Bailout Policies, bank competition and bank risk-taking in crisis periods

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Abstract: *This paper analyzes the impact of government bailout policies on the risk of the banking sector in OECD countries between 2005 and 2013. We use the 2007-2008 financial crisis as an exogenous source of change in bailout expectations. First, in line with the moral hazard hypothesis, we verify that financial institutions with high bailout expectations assume higher risks than others. Second, we find that, in normal times, rescue guarantees to large financial institutions distort competition in the sector and increase the risk of the other institutions. However, during the recent financial crisis, increases in the rescue expectation of competitors of an institution, to the extent that they represent a reduction in its chance of bailout, decrease its risk taking. Additionally, in a crisis period, it is also evident that the deterioration in the countries' sovereign capacity to bailout banks is associated with lower risk taking; on average, the increase in risk taking is higher in countries with a better fiscal condition.*

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1 Introduction

The expansion of the banking safety net is often used in the management of financial crises (Schich, 2012). However, there is still no consensus in the literature about the *ex-ante* and *ex-post* effects of increased guarantees to bank liabilities on bank risk taking. Under the hypothesis of market discipline, the safety net reduces the incentive for depositors, creditors, and shareholders to monitor the behavior of financial institutions (FIs), thereby leading to an increase in their risk level (Flannery, 1998). In this line, Dam and Koetter (2012) find evidence that increases in government rescue expectations are associated with greater risk taking. Gropp, Hakenes, and Schnabel (2011) find that government bailout policies distort competition, i.e., because banks with implicit or explicit rescue guarantees fundraise at lower costs and, therefore, can lend at lower rates and retain the best borrowers, they end up pressuring the profit margin of other banks, that are forced to lend to riskier clients.

In turn, the charter value hypothesis states that government guarantees make protected banks to act more conservatively (Keeley, 1990). Bailout expectations are higher for a select group (too big to fail, or systemically important) of banks. Gropp, Hakenes and Schnabel (2011) shows that competition distortions causes unprotected banks to take even more risk.

A number of recent studies examine the risk taking of banks and the banking safety net in periods of financial stress. Damar, Gropp, and Mordel (2012) verify that, while banks that have a greater bailout probability take more risks in normal periods, the banks that *ex-ante* had a low probability of being rescued assume greater risks than those considered to be potentially secured in periods of crisis. This result is partially confirmed by Marques, Correa, and Sapriza (2013), who suggest that the relationship between rescue probability and bank risk is positive, both in the analysis interval called normal, between 2003 and 2004, and in the crisis interval, between 2009 and 2010. Duchin and Sosyura (2014) examine the impacts of the Troubled Asset Relief Program (TARP) on banking risk, and find evidence of an increase in moral hazard. Along these lines, Acharya, Anginer, and Warburton (2013) find that, there

is a positive relationship between risk and cost of funding for small and medium FIs, whereas for large FIs, the cost of funding is almost unrelated to its risk. Oliveira, Schiozer, and Barros (2015) show that, in 2008, there was a significant migration of deposits to systemically important institutions in Brazil. This migration is not a reflection of bank fundamentals, but instead of the depositors' perception that large institutions had an implicit rescue guarantee. Their results imply that this too-big-to-fail policy distorts competition by allowing these banks increased access to liquidity when it is most scarce. Taken together, these studies suggest that the incentives for bank risk taking (i.e., the variables that affect risk taking) in normal times may be different from those in a financial crisis.

As opposed to the market discipline hypothesis, the charter value hypothesis suggests that the safety net provides a reduction in bank risk. Keeley (1990) analyzes the effects of banking competition in the US starting in the 1950s and concludes that regulatory restrictions on new entrants and competition make the charter value significant, which reduces the appetite for risk.

Based on what has been described above, one can infer that the financial soundness of the government is an important element in explaining risk taking in the financial system because it makes a possible bailout more or less credible. For example, for the interval from 1991 to 2008, Demirgüç-Kunt and Huizinga (2013) find a significant negative relationship between the stock prices of banks and the countries' fiscal deficit and a positive relationship between the banks' CDS and the dependent variable.

This study investigates whether the bailout expectation of a given bank affects the level of risk of the protected bank and its competitors, both in normal times and during financial crises. Since the theory suggests that bailout expectations may either increase or decrease bank risk, the relative magnitude of the forces that drive bank risk may be different in normal times and during the financial crisis. Beyond that, individual bank bailout expectations may be altered during a financial crisis for at least two reasons. First, governments and regulators may have a different assessment of the possible macroeconomic effects of a bank's demise during a financial crisis as compared to a bank failure in normal times. A bank failure in turbulent times might trigger bank runs to other healthy banks, resulting in the deepening of a

systemic crisis. As such, it is quite reasonable that authorities will be more likely to rescue banks during a financial crisis than in normal times.

On the other hand, agents may reassess their expectations of bank bailouts during a financial turmoil because countries may have limited financial capacity to rescue banks. During a financial crisis, it is likely that many banks will be in financial distress and, therefore, authorities will have to establish priorities about which banks should be rescued, based on the possible consequences of their failures. Systemically unimportant banks may be perceived as being particularly less likely to be bailed out during a financial turmoil. Therefore, the bailout expectation of competitor banks may be an important determinant of bank risk taking. Likewise, countries with greater financial capacity are more likely to inject large amounts of resources into the financial system (i.e., rescue financial institutions), than countries with smaller financial capacity. In financially constrained countries, the bailout expectation of competitors may be even more important for bank risk taking.

