n} * X_{B,j,c,n,T} \quad (4)$$

where:

- $S_{b,j,c,t}$ is the “systematic” shock of the index value Y for bank “b” on segment “j” in country “c” at time “t”;
- $(\beta_{b,j,c,0}, \beta_{b,j,c,1}, \beta_{b,j,c,2}, \dots)$ is a set of regression coefficients to be estimated specifically for bank “b” on segment “j” in country “c”;

²⁶ The Probit functional form is chosen over linear and exponential representations because, for any value of the index Y, it yields a probability between [0,1]. This is important, because it is a desirable property for a probability to have. Moreover, the Probit functional form is chosen over Logit functional form used by Wilson (1997 a, b) in his credit portfolio view approach, because it offers on average a better fit.

- $(X_{B,j,c,1,T}, X_{B,j,c,2,T}, \dots)$ is a set of macroeconomic and financial factors for the banking system “B” on segment “j” in country “c” at time “T” (where T can be t, t-1, ..., t-n) ;

The economic interpretation of $S_{b,j,c,t}$ is the following. $S_{b,j,c,t}$ is driven by the state of the economy in which bank “b” and all banking system “B” operate. For instance, higher defaults are associated with or even caused by weaker economies or recessions: as GDP growth declines, defaults and insolvencies tend to increase. Then, “b” and all other banks belonging to “B” share this macroeconomic component in their PD on loan portfolios. However, the sensitivity (β 's) of loan PD to macroeconomic variables will vary among the banks in the country.

$I_{b,j,c,t}$ is specified through the following form:

$$I_{b,j,c,t} = (\varphi_{b,j,c} * I_{b,j,c,t-k}) + \varepsilon_{b,j,c,t} = \text{management risk} + \text{residual risk} \quad (5)$$

where:

- $I_{b,j,c,t}$ is the “idiosyncratic” shock of the index value Y for bank “b” on segment “j” in country “c” at time “t”;
- $(\varphi_{b,j,c})$ can be one or more autoregressive coefficients to be estimated specifically for bank “b” on segment “j” in country “c”;
- $\varepsilon_{b,j,c,t}$ is a random variable assumed to be independent and identically normally distributed, i.e.: $\varepsilon_{b,j,c,t} \sim iid N(0, \sigma_\varepsilon^2)$.

The economic interpretation of $I_{b,j,c,t}$ is the following. $I_{b,j,c,t}$ represents the unobservable idiosyncratic risk of the loan portfolio, that is the risk related to the individual behaviour of each bank and its borrowers. $I_{b,j,c,t}$ cannot be directly observed by “known” variables, but we assume can be explained by two elements in eq.5. $\varepsilon_{b,j,c,t}$ represents that element of idiosyncratic risk that is unique of borrowers on segment “j”, in country “c” and is not correlated with the systematic risk factor. We call it “Residual” risk. There is also another element of idiosyncratic risk, $(\varphi_{b,j,c} * I_{b,j,c,t-k})$ that we call “Management” risk. Because idiosyncratic risk is determined by the intrinsic characteristics of each borrower and lending institution, this type of portfolio risk can be partially adjusted by banks for various purposes. This means to say that although the macroeconomic environment affects the portfolios of all banks, they react differently to obtain the best opportunities or to protect themselves. Business decisions of bank managers should partially reflect this factor: a partial control over

idiosyncratic risk from managers can be used to offset changes in the the level of PD²⁷. For assumption manager cannot interfere in the systematic risk, but can interfere only in the idiosyncratic risk. At time t-k, increases in $I_{b,j,c,t-k}$ and, then, in $PD_{b,j,c,t-k}$ should induce actions by banks to lower their idiosyncratic risk from $I_{b,j,c,t-k}$ to $I_{b,j,c,t}$ and thus to maintain, at time t, their overall risk $PD_{b,j,c,t}$ at the desired levels established by management. It means that, in eq.5, only if the coefficient $\varphi_{b,j,c}$ has negative sign, the managerial decisions are correctly offsetting changes in the level of PD. Overall, the entity and the sign of $\varphi_{b,j,c}$ will vary among banks due to managerial abilities and strategies.

Recalling eq.3, the relative importance of $S_{b,j,c,t}$ and $I_{b,j,c,t}$ to explain $\Delta Y_{b,j,c,t}$ varies among sectors and banks. For instance, it is expected that the corporate sector will be more influenced by macroeconomic factors than the retail sector. On the contrary, the idiosyncratic factor should weigh more in the retail sector than in the corporate sector. Moreover, as described above, regarding the idiosyncratic factor is important to distinguish between the component related to business actions undertaken by each bank and the residual component related to intrinsic features of the borrowers.

Then, recalling eq. 3, 4 and 5 we have:

$$\Delta Y_{b,j,c,t} = S_{b,j,c,t} + I_{b,j,c,t} = \beta_{b,j,c,0} + \sum_{n=1}^{n^{\text{of macro-financial variables}}} \beta_{b,j,c,n} * X_{B,j,c,n,T} + (\varphi_{b,j,c} * I_{b,j,c,t-k}) + \varepsilon_{b,j,c,t} \quad (6)$$

In summary, this theoretical approach models default rates of a loan portfolio by estimating their systematic and idiosyncratic risk components to allow to simulate the evolution over time of the default rates by applying a macroeconomic scenario to the model. The simulation of default rates in turn make it possible to obtain estimates of PD-pit for a loan portfolio under a given scenario.

4. ESTIMATION PROCEDURE

This section presents the estimation procedure of the satellite model described in previous section.

First of all, it is worth nothing the case of Italy where not all banks have available DR time-series sufficiently long to estimate robust econometric regressions. Usually, DR time-series cover only recent years. However, Italian banks have available time-series of “tassi di decadimento” (DECR) from 1996 to the present. The difference between DR and DECR is that, the first consider the entire

²⁷ Although it is impossible for banks to alter the idiosyncratic risk of each borrower or loan, they can modify the idiosyncratic risk at the portfolio level.

flow of loans from bonis to default, while the latter consider only the flow of loans from bonis to the worst bucket of default classification, called “sofferenze”. In Italy, there are 3 buckets of default called “scaduti”, “incagli” and “sofferenze”. In such a case, it is possible to extend the DR series using DECR series with different techniques²⁸. We introduce a new statistical technique, called “target standardized approach” (TSA), which aims to reduce the degree of discretion of the credit analyst respect to the other approaches in note 27. The TSA consists of 4 steps:

- 1) definition of target (benchmark) values of mean and standard deviation of DR series for the period with missing data;
- 2) standardization of DECR time-series:

$$Z(DECR)_t = \frac{DECR_t - \text{Mean}[DECR_{\text{extension period}}]}{\text{St.dev.}[DECR_{\text{extension period}}]} \quad (7)$$

where t includes all the observations in the chosen extension period for DR;

- 3) substitution in Z(DECR) of step 2, the target values of mean and standard deviation of DR series of step 1, to find DR values on the extension period:

$$DR_t = Z(DECR)_t * \text{Target St. dev. [DR]} + \text{Target Mean[DR]} \quad (8)$$

Due to the aim to estimate PD-pit consistent with the scope of ECB stress testing, we apply some constraints:

- not allowing for lagged dependent variables among regressors;
- allowing at most of one lag on independent variables;
- among the macro-financial variables, we use only those included in EBA macroeconomic baseline and adverse scenarios: GDP, inflation, unemployment, long-term interest rates and house prices;
- to apply directly EBA “yearly” shocks, without ex-post transformation of the frequency of variables, the following techniques are applied. Variables, expressed in levels, are computed by annualizing the observations and maintaining the quarterly frequency. Variable, expressed in first differences, are computed as the difference between the observation on a quarter of year x and the observation on the same quarter on year x-1.

²⁸ For instance, conversion factors based on regression approach or DR/TDEC ratio approach.

In order to be included in the estimation of our satellite model, each relevant macro-financial variable has to fulfil two conditions:

1. its coefficient must be statistically significant and improve the overall model's performance;
2. its coefficient must be consistent with economic theory.

As it concerns condition 1, in the empirical analysis we impose a threshold of significance at 15% on the p-value of each coefficient. As it concerns condition 2, we report, in Table 1.1, the expected sign of the relation between the macro-financial variables and the default rates on loans. The economic interpretation on the expected sign of the coefficient for each macro-financial variable is the following:

- GDP growth: a reduction of GDP indicates a contraction phase of economic cycle and, then, a reduction of aggregate demand, resulting in an increase of default rates on loans;
- Inflation rate: a reduction of inflation explained by the impact on prices of a weaker aggregate demand is associated to an increase of default rates on loans;
- Unemployment rate: an increase of unemployment rate denotes a contraction phase of economic cycle and, then, a reduction of disposable income, resulting in an increase of default rates on loans;
- Property price index: a reduction of property prices caused by a reduction of consumer confidence and, then, as a reaction to a general deterioration of economic outlook is associated to an increase of default rates on loans;
- Long term Italian government interest rate: an increase of interest rates, due to an increase of risk aversion and reduction of liquidity, indicates the uncertainty of the markets in economic and political outlook of a country and, then, it is associated to an increase of default rates on loans;
- Sovereign spread: an increase of sovereign spread denotes and increase of credit risk of a country due to negative expectations on the economic outlook and, then, it is associated to an increase of default rates on loans.

Due to the presence of $I_{b,j,c,t-k}$ among regressors in eq.6, the estimation procedure of our satellite model consists of two steps:

1. STEP: we regress the bank-segment-country specific index ($Y_{bjc,t}$) in first differences on stationary macroeconomic variables ($X_{B,j,c,n,T}$) to find the estimates of the idiosyncratic component ($I_{b,j,c,t}$) over time:

$$\Delta Y_{bjc,t} = \gamma_{b,j,c,0} + \sum_{n=0}^{n^{\circ}of\ macro-financial\ variables} \gamma_{b,j,c,n} * X_{B,j,c,n,T} + I_{b,j,c,t} \quad (9)$$

2. STEP: we model the idiosyncratic component as an autoregressive process, that means we estimate eq.9 under the following form:

$$\Delta Y_{bjc,t} = \beta_{b,j,c,0} + \sum_{n=0}^{n^{\circ}of\ macro-financial\ variables} \beta_{b,j,c,n} * X_{B,j,c,n,T} + (\varphi_{b,j,c} * I_{b,j,c,t-1}) + \varepsilon_{b,j,c,t} \quad (10)$$

Once the parameters $\beta_{b,j,c,n}$ and $\varphi_{b,j,c}$ have been estimated, we apply EBA macro-financial scenarios to obtain PD-pit estimates.

5. DATASET

The dataset covers the period 1Q1996-4Q2015²⁹. Variables are retrieved from the databases of Bank of Italy, Eurostat, Bank of International Settlements.

The macro-financial variables considered are those included in the macro-financial adverse scenario of EBA (February 2016): GDP, inflation, unemployment, long-term interest rates and house prices. It is fundamental to highlight that we express all variables on yearly basis. The reason is that EBA forecasts are provided on yearly variables (for 2016, 2017 and 2018), then, when we use these forecasts to estimate PD-pit, we need a model with all variables expressed in years. Simply switching time-series from quarterly to yearly frequency is not recommendable due to the high reduction of observations. Then, we express variables on yearly basis but maintaining quarterly frequency, through transformation techniques explained below.

We consider the following macro-financial variables:

- Yearly Italian and EU³⁰ real GDP growth rate on quarterly frequency:

$$GDP_q = \ln \left[\frac{(GDPV_q + GDPV_{q-1} + GDPV_{q-2} + GDPV_{q-3})}{(GDPV_{q-4} + GDPV_{q-5} + GDPV_{q-6} + GDPV_{q-7})} \right]$$

where:

- $q = \text{quarters}$

²⁹ Due to the data availability of the dependent variable, i.e. default rates on Italian corporate, small business and retail loans.

³⁰ European Union 28 countries.