We exploit the exogenous variation in the bailout expectations of banks caused by the global financial crisis to assess the effects of these expectations on bank risk taking. We also benefit from the cross sectional variation in the banks' systemic importance and in countries' financial capacity (that results in heterogeneous variation in bailout expectations across countries) to assess this effect. More specifically, this paper aims to answer the following research questions:

1. Is there a relationship between the risk level of a bank and the rescue expectation of its competitors? If so, does it change in periods of banking crisis?

2. Is there a relationship between the financial capacity of a country and the risk of local banks? If so, does it change in periods of banking crisis?

In line with the moral hazard hypothesis, we find that increases in the bailout expectation of a financial institution are associated with greater risk taking. In addition, the rescue expectation of the competitors of an institution also increases its risk taking in normal times. However, during a financial crisis, this effect is mitigated. Therefore, in normal periods, there is a predominance of the channel through which increases in the protection of competitors distort competition, reduce profit margins, and increase the risk taking of small institutions. During

crises, however, increases in the bailout expectations of competitors are associated to lower risk taking. This result is consistent with the rationale that, the higher the bailout expectation of a bank's competitors, the lower the bank's relative importance in the financial system and thus the lower its prospect of eventually being rescued in a crisis. Therefore, there is incentive for greater conservatism among banks with a lower probability of being rescued.

Similarly, we observe that the association between sovereign CDS spread and banking risk is modified during the crisis. While in normal times, increases in the sovereign CDS are linked to higher levels of banking risk (which may stem from an increase in the risk of the banks' assets directly), this effect is reversed during the crisis. That is, in line with the moral hazard hypothesis, an increase in the CDS (a reduction in the financial ability of a country to undertake bank bailouts), creates incentives for less risk taking in banks. Finally, corroborating previous results, we find that, during the crisis, institutions located in countries with lower ex-ante bailout capacity (i.e. higher pre-crisis CDS spreads) have, on average, an increase in their risk level that is lower than that in countries with low pre-crisis CDS spreads.

We add to Gropp, Hakenes, and Schnabel (2011) by analyzing a crisis periods an testing whether, in a period of financial stress, their observed effect is homogeneous among banks. In turn, Damar, Gropp, and Mordel (2012) and Marques, Correa, and Sapriza (2013) also assess banking risk in the pre- and post-crisis period, but they examine the direct effect of bailout expectation on bank risk taking, whereas we also examine the effect of bailout expectations on the banks' competitors.

The remainder of this paper is as follows: the second section discusses the hypotheses and methodology, followed by a description of the data and variables in the third section. Section 4 discusses the estimation results and the last section concludes.

2 Identification strategy

2.1 Hypotheses and Models

The criteria used by governments in their decision to rescue a certain bank are not fully transparent to investors and bankers.[1] Typically, investors rely on the assessment that factors such as size, interconnectedness, type of ownership, and sometimes political connections raise the chance of bank bailout (Brewer and Jagtiani, 2013). In the following excerpt, Moosa (2010) examines the 2007 crisis and confirms this subjectivity:

> […] Let us examine the recent record to see why the bailout practice has indeed been cherry picking. In 2008 Lehman Brothers was allowed to fail (by filing for bankruptcy) but Merrill Lynch and Bear Stearns were saved from bankruptcy by government-assisted and partly-financed mergers with Bank of America and JP Morgan, respectively. Citigroup and AIG (and Goldman Sachs indirectly) were saved by massive direct injection of cash from the U.S. Treasury. Yet in 2009 alone more than 150 other U.S. banks were allowed to fail. The TBTF status was given to Continental Illinois in the 1980s, but not to Drexel Burnham Lambert in the 1990s. Consider also the case of the hedge fund LTCM, which was saved by the intervention of the New York Fed that engineered (with a lot of arm twisting, some would say) a very attractive deal for the failed management, but another fund (Amaranth) that was twice as big was allowed to go down.

Thus, under a stress situation in the banking system, it is reasonable to expect that the lower the market share of a bank in the financial system, the lower its chance of eventually being bailed out, because its failure is less likely to pose a systemic risk.

We follow Gropp, Hakenes, and Schnabel (2011), and measure the distortion on competition caused by the protection to competitor banks as the market share of insured competitors (MSIC), whose computation is detailed below. In line with the theory of moral hazard and considering the aspects previously presented, it is possible that, in times of crisis, an increase in the MSIC of a bank contributes to reducing its risk taking. Or, similarly, a decline in the MSIC of a bank, to the extent that it puts the bank in a better position to obtain government assistance, is linked to greater risk taking. Thus, the channel through which increases in the MSIC contribute to distorting competition and increasing risk taking, as noted by Gropp, Hakenes, and Schnabel (2011), would be mitigated or even be overcome by this new channel, which is designated here as the *channel of relative rescue expectation*. Therefore, based on

[1] A few countries (such as Japan) have an explicit list of systemically important banks, but these are the exception and not the norm.

the discussion above, we hypothesize that, in normal times, increases in the MSIC are related to greater bank risk taking. However, in a period of crisis, this effect is mitigated.