- $GDPV_q = \text{quarterly total value of real GDP}$

- Yearly Italian and EU inflation rate on quarterly frequency:

$$HICP_q = \ln \left[\frac{(HICPV_q)}{(HICPV_{q-4})} \right]$$

where:

- $HICPV_q = \text{quarterly value of Harmonised Index of Consumer Prices (HICP)}$

- Yearly average of Italian and EU unemployment rate on quarterly frequency:

$$UR_q = \frac{UNM_q + UNM_{q-1} + UNM_{q-2} + UNM_{q-3}}{4}$$

where:

- $UNM_q = \text{quarterly average of unemployment rate}$

- Yearly growth rate of Italian residential house prices Index on quarterly frequency:

$$HPI RE_q = \ln \left[\frac{(HIRE_q)}{(HIRE_{q-4})} \right]$$

where:

- $HIRE_q: \text{quarterly average of Italian residential house prices index}$

- Yearly growth rate of Italian Commercial residential house prices Index on quarterly frequency:

$$HPI CO_q = \ln \left[\frac{(HICO_q)}{(HICO_{q-4})} \right]$$

where:

- $HICO_q: \text{quarterly average of Italian Commercial house prices index}$

- Yearly average of Italian long term interest rate on quarterly frequency:

$$LR_q = \frac{LTIR_q + LTIR_{q-1} + LTIR_{q-2} + LTIR_{q-3}}{4}$$

where:

- $LTIR_q = \text{quarterly average of long term interest rate 10Y}$

- Yearly average of Sovereign Spread Italy vs Germany on quarterly frequency:

$$SOV_q = \frac{SOVS_q + SOVS_{q-1} + SOVS_{q-2} + SOVS_{q-3}}{4}$$

where:

- $SOVS_q =$ quarterly average of sovereign spread Italy vs Germany 10Y

Available default measures on loans, precisely “tassi di decadimento dei finanziamenti per cassa” (DECR), of Italian banking system are retrieved from the statistical database of Bank of Italy. DECR on single banks are not public available. Thus, aggregate DECR of Italian banking system are collected for corporate segment (excluding financial companies), small business segment and retail segment. Bank of Italy computes quarterly DECR (in %) as the ratio:

$$DR_q = \frac{n^\circ \text{ of counterparties in bonis at the beginning of the quarter } (q) \text{ and in default at the end of the quarter } (q)}{n^\circ \text{ of counterparties in bonis at the beginning of the quarter } (q)}$$

Then, we apply the following formula to have yearly DECR on quarterly frequency for each loan segment:

$$DECR_q = 1 - [(1 - DECR_q) * (1 - DECR_{q-1}) * (1 - DECR_{q-2}) * (1 - DECR_{q-3})]$$

Where:

$DECR_q =$ quarterly rates of “decadimento” from Bank of Italy

As explained in the estimation procedure section, we need to transform DECR in DR, because only the second consider the entire flow of loans from bonis to default, while the first consider only the flow of loans from bonis to the worst bucket of default classification, called “sofferenze”. We compute this transformation using the technique “TSA” introduced before. TSA requires selecting DR target values of mean and standard deviation (shown in table 1 for each of three segments³¹) over the period 1996-2015.

Table 1 reports the descriptive statistics of the dataset.

It is worth nothing that, when necessary, further transformations are applied to ensure that all variables are stationary and expressed in the same unit of measure, that is as percentage variation respect to the year before. Following these requirements, we transform in first differences the Italian

³¹ The chosen target averages of DR for Corporate, Small Business and Retail segments are respectively 5%, 4.5% and 4%, in accordance with market benchmarks. For the target standard deviations, we used the values of standard deviations of the DECR.

and European unemployment rates, the Italian long-term interest rate and the Italian Sovereign Spread as well as the dependent variables, that is, the probit of default rates for each loan segment.

6. EMPIRICAL ANALYSIS

In this section, we present the empirical findings of our satellite model on the Corporate, Small Business and Retail loan segments for the aggregate Italian banking system. We will present 4 panels, each one with a focus on a different metric to compare the results on the three loan segments.

In panel 1, we report the parameters' estimates for each segment.

The performance of the model is satisfactory, with an Adj. R-Square of 92% for Corporate segment, 69% for Small Business segment and 68% for Retail segment. This evidence is also supported in Panel 2, where we report for each loan segment, a graph with the evolution of observed vs predicted values of Italian banking system index (probit of default rates) in first differences.

The models selected are those satisfying, for each segment, the conditions 1 and 2 described in the section of estimation procedure³². Moreover, we have preferred models influenced more by macroeconomic than financial variables, in accordance with the interpretation that the real economy has stronger impact than financial economy on loan defaults. Corporate and Small Business defaults rates are influenced by the same variables: Italian GDP growth rate, Italian Harmonized Index of Consumer Price, European unemployment rate and Italian long-term interest rate. Retail segment is influenced by the same variables, except for the price level indicator that, not surprisingly for this segment, is represented by the Italian residential House Price Index. For all segments, long-term interest rate has 1 year lag, since financial variables usually anticipate default rates. Unemployment rate is lagged one year too, in accordance with the expectation that the effect of job lost on default rates spreads out not immediately, due to the existence of welfare mechanisms.

For Corporate and Small Business sectors, the Idiosyncratic factor related to management actions is highly significant (in both case under 1%) and with negative sign, showing a partial control of aggregate Italian banks to offset changes in the level of PD to maintain portfolio risk at desiderated levels. While for retail loans, it is not significant and also with positive sign, showing inability of aggregate Italian banks to adjust the portfolio risk on this segment. These evidences are supported in table 2, where we report, for each segment, the correlation between Systematic factor

³² It is worth nothing that other model specifications, in terms of variables selected, fulfil conditions 1 and 2 described in the section of estimation procedure. Then, the model specifications presented are illustrative. In a real application, other model specifications can be chosen according to the risk-profile of the portfolio and the credit risk manager experience.

$(\beta_{b,j,c,0} + \sum_{n=1}^{n^{\circ} \text{ of macro. var.}} \beta_{b,j,c,n} * X_{B,j,c,n,T})$ and Idiosyncratic factor, distinguishing between management related $(\varphi_{b,j,c} * I_{b,j,c,t-1})$ and residual $(\varepsilon_{b,j,c,t})$. For Corporate and Small Business segments we find a negative correlation between the idiosyncratic factor related to management and the systematic factor, suggesting some ability of Italian banking system to offset the macroeconomic impact on default rates. This ability is higher for corporate sector (-11%) than small business sector (-4). While for retail segment the correlation is positive, suggesting no ability of Italian banking system to reduce the macroeconomic impact on default rates. It is also worth nothing that, in accordance with the theoretical assumptions of our model, there is absence of correlation between the residual idiosyncratic factor and the systematic factor, for all segments.