To test the extent to which bank risk respond to MSIC, we exploit an exogenous variation in the perception of bailout probabilities caused by the global financial crisis, and estimate the following model:

$$\text{Z-score}_{i,t} = \beta_0 + \beta_1 \text{MSIC-}_{i,t} + \beta_2 \text{Crisis}_{t,j} + \beta_3 \text{MSIC-}_{i,t} * \text{Crisis}_{t,j} + \beta_4 X^{T}_{i,t} + \varepsilon_{i,t} \qquad (1)$$

The main dependent variable is a measure of bank risk defined in line with Damar, Gropp, and Mordel (2012); Soedarmono, Machrouhb, and Tarazi (2013); and Marques Correa, and Sapriza (2013), the logarithm of the Z-score, defined as in equation (2).

$$Z\text{-}score = \frac{\text{ROA} + \text{Capital to Assets Ratio}}{\sigma(\text{ROA})} \qquad (2)$$

The market share of insured competitors (MSIC) variable is based on Gropp, Hakenes, and Schnabel (2011), which aims to capture the association between the market share (total assets of the institution over the total assets of its country's financial system) of each bank and the rescue probability of its competitors. The formula for calculating the MSIC of bank *i* in country *j* in period *t* is given by equation (3):

$$MSIC_{-i,j,t} = \sum_{k \neq i}^{N_j} p_{k,j,t} \, \frac{a_{k,j,t}}{A_{k,t}} \qquad (3)$$

Where $a_{k,j,t}$ is the total assets of competitor bank *k* in country *j* in period *t*, and $A_{j,t}$ is the sum of total assets of the banking system in country *j* in period *t*. The competitor bank k probability of bailout in period *t* is given by $p_{k,j,t}$. This probability is based on the support rating of the bank, estimated by Fitch Rating and refers to the possibility of bailout if the bank is about to be unable to meet its financial commitments. Ratings from 1 to 5 are given to each bank, and the higher the rating is, the lower its chance of receiving financial bailout if needed. This support can come from both shareholders and governmental authorities in the countries where the banks are headquartered. With respect to classification criteria, Fitch reports that

the classification is based on the financial ability and the potential guarantors' propensity to bailout.

Similar to Gropp, Hakenes, and Schnabel (2011), each rating is associated with a rescue probability. Because not all banks in the database had support indicators estimated by Fitch Rating – in most cases, due to their low representativeness – such institutions are assigned a support level of 5, which equals a rescue probability of zero. Table 1 details the support rating variable and the respective adopted rescue probability.

Table 1 - Description of support rating by Fitch and assignment of bail-out probabilities

Support rating	Description by Fitch	Assigned bail-out probability
1	A bank for wich there is an extremely high probability of external support. The potetnital provider of support is highly rated and has a very high propensity to provide support to the bank in question.	1
2	A bank for wich there is a high probability of external support. The potetnital provider of support is highly rated and has a high propensity to provide support to the bank in question.	0.9
3	A bank for wich there is a moderate probability of support because of uncertainties about the ability or propensity of the potential provider to do so	0.5
4	A bank for wich there is a limited probability of support because of significant uncertainties about the ability or propensity of any possible provider of support to do so	0.25
5	A bank for wich external support, although possible, cannot be reliable upon. This may be due to a lack of propensity to provide support or to very weak financial ability to do so.	0

$Crisis_{t,j}$ is a dummy for the financial crisis period. It is a dummy that takes the value of 1 for crisis years in the countries that were deeply affected by the global financial crisis. We use Laeven and Valencia's (2012) assessment to define whether a certain country was hit by the financial crisis.[2] Therefore for countries unaffected by the global financial crisis, this dummy assumes value 0 for the whole sample period; for countries whose financial systems were affected by the crisis, it takes value 1 from 2008 to 2010, and 0 for all other years.

[2] Their assessment relies two criteria to consider that a country's financial system was affected by the crisis: a) strong signs of difficulties in the financial system, indicated, for example, by bank runs or bank liquidations; and b) significant government intervention in the banking system in response to the losses suffered by the institutions.

The bank-specific control variables include the following: size, the bank's own rescue probability, and liquidity. The first variable is estimated by the natural logarithm of the total assets of the institution. The support probability is based on Fitch's support rating, as shown in Table 1. To mitigate possible endogeneity between the rescue probability and the institution's degree of risk-taking, the bank's own bailout probability enters the model with a one-year lag. The liquidity variable is defined by the ratio between liquid assets and short-term liabilities.

Regarding the macroeconomic environment of the analyzed countries, the following controls are used in robustness checks: Gross Domestic Product (GDP) growth, the relationship between total bank credit and GDP, GDP per capita and sovereign Credit Default Swap (CDS) spread. Additionally, the Herfindahl concentration index is used as a proxy for the bank competition level in the countries. The GDP per capita variable is used as a proxy for the sophistication of the financial system. Considering that differences related to the impact of economic growth on bank risk may be due, for example, to some lagged effect, this study includes this variable with a lag of 1 and 2 years. In line with Anginer, Demirgüç-Kunt, and Zhu (2014), the domestic credit index to the private sector over the GDP aims to control the importance of the financial system to the countries' economies. We use the 5-year CDS spread, in US dollars, as a measure of the countries' sovereign risk. We use average trading spread calculated on the last business day of the year.

The interpretation of the regression coefficients is straightforward: β_1 captures the average effect of the distortion in competition caused by the protection of competitor banks, β_2 is the average effect of the global financial crisis on bank risk (in the countries affected by the crisis), and β_3 is the main coefficient of interest, and is interpreted as the marginal effect of the financial crisis (i.e. an increase in bailout expectation for large banks) on the relationship between the MSIC and bank risk.