In table 3, we report the relevance of all variables to explain the volatility of default rates. This indicator assign a weight (%) to each variable³³ and is computed as:

$$\mathbf{Weight\ of\ variables}_i = \frac{|\beta_i| * sdX_i / sdY}{|\beta_1| * \frac{sdX_1}{sdY} + \dots + |\beta_i| * \frac{sdX_i}{sdY} + \dots + |\beta_n| * \frac{sdX_n}{sdY}} \quad (11)$$

Where:

- sd is the standard deviation;
- β is the estimated coefficient of each variable;

The weight of the systematic factor is equal to the sum of the weights of all macroeconomic and financial variables. The weight of the idiosyncratic factor is equal to the sum of the weights of management related and residual idiosyncratic factors.

Among the main evidences on the results of weights, first, the systematic factor has higher impact on corporate defaults rates (77%) rather than on Small Business and retail defaults (64%). Second, a relevance of management actions on default behaviour is present only for Corporate (10%) and Small Business (11%) loans, while for retail segment is negligible (3%). This is line with the previous results where we found no statistically significance of management actions on retail loans. Third, the Small Business and, in particular, the Retail loans are the segments where we find the higher influence (25% and 33% respectively) of intrinsic values of borrowers to explain the dynamics of default rates.

These evidences on weights are supported in panel 3, where we report, for each segment, the evolution of the estimated Systematic and Idiosyncratic factors. The three graphs show that as we move from segments with higher loans' exposure amount, Corporate and Small Business, to a segment with lower exposures, Retail, the values of the systematic component of default rates are smaller and less

³³ For the "residual" idiosyncratic factor the " β " is put equal to one.

volatile, as well as for the idiosyncratic factor related to management. On the contrary, the values related to the intrinsic features of the borrowers (residual idiosyncratic factor) are bigger and more volatile.

7. STRESS TEST RESULTS

In this section, we show the PD-pit estimates of our satellite model, after the application of a macroeconomic scenario, in order to fulfil BCE requirements for stress testing exercises on impairments for credit losses.

The scenarios (both baseline and adverse) used in this paper have been defined by the EBA for the 2016 stress-test exercise. The objective is to assess the resilience of a large sample of banks in the EU against an adverse but plausible scenario. The scenario assesses banks against a deterioration from the baseline forecast in the main macroeconomic variables such as GDP, inflation and unemployment. Moreover, changes in interest rates also affect the cost of funding for banks under the stress. Table 4 presents EBA baseline and adverse scenarios for the macro-financial variables selected in this paper for stress testing PD on corporate, small business and retail loans owned by the Italian banking system.

In panel 4, we present the estimated evolution in 2016, 2017 and 2018, of the one year default rates of corporate, small business and retail loans under baseline and adverse scenarios for Italian banking system.

In table 5, we report PD-pit estimates under baseline and adverse scenarios for Corporate, Small Business and Retail segments of Italian Banking System. These results, for each segment, show the sensitivity of the PD to the macroeconomic variables in the scenario.

For the aggregate corporate loan portfolio of Italian banking system under the baseline scenario, the default rate has decreased to 6% in 2018; by contrast, the PD reaches 7.4% under the adverse scenario. In terms of impact respect to 2015, under the adverse scenario the PD increases of 15% in 2018, while in the baseline decrease of 6% in 2018.

For aggregate small business loan portfolio, under the baseline scenario, the default rate has decreased to 4.4% in 2018, which means an impact on 2015 of minus 10%. While under the adverse scenario, the impact is plus 6% with a PD in 2018 at 5.3%.

For aggregate retail loan portfolio, the impact respect to 2015 is minus 10% under the baseline scenario and plus 4% under the adverse scenario.

Overall, the model forecasts, for all segments, a monotonic trend of PD, increasing in the adverse scenario and decreasing in the baseline scenario, coherently with regulator's expectations. In case of deterioration of the economic environment, the corporate default rates present higher sensitivity than small business and retail default rates.

8. CONCLUSIONS

In this paper, we present a satellite model for stress testing the PD of a loan portfolio, in compliance with current regulatory guidelines. Our model is based on the assumption that the PD of a loan portfolio is influenced by systematic and idiosyncratic risk factors. A distinguish feature of our model is to extract these two factors and, then, disentangle the idiosyncratic factor in the Management component and in the residual component. The first component models the control over idiosyncratic risk from managers used to offset changes in the level of PD. For instance, an increase of macroeconomic risk generates an increase of default correlations across borrowers, but the entity of this increase will vary among banks according to their strategic decisions. The second component is related to the intrinsic features of the borrowers and is not correlated with the systematic factor.

First, we find that the economic conditions of the country are more relevant on the level of default rates in the corporate segment than in the Small Business and retail segments. Second, only in the corporate and small business segments we find an ability of manager to control, at least to some extent, the idiosyncratic risk to offset changes in the portfolio PD. Third, the default rates in retail segment are the most influenced by the idiosyncratic features of the borrowers. Indeed, historically, retail default rates are the most difficult to model and to link to macroeconomic variables.

Nonetheless, for all segments, our specification for the satellite model show a significant and robust relationship between key macroeconomic variables and loan default rates. Overall, the results on the stress test exercise suggest a coherent monotonic evolution of PD-pit under EBA scenarios, decreasing at most of 10% under the baseline scenario, and increasing at most of 15% under the adverse scenario.

For the empirical analysis, the model has been simulated in the contest of ECB stress test, but the theoretical and econometric approach to build this satellite model are particularly well suited to other regulatory exercises such as ICAAP report and the expected credit losses impairment" modelling under the new accounting standard IFRS 9. Moreover, the model can be adopted not only for risk management activities, but also for budget and planning operations.

9. TABLES

Table 1: Descriptive Statistics of Dataset (yearly variables on quarterly frequency) 1Q1997-4Q2015

Variables	Mean	Median	Std Dev	Minimum	Maximum
Default Rates Corporate Loans of aggregate Italian Banking System	0.050	0.048	0.007	0.043	0.066
Default Rates Small Business Loans of aggregate Italian Banking System	0.045	0.044	0.005	0.039	0.058
Default Rates Retail Loans of aggregate Italian Banking System	0.040	0.038	0.007	0.033	0.062
European GDP Growth Rate	0.016	0.020	0.017	-0.045	0.041
Italian GDP Growth Rate	0.004	0.007	0.020	-0.059	0.039
European HICP Rate	0.026	0.024	0.015	-0.003	0.080
Italian HICP Rate	0.020	0.021	0.009	-0.002	0.039
European Unemployment Rate	0.097	0.095	0.013	0.074	0.120
Italian Unemployment Rate	0.093	0.088	0.019	0.061	0.127
Italian House Price Index Rate Residential	0.056	0.023	0.082	-0.057	0.204
Italian House Price Index Rate Commercial	0.029	0.034	0.040	-0.033	0.133
Italian Government Long term Interest rates 10Y	0.046	0.046	0.011	0.017	0.084
Sovereign Spread Italy vs Germany 10Y	0.010	0.004	0.011	0.001	0.042

Table 2.1: Expected sign between EBA macro-financial variables and default rates on loans. All series are expressed as variation respect to the previous year.