First, the regression is estimated using the ordinary least squares method and includes bank fixed effects, which aims to capture unobserved heterogeneity that is relatively stable over time, such as quality of management, corporate governance and ownership structure. Since market shares and bailout probabilities are relatively stable through time, bank fixed effects could be capturing a large part of the effect of MSCI on bank risk. To consider this issue and

still address the problem of unobserved heterogeneity among countries, regressions with country fixed effects are performed. Thus, it is sought to control the effects that omitted variables, such as regulation and supervision in the countries, may have on the risk taking of local banks. Because the individual error term can contain common elements in all periods of analysis, we use robust clustered standard errors at the bank level.

In addition, studies suggest that the financial capacity of the country (i.e., fiscal condition) is an effective constraint to the depth and breadth of bank bailouts (e.g., Alter and Schüler, 2012; Demirgüç-Kunt and Huizinga, 2013; Duttagupta and Cashin, 2011; Schich and Lindh, 2012). We further investigate the effect of bailout expectation on risk taking by looking at a country's ability to rescue its financial institutions in case of necessity. In a study that assesses the causes and consequences of banking crises in several countries between 1980 and 2002, Demirgüç-Kunt and Detragiache (2005) indicate that economies with a deteriorated fiscal position, inflation, and high real interest rates are associated with higher bank instability. One reason for this finding perhaps lies in the fact that deficit governments choose to postpone measures for strengthening the FIs, as Lindgren, Garcia, and Saal (1996 apud Demirgüç-Kunt and Detragiache, 1998) suggest:

> ... supervisors often are prevented from intervening in banks because this would bring problems out in the open and "cause" government expenditure. Typical justification for inaction are that there is "no room in the budget" or that fiscal situation is "too weak" to allow for any consideration of banking problems.

Additionally, Demirgüç-Kunt and Detragiache (2005) emphasize that the variable GDP per capita, considered an indicator of the quality of a country's institutions, is negatively correlated with systemic risk, which is corroborated by Berger, Klapper, and Ariss (2009). In this sense, considering the strong negative relationship between GDP per capita and the sovereign risk rating (Cantor and Packer, 1996), it is reasonable to assume that the link between country risk and banking risk is also affected by factors that are often related to the GDP per capita variable. That is, in line with Kane (2000), Angkinand and Wihlborg (2010),

and La Porta, Lopez-de-Silanez, and Shleifer (2002), it may be said that, in less developed countries, there are lower respectability of contracts, higher corruption levels, and lower governance, suggesting greater risk taking by the FIs. As a result, it is assumed that, in normal times, the CDS spread variable is related to greater risk taking.

By contrast, in periods of crisis, when there is a greater concern for the countries' rescue ability, it is possible that the relationship between sovereign CDS and bank risk is changed because the bankruptcy of a FI occurs due to insolvency or illiquidity or because the government or the lender of last resort ultimately allows bankruptcy to happen, either because it is not convenient to do the bailout or because it has no financial ability to do so. Therefore, as a deterioration in the country's financial capacity suggests a lower guarantor credibility and strength, an increase in the sovereign risk tends to impact the risk taking of FIs (Schich and Lindh, 2012).

In Correa, Lee, Sapriza, and Suarez (2014), for example, the link between sovereign risk and rescue prospects is observed by analyzing the relationship between stock prices in the banking sector and the lowering of the countries' credit risk classification. Examining banks from 37 countries between 1995 and 2011, it is found that downgrades in the country risk rating have a strong negative effect on the price of the evaluated papers. In addition, the stronger the bailout expectation of the FI, the stronger this relationship is. In the excerpt below, Demirgüç-Kunt and Huizinga (2013) find a similar result and highlight its validity especially during the crisis:

> Especially at a time of financial and economic crisis, there are doubts about countries' ability to keep their largest banks afloat. For 2008, we present evidence that the share prices of systemically large banks were discounted relatively more on account of systemic size in countries running large fiscal deficits. This is evidence that systemic banks located in countries with stressed public finances saw their contingent claim on the financial safety net reduced relatively more in 2008, which is evidence that they have grown 'too big to save'.

Thus, considering the different channels through which the sovereign risk rating and the bank bailout prospect are related and, moreover, how they affect the risk of FIs, we hypothesize

that, in normal periods, increases in CDS spreads are related to greater bank risk taking. However, in periods of crisis, this effect is reducued.

Below, the general model to be estimated is presented:

$$Z\text{-score}_{i,t} = \beta_0 + \beta_1 CDS_{i,t} + \beta_2 MSIC_{i,t} + \beta_3 Crisis_t + \beta_4 CDS_{i,t*} Crisis_t + \beta_5 MSIC_{i,t*} Crisis_t + \beta_6 X^T_{i,t} + \varepsilon_{i,t}$$

If it is confirmed that, in times of crisis, increases in the CDS spread contribute to reducing the bank risk taking, then it will be tested whether this effect is homogeneous between FIs located in countries considered to have a high credit risk and FIs located in low credit risk countries. Considering the hypothesis that a high sovereign CDS spread is associated with a lower ability to perform occasional bank bailouts, it would be consistent with previous conjectures to hypothesize that, in periods of crisis, FIs located in countries with low CDS spreads will increase their risk more (or decrease it less) than FIs located in countries with high CDS spreads.