Regressors	Expected sign on the regressor coefficient
GDP EU and IT	-
HICP EU and IT	-
UR EU and IT	+
HPI RE and CO	-
LTIR IT	+
SOV SPREAD	+

Table 2: Correlations between Systematic and Idiosyncratic factors of Corporate, Small Business and Retail segments

CORRELATIONS		
CORPORATE	Idiosyncratic	
	Management	Residual
Systematic	-11%	0%
SMALL BUSINESS	Idiosyncratic	
	Management	Residual
Systematic	-4%	0%
RETAIL	Idiosyncratic	
	Management	Residual
Systematic	15%	0%

Table 3: Weight of variables to explain the volatility of default rates for Corporate, Small Business and Retail segment

Segment	Factor	Weight	Variable	Weight
CORPORATE	SYSTEMATIC	77%	GDP IT	35%
			HICP IT	11%
			UR EU	18%
			LR IT	13%
	IDIOSYNCRATIC	23%	Management	10%
			Residual	14%
SMALL BUSINESS	SYSTEMATIC	64%	GDP IT	25%
			HICP IT	9%
			UR EU	17%
			LR IT	13%
	IDIOSYNCRATIC	36%	Management	11%
			Residual	25%
RETAIL	SYSTEMATIC	64%	GDP IT	9%
			HPI RE IT	40%
			UR EU	8%
			LR IT	7%
	IDIOSYNCRATIC	36%	Management	3%
			Residual	33%

Table 4: EBA Baseline and Adverse scenario on selected macroeconomic and financial variables

EBA forecasts	BASELINE			ADVERSE		
	2016	2017	2018	2016	2017	2018
Variables (%)						
IT GDP growth rate	1.5	1.4	1.7	-0.4	-1.1	0.0
Italian HICP rate	1.0	1.9	2.8	-0.1	0.3	0.3
Italian HPI Residential	2.0	4.1	5.9	-9.6	-2.1	1.5
European Unemployment rate	9.20	8.90	8.90	9.90	10.80	11.60
Italian Long-term interest rate	1.8	2.0	2.1	2.9	3.0	3.0

Table 5: PD-pit estimates under baseline and adverse scenarios for Corporate, Small Business and Retail segments of Italian Banking System

Segment	Scenario	PD-pit			Impact		
		2016	2017	2018	2016 / 2015	2017 / 2015	2018 / 2015
Corporate	Baseline	6.4%	6.2%	6.0%	0.99	0.95	0.94
	Adverse	6.7%	7.0%	7.4%	1.03	1.08	1.15
Small Business	Baseline	4.8%	4.6%	4.4%	0.97	0.92	0.90
	Adverse	4.9%	5.1%	5.3%	1.00	1.03	1.06
Retail	Baseline	3.7%	3.6%	3.5%	0.97	0.93	0.90
	Adverse	3.9%	4.0%	4.0%	1.02	1.03	1.04

10. PANELS

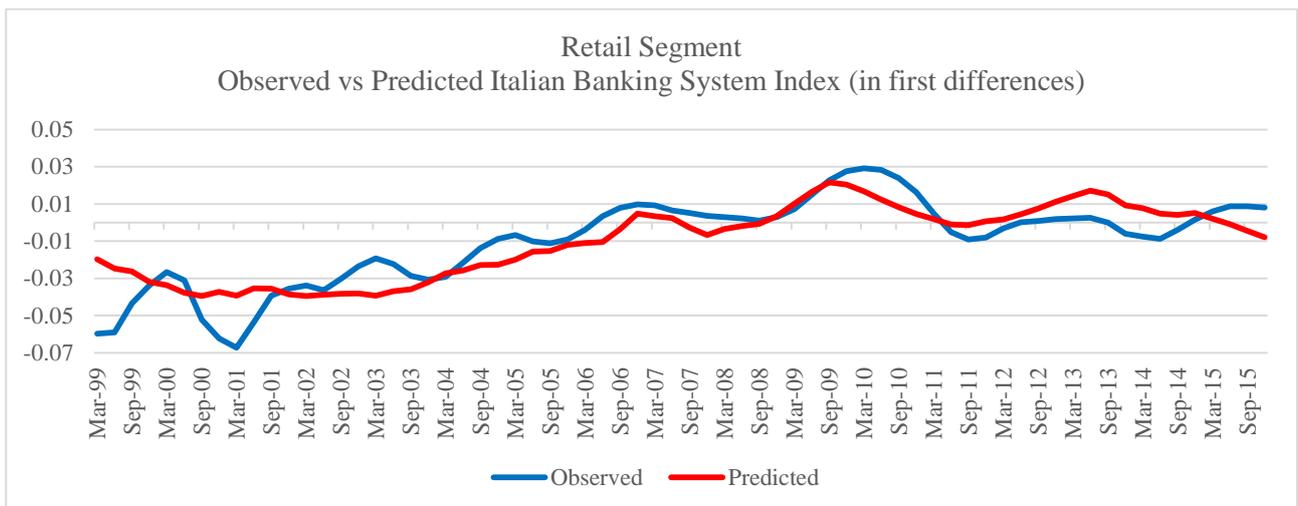
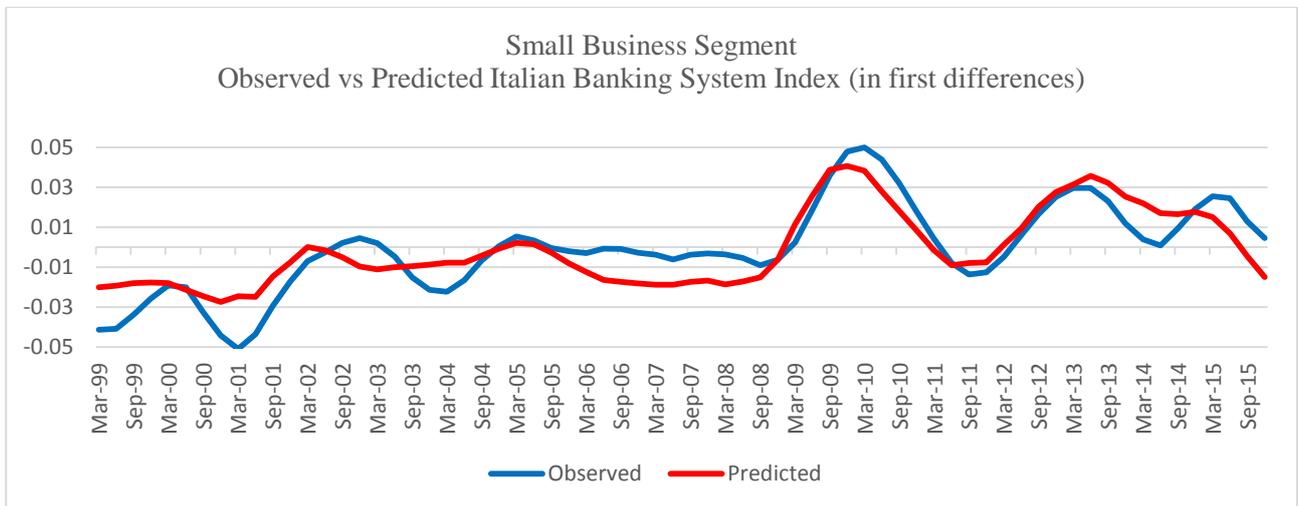
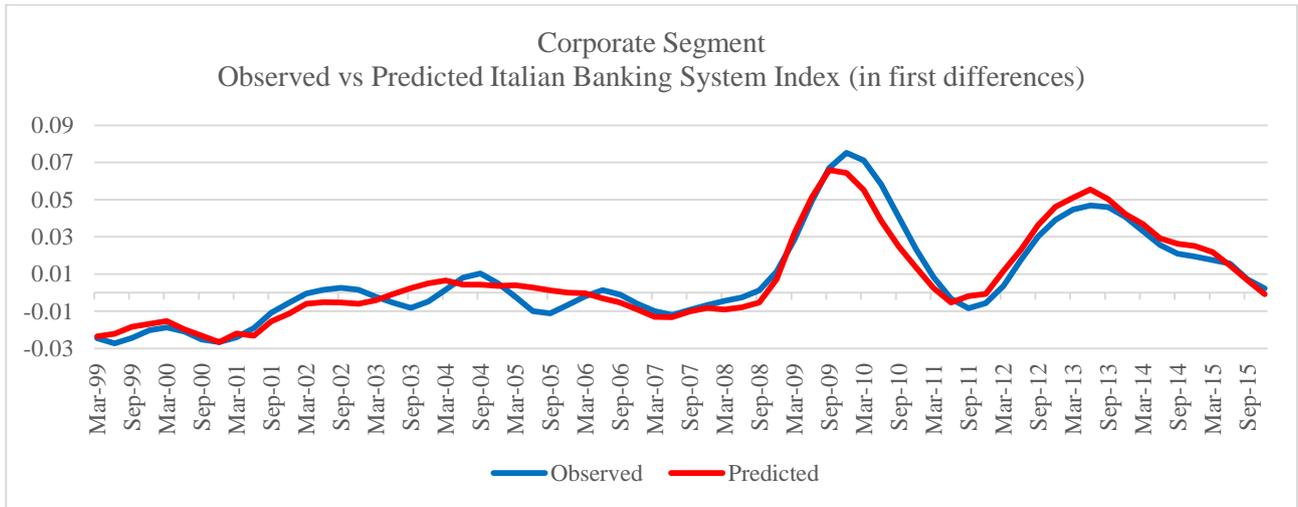
Panel 1: Parameters estimates for Corporate, Small Business and Retail segments

Dependent Variable: Probit of Corporate Default Rates in first difference			
Regressors	Lag	Coef. Estim.	p-Value
Intercept		0.0213	<.0001
GDP IT	0	-0.8182	<.0001
HICP IT	0	-0.5263	<.0001
UR EU	1 Year	1.1851	<.0001
LR IT	1 Year	0.7351	<.0001
Idiosyncratic (Management)	1 Year	-0.4593	<.0001
Adj. R-Square	92%		