We test this hypothesis using a difference-in-differences model. In this case, based on the availability of CDS spreads data from 2007, countries are divided into 2 groups: low credit risk, i.e., the 8 countries (Austria, Belgium, France, Germany, Slovenia, Switzerland, United Kingdom, United States) with the lowest CDS spreads in 2007; and high credit risk, i.e., the 18 countries with the highest CDS spreads in 2007. We estimate the following model:

$$Z\text{-score}_{i,j,t} = \beta_0 + \beta_1 LCDS_j + \beta_2 Crisis_{j,t} + \beta_3 LCDS_{j*} Crisis_{j,t} + \varepsilon_{i,t} \quad (5)$$

The $LCDS_j$ variable is a dummy variable that takes the value of 1 if the FI is in a country *j* with low CDS spread in 2007; otherwise, the value is 0. The Crisis variable is defined above. Thus, the hypothesis is corroborated if the coefficient of the interaction between LCDS and Crisis is negative and significant.

3 Data and summary statistics

3.1 Data and sample

Bankscope is the main source of data for this work. The annual financial and unconsolidated statements of commercial banks, savings banks, cooperative banks, mortgage banks, and government credit institutions operating in the OECD (Organization for Economic Cooperation and Development) member countries were collected from this database. For simplicity, we call all these financial institutions "banks", unless explicitly mentioned otherwise. The criterion for choosing this set of institutions is the fact that commercial banks represent a significant part of the banking sector in these countries whereas the other banks are their main competitors (Gropp, Hakenes, and Schnabel, 2010).

Although the period of analysis extends from 2005 to 2013, the sample covers a broader range, extending from 2003 to 2013. The difference reflects the need to have a window with 3 basis dates to estimate the standard deviation of the return on assets (ROA) of the banks; that is, to calculate the z-score in year t, data between t and t-2 are used. The selection of this research range aims to analyze the pre-financial crisis period of 2007, mitigating the effects of possible distortions caused by the change in the accounting standards of FIs in the mid-2000s, that is, the adoption of the International Financial Reporting Standards (IFRS) in some of the countries in the sample (Marques, Correa, and Sapriza, 2013).

In 2005, the database was composed of 14,000 institutions. However, due to a lack of relevant data, such as the capitalization rate or return on assets of some institutions, this total was reduced to 3,337 institutions. During the overall period, the number of observations reaches 41,632 (bank-years). Table 2 shows the number of sample banks distributed by type and country in 2005 and 2013. At the beginning of the investigation, the countries with the highest number of observations are Germany, Japan, and the United States. However, there is a distinct difference between the types of banks that predominate in these countries. In the first 2 countries, cooperative banks prevail, whereas in the third, commercial banks prevail.

Macroeconomic country-level data, such as GDP growth, GDP per capita and the credit to GDP ratio are obtained from the World Bank. The 5-year sovereign CDS spreads are obtained from Bloomberg.

Table 2 - Distribution of banks by type and country

	Commercial Banks		Cooperative Banks		Mortgage Banks		Savings Banks		Government Credit Institutions		Total	
Country	**2005**	**2013**	**2005**	**2013**	**2005**	**2013**	**2005**	**2013**	**2005**	**2013**	**2005**	**2013**
Australia	1	15	0	2	0	7	0	1	1	8	**2**	**33**
Austria	53	58	95	70	9	15	51	48	2	3	**210**	**194**
Belgium	12	21	5	5	2	2	3	4	1	0	**23**	**32**
Canada	10	30	3	34	0	3	0	4	0	2	**13**	**73**
Chile	1	22	0	0	0	2	0	4	0	0	**1**	**28**
Czech Republic	12	17	0	2	2	2	0	0	2	2	**16**	**23**
Denmark	29	29	1	6	7	8	13	32	2	2	**52**	**77**
Estonia	5	7	0	0	0	0	0	0	0	0	**5**	**7**
Finaland	3	20	0	1	0	3	1	3	0	2	**4**	**29**
France	70	96	25	67	11	21	16	20	4	8	**126**	**212**
Germany	72	104	652	929	32	37	389	494	11	27	**1.156**	**1.591**
Greece	7	8	1	1	0	0	0	0	0	0	**8**	**9**
Hungary	17	18	0	1	2	3	0	0	2	2	**21**	**24**
Iceland	0	4	0	0	0	1	6	2	0	0	**6**	**7**
Ireland	5	8	0	0	2	3	0	0	0	0	**7**	**11**
Israel	10	9	0	0	0	0	0	0	0	0	**10**	**9**
Italy	3	69	1	400	0	3	0	31	1	7	**5**	**510**
Japan	115	125	443	426	0	0	1	0	2	2	**561**	**553**
Luxembourg	41	43	2	1	0	0	0	2	0	0	**43**	**46**
Mexico	21	40	0	4	4	5	0	5	3	4	**28**	**58**
Netherlands	12	28	1	1	1	2	0	1	1	2	**15**	**34**
New Zealand	0	13	0	4	1	3	0	0	0	0	**1**	**20**
Norway	3	14	0	0	1	9	33	106	1	2	**38**	**131**
Poland	19	33	1	1	0	0	1	1	1	1	**22**	**36**
Portugal	4	18	0	2	0	1	0	4	0	0	**4**	**25**
Slovak Republic	6	9	0	0	0	2	1	2	2	2	**9**	**15**
Slovenia	11	13	0	2	0	0	0	1	0	1	**11**	**17**
South Korea	1	13	1	1	0	0	0	3	1	3	**3**	**20**
Spain	6	30	6	31	1	0	5	14	1	2	**19**	**77**
Sweden	11	19	0	0	5	8	43	51	3	2	**62**	**80**
Switzerland	108	92	6	7	4	4	179	187	24	26	**321**	**316**
Turkey	12	26	0	0	0	0	0	0	1	3	13	29
United Kingdom	63	100	0	1	29	49	0	3	2	1	**94**	**154**
United States	402	396	11	6	5	4	8	8	2	2	**428**	**416**
Total	**1.145**	**1.547**	**1.254**	**2.005**	**118**	**197**	**750**	**1.031**	**70**	**116**	**3.337**	**4.896**