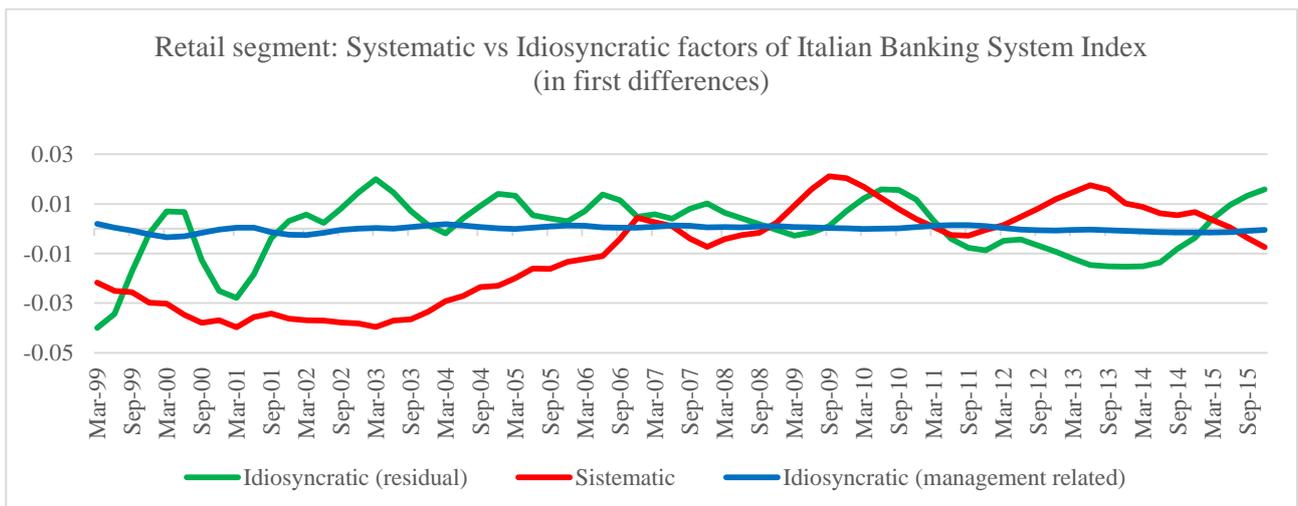
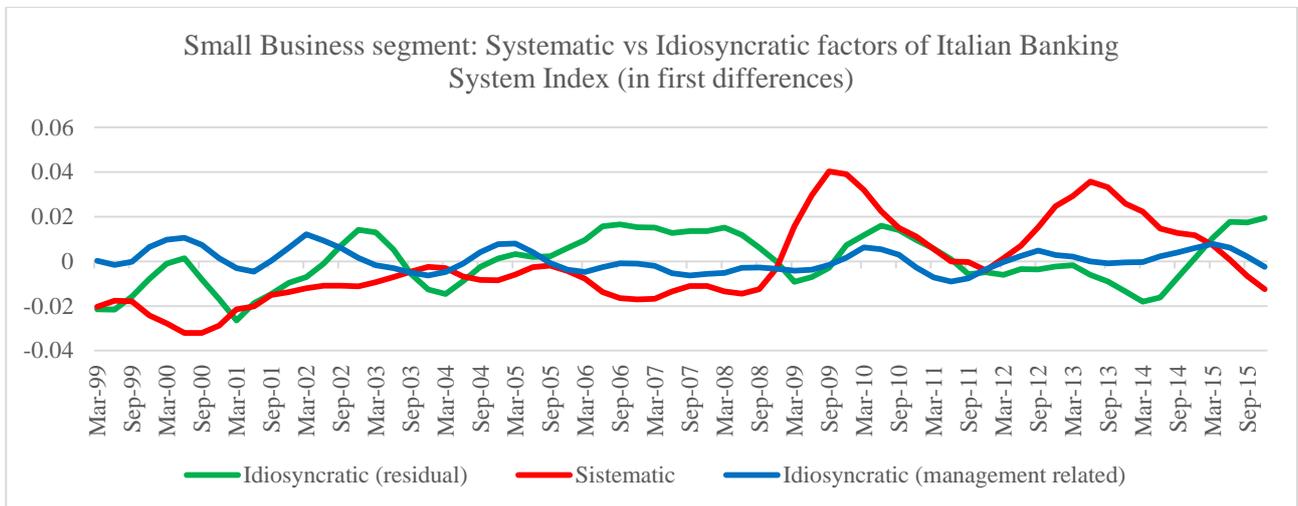
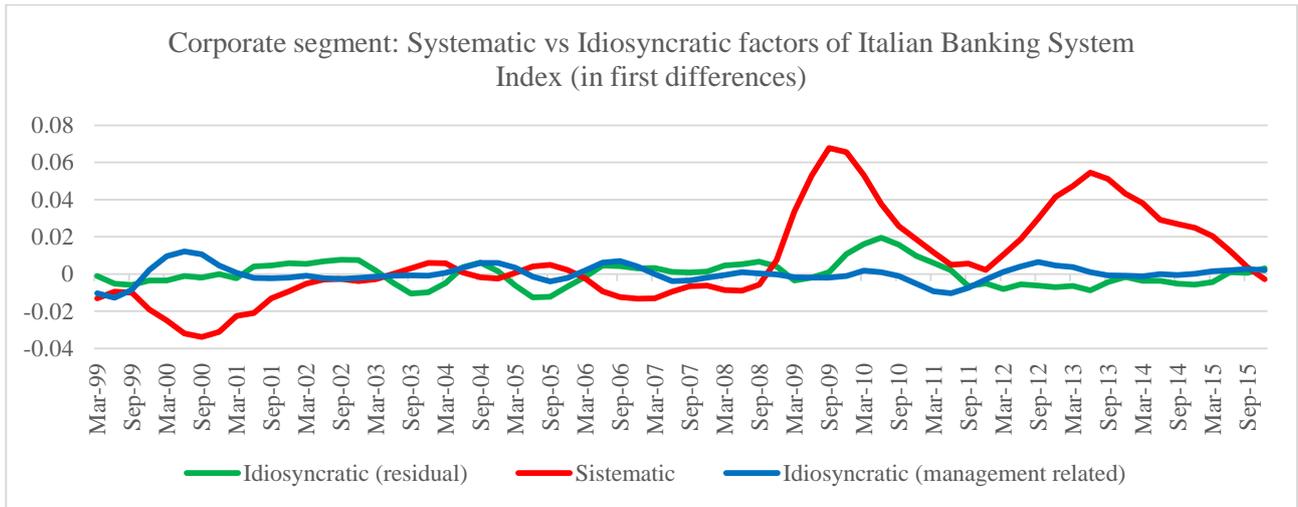
Dependent Variable: Probit of Small Business Default Rates in first difference			
Regressors	Lag	Coef. Estim.	p-Value
Intercept		0.0090	0.0360
GDP IT	0	-0.5528	<.0001
HICP IT	0	-0.3995	0.0380
UR EU	1 Year	1.0790	<.0001
LR IT	1 Year	0.7180	0.0015
Idiosyncratic (Management)	1 Year	-0.4127	0.0024
Adj. R-Square	69%		

Dependent Variable: Probit of Retail Default Rates in first difference			
Regressors	Lag	Coef. Estim.	p-Value
Intercept		0.0005	0.8085
GDP IT	0	-0.1746	0.0637
HPI RE IT	0	-0.1871	<.0001
UR EU	1 Year	0.4266	0.0814
LR IT	1 Year	0.3004	0.1414
Idiosyncratic (Management)	1 Year	0.0925	0.4732
Adj. R-Square	68%		