3.2 Descriptive statistics

Table 3 shows that during the pre-crisis period the average standard deviation of the ROA is equal to 0.27. However, during the crisis the ratio increases to 0.43. Similarly, there is also a significant increase in the banks' liquidity. Median liquidity is virtually unchanged, indicating that some (but not all) banks have significantly held on to liquidity. There is no relevant modification in the capital to assets ratio between the crisis and the time that precedes it. The average return on assets (ROA) during the crisis is 0.33%, less than half of the 2005-2007 average. As a result of the reduction of the ROA and the increase of its variability, there is a significant drop in the z-score during the crisis, with the average risk in the banking sector going from 0.26 to -0.36. Apparently this decrease has affected a large part of the banks analysed, given that the standard deviation of the z-score during the crisis remains similar to the pre-crisis period.

Table 3 - Banks' Descriptive Statistics

Variables	2005-2007				2008-2010				2005-2013			
	Observations	Mean	Median	Std. Dev	Observations	Mean	Median	Std. Dev	Observations	Mean	Median	Std. Dev
Standard Deviation of ROA	12.441	0,27	0,07	1,09	14.335	0,43	0,12	2,11	41.632	0,36	0,09	2,23
Liquidity (%)	12.237	21,02	15,99	18,47	14.088	27,59	16,48	54,71	40.896	20,65	15,38	18,28
MSIC	12.441	0,58	0,58	0,17	14.335	0,59	0,56	0,14	41.632	0,57	0,54	0,15
Capital to Assets Ratio (%)	12.441	9,20	6,86	10,09	14.335	9,24	7,09	11,94	41.632	9,47	7,37	11,93
Probability of Support	12.441	0,04	0,00	0,20	14.335	0,05	0,00	0,21	41.632	0,05	0,00	0,21
ROA (%)	12.441	0,70	0,38	2,18	14.335	0,33	0,26	3,32	41.632	0,46	0,30	2,65
Total Assets (US$ billions)	12.441	16,40	0,89	110,00	14.335	22,60	1,05	146,00	41.632	21,50	1,04	140,00
Z-Score*	12.441	0,26	0,13	1,59	14.335	-0,36	-0,40	1,56	41.632	0,02	-0,05	1,63

*Natural log of Z-score winsorized at 1%

According to the Table 4, the average CDS spread rises from 5.35 to 71.94 between the pre-crisis and crisis period. The data show that, on average, countries have negative GDP growth during the crisis. However, from 2010 on, a positive GDP growth is observe (unreported).

Table 4 - Countries' Descriptive Statistics

Variables	2005-2007				2008-2010				2005-2013			
	Observations	Mean	Median	Std. Dev	Observations	Mean	Median	Std. Dev	Observations	Mean	Median	Std. Dev
CDS Spread	9.876	5,35	3,63	5,29	13.381	71,94	55,71	71,94	37.670	68,52	39,59	183,50
Herfindahl Index	12.441	0,12	0,07	0,11	14.335	0,11	0,07	0,08	41.632	0,10	0,07	0,08
GDP Per capita (US$ th)	12.441	40,87	37,71	11,79	14.335	45,43	41,72	14,36	41.632	45,05	43,09	15,15
GDP Growth (%)	12.339	2,75	2,99	1,28	14.335	-0,29	0,49	3,50	41.632	1,10	1,61	2,67
Domestic Credit (%GDP)	12.339	105,46	101,83	32,39	13.939	112,89	106,89	37,403	40.478	109,52	105,04	35,52

4 Results

4.1 Market Share of Insured Competitors

In column 1 of Table 5, the effects of the MSIC variable, the crisis dummy variable, and the interaction between them on banking risk are tested. In line with what was indicated by Gropp, Hakenes, and Schnabel (2011), we find that, in normal times, the MSIC coefficient is significant and negative. Thus, an increase of 0.1 in the MSIC is associated with an average decrease of 12.15% in the Z-score. Therefore, considering the reduction in the explained variable, there is evidence that increases in the MSIC stimulate greater risk taking. The crisis dummy is also linked to increases in bank risk, and it signals that, for a bank with a zero MSIC, there is an average decrease of 47.8% ($e^{-0.651} - 1$) in the Z-score. Unlike the effect of the MSIC variable when taken alone, its interaction with the crisis dummy has a positive and significant impact on the risk level. Thus, although in periods of crisis the net effect of the MSIC variable on the Z-score is still negative, there is an important reduction in its effect. That is, in the banking turbulence phase, for an increase of 0.1 in the MSIC, there is an average decrease of 8.64% in the Z-score. Thus, there is evidence corroborating hypothesis 1: due to the governments' limited ability to bailout financial institutions in times of crisis, an increase in the MSIC may have a marginal effect that contributes to the decline in risk taking.