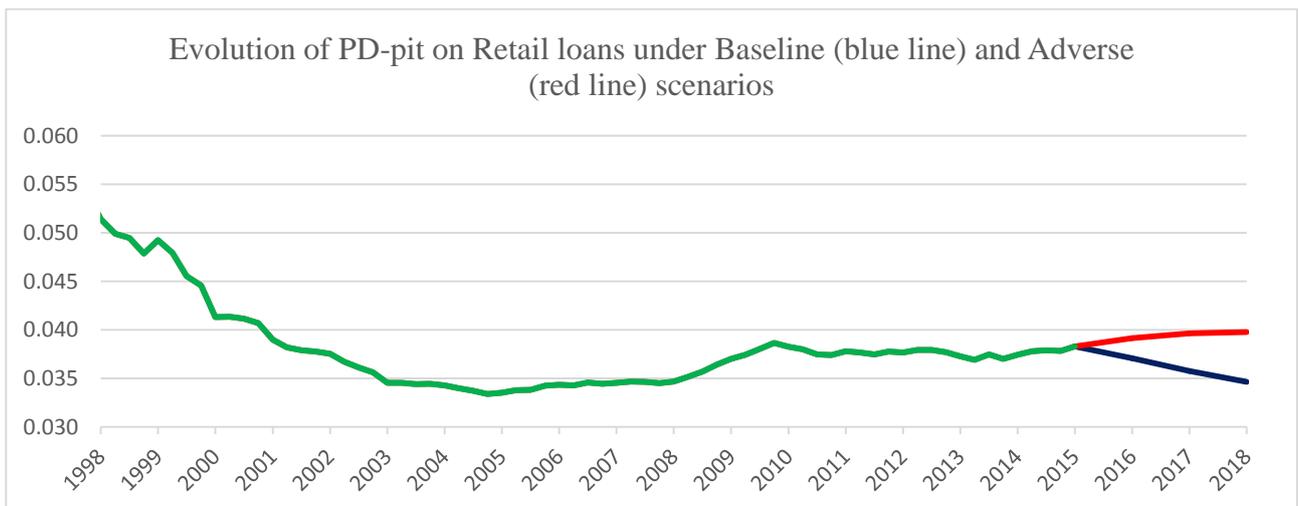
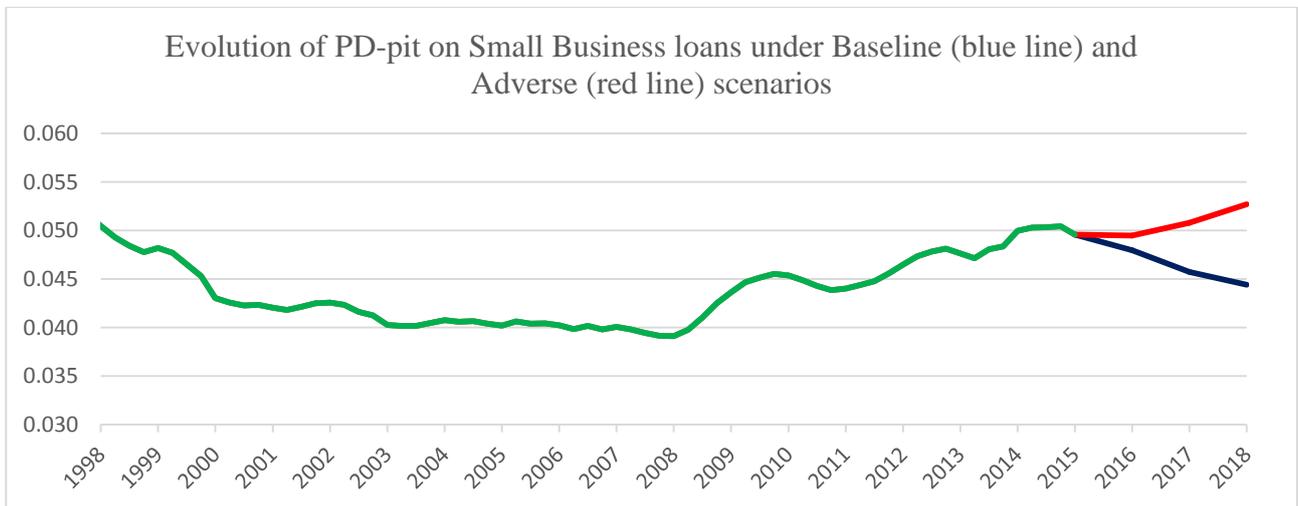
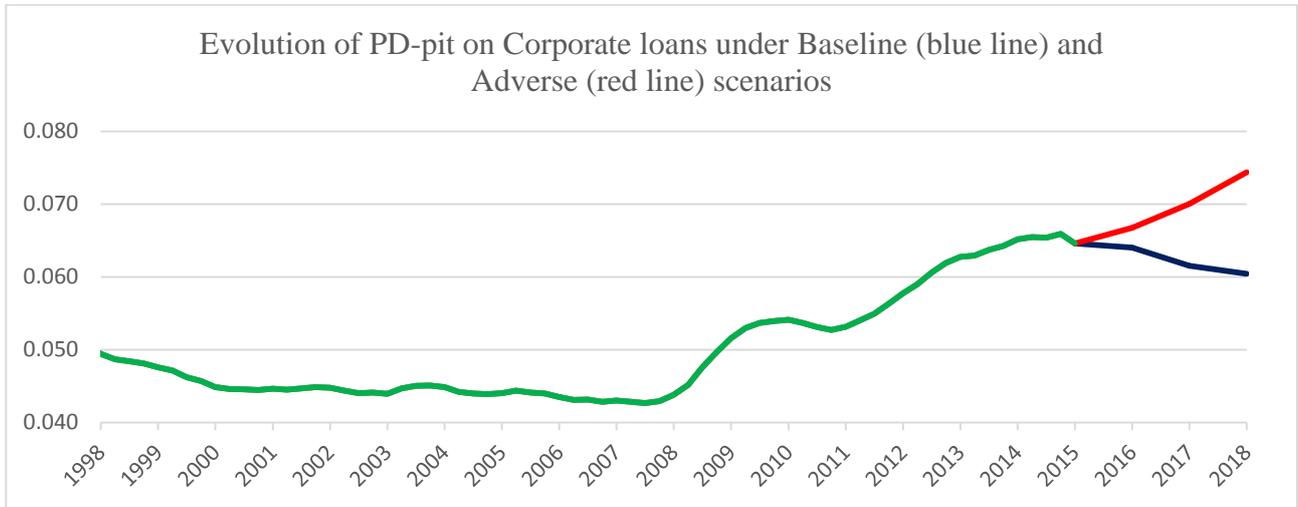
Panel 2: Observed vs Predicted values of Probit of Corporate, Small Business and Retail Default Rates in first differences from March 1999 to December 2015



Panel 3: Systematic factor vs Idiosyncratic factors of Probit of Corporate, Small Business and Retail default rates in first differences from March 1999 to December 2015



Panel 4: Observed default rates (green line) and Baseline (blue line) vs Adverse (red line) PD-pit estimates of Corporate, Small Business and Retail loans of Aggregate Italian Banking System



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