Table 5 - Bank Risk Taking and MSIC

Variables	Z-score					
	(3.1)	(3.2)	(3.3)	(3.4)	(3.5)	(3.6)
MSIC	-1,215***	-1,484***	-1,829***	-0,650***	-1,063***	-1,495***
	(0,130)	(0,142)	(0,161)	(0,122)	(0,126)	(0,148)
Crisis	-0,651***	-0,700***	-0,557***	-0,658***	-0,702***	-0,570***
	(0,072)	(0,076)	(0,085)	(0,069)	(0,072)	(0,080)
MSIC * Crisis	0,351***	0,485***	0,409***	0,341***	0,480***	0,479***
	(0,114)	(0,119)	(0,132)	(0,109)	(0,114)	(0,124)
Assets		-0,287***	-0,396***		0,005	0,006
		(0,038)	(0,052)		(0,009)	(0,009)
Prob. Sup.		-0,441***	-0,361***		-0,512***	-0,570***
		(0,137)	(0,135)		(0,078)	(0,083)
Liquidity		-0,081***	-0,106***		-0,160***	-0,163***
		(0,018)	(0,020)		(0,016)	(0,017)
Herfindahl Ind.			0,831***			0,771***
			(0,254)			(0,246)
GDP per capita			0,599***			0,105
			(0,120)			(0,100)
GDP Growth $_{t-1}$			0,018***			0,026***
			(0,003)			(0,002)
GDP Growth $_{t-2}$			-0,030***			-0,032***
			(0,002)			(0,002)
Banking Credit (%GDP)			-0,005***			-0,005***
			(0,001)			(0,001)
Constant	0,848***	5,315***	1,284	0,527***	1,147***	0,815
	(0,075)	(0,574)	(1,162)	(0,071)	(0,156)	(1,100)
Nº Clusters	5.639	5.346	4.881	34	34	34
Nº Fixed Effects	5.639	5.346	4.881	5.639	5.346	4.881
Nº Observations	41.632	38.084	34.731	41.632	38.084	34.731
F	275,800	126,960	104,920	296,630	144,560	118,840
R^2	0,597	0,610	0,619	0,210	0,235	0,249
R^2 Adjusted	0,534	0,546	0,557	0,209	0,235	0,248

Column (3.1) to (3.3) report results for FI fixed effects regressions. Column (3.4) to (3.6) report results for country fixed effects regressions Standard errors are reported in parentheses below their coefficient estimates and are adjusted for both heteroskedasticity and within correlation clustered at the FI level. ***, **, * indicate 1%, 5%, and 10% significance, respectively

In column 2, the control variables are added at the FI level. In line with Boyd and Runkle (1993), it is found that higher total assets are related to higher risk. Consistent with Dam and Koetter (2012), the coefficient of the support probability variable is negative and significant. Therefore, there is support for the moral hazard theory. Similarly, increases in liquidity are also linked to a higher risk level. In this sense, the hypothesis that more aggressive banks hold assets of greater liquidity in their portfolios, for example, to respond to margin calls or a run to their liabilities in a timely manner (Marques, Correa, and Sapriza, 2013) is corroborated.

By including controls at the country level in the regression, as column 3 highlights, we find that the Herfindahl index has a positive and significant coefficient, which suggests that more concentrated banking systems are related to increased robustness. Therefore, there is evidence

that is consistent with the charter value hypothesis. Similarly, increases in the GDP per capita variable are also associated with lower risk taking.

With respect to the effects of economic growth on bank risk taking, it is observed that, depending on the lag period, it is possible to have very different results. That is, when the explanatory variable is used with a one-year lag, a 1% increase in economic growth is reflected in lower risk taking, with an increase of 0.18% in the Z-score. By contrast, when a 2-year lag is applied, the response is a decrease of 0.32% in the Z-score. Thus, there is evidence pointing to the existence of the effect known in the literature as boom and bust (e.g., Hardy and Pazarbasioglu, 1998; Schularick and Taylor, 2009).

More importantly, the addition of bank-level and country-level control variables to our specification does not materially alter the coefficients of our variables of interest β_1, β_2 and β_3.

The columns 4 to 6 of Table 5 report the regressions with country fixed effects instead of bank fixed effects. Although the magnitude of β_1 is slightly reduced as compared to the counterpart regressions with bank fixed effects, our inferences remain the same.

4.2 Credit Default Swap Spread

In column 1 of Table 6, we test our second hypothesis that the country CDS spread is associated with an increase in bank risk taking in normal times, but less so during financial crises. The regression is estimated using bank fixed effects. First, we find that, in normal times, the CDS coefficient is significant and negative. Thus, an increase of 100 bp in the CDS is associated with an increase in risk taking, with an average decrease of 2.7% in the Z-score. The negative and significant coefficient of the crisis dummy, in turn, indicates that times of stress in the financial sector are linked to increases in risk taking. Finally, unlike the CDS taken in isolation, there are signs that, in times of instability, an increase in the CDS reduces risk taking. Thus, by raising the CDS spread by 100 bp, there is an average increase of 3.0% in the Z-score. Therefore, there is evidence that a reduction in the country’s ability to pay its debts is linked to a decrease in banking risk.

In column 2 of Table 6, we add the MSCI and its interaction with the crisis dummy. Not only the inferences on the influence of the CDS spread are maintained (relative to column 1, Table 6), but also the coefficients of the added variables are very similar to the ones obtained in Table 5. Finally, adding bank and other country-level control variables (column 3) and using country fixed effects instead of bank fixed effects (columns 4 to 6) do not qualitatively alter our inferences.

Table 6 - Bank Risk Taking, MSIC and CDS

Variables	Z-Score					
	(4.1)	**(4.2)**	**(4.3)**	**(4.4)**	**(4.5)**	**(4.6)**
CDS	-0,027***	-0,040***	-0,045***	-0,030***	-0,039***	-0,046***
	(0,004)	(0,004)	(0,004)	(0,004)	(0,003)	(0,004)
MSIC		-1,586***	-2,288***		-1,145***	-1,929***
		(0,133)	(0,154)		(0,122)	(0,137)
Crisis	-0,485***	-0,848***	-0,807***	-0,506***	-0,898***	-0,883***
	(0,021)	(0,083)	(0,094)	(0,020)	(0,078)	(0,089)
CDS*Crisis	0,057***	0,085***	0,124***	0,071***	0,086***	0,142***
	(0,015)	(0,017)	(0,021)	(0,015)	(0,016)	(0,020)
MSIC*Crisis		0,591***	0,668***		0,651***	0,818***
		(0,133)	(0,151)		(0,126)	(0,151)
Assets			-0,422***			0,008
			(0,060)			(0,010)
Prob. Sup.			-0,337**			-0,607**
			(0,139)			(0,087)
Liquidity			-0,111***			-0,146***
			(0,020)			(0,017)
Herfindahl Ind.			0,851***			0,596**
			(0,296)			(0,296)
GDP per capita			0,827***			0,323***
			(0,135)			(0,114)
GDP Growth $_{t-1}$			0,016***			0,023***
			(0,003)			(0,002)
GDP Growth $_{t-2}$			-0,034***			-0,037***
			(0,002)			(0,002)
Banking Credit (%GDP)			-0,002			-0,001
			(0,001)			(0,001)
Constant	0,183***	1,094***	-0,822	0,188***	0,846***	-1,776
	(0,006)	(0,076)	(1,354)	(0,015)	(0,071)	(1,263)
Nº Clusters	5424	5424	4758	5424	5424	4758
Nº Fixed Effects	5424	5424	4758	32	32	31
Nº Observations	37.670	37.670	32.061	37.670	37.670	32.061
F	243,69	184,07	94,520	286,83	194,65	105,410
R^2	0,603	0,607	0,627	0,212	0,215	0,251
R^2 Adjusted	0,536	0,541	0,561	0,211	0,214	0,250

Column (4.1) to (4.3) report results for FI fixed effects regressions. Column (4.4) to (4.6) report results for country fixed effects regressions. Standard errors are reported in parentheses below their coefficient estimates and are adjusted for both heteroskedasticity and within correlation clustered at the FI level. ***, **, * indicate 1%, 5%, and 10% significance, respectively

The results shown in Table 7 have the triple interaction between CDS, MSIC, and Crisis as its main variable of interest. In the specification with country fixed effects, in column 2, the triple interaction is positive and significant at the 10% level. In this line, with a fixed MSIC index, we find that, in periods of crisis, increases in the CDS are associated with lower risk taking, or, similarly, with a fixed CDS spread, increases in the MSIC are associated with lower bank risk taking during crises.

Table 7 - Bank Risk Taking and Triple Interaction

Variables	**Z-Score**	**Z-Score**
	(5.1)	**(5.2)**
CDS	-0,056***	-0,052***
	(0,004)	(0,004)
MSIC	-1,641***	-1,197***
	(0,133)	(0,123)
Crisis	-0,814***	-0,801***
	(0,101)	(0,091)
CDS*MSIC	0,121***	0,100***
	(0,021)	(0,020)
CDS*Crisis	0,024	-0,028
	(0,070)	(0,047)
MSIC*Crisis	0,564***	0,506***
	(0,163)	(0,150)
CDS*MSIC*Crisis	0,042	0,149*
	(0,110)	(0,079)
Constant	1,094***	0,850***
	(0,076)	(0,071)
Nº Clusters	5424	5424
Nº Fixed Effects	5424	32
Nº Observations	37.670	37.670
F	129,470	140,320
R^2	0,608	0,215
R^2 Adjusted	0,542	0,215

Column (5.1) reports results for FI fixed effects regressions. Column (5.2) reports results for country fixed effects regressions. Standard errors are reported in parentheses below their coefficient estimates and are adjusted for both heteroskedasticity and within correlation clustered at the FI level. ***, **, * indicate 1%, 5%, and 10% significance, respectively

However, when we use bank fixed effects, the coefficient of the triple interaction is positive, but not statistically significant. It is indeed possible that the bank fixed effects capture most of the cross sectional variation in MSIC, which makes the remaining time-variation too small to

be "spread" among three variables (MSIC, its interaction with *Crisis* and the triple interaction). In addition, most of our previous inferences are maintained.

4.3 Robustness Tests

We follow Beck, De Jonghe, and Schepens (2013) and separately analyze the effects of our independent variables in the 3 components of the z-score: ROA, Capital to Assets ratio, and the standard deviation of ROA. The results are shown in Table 8.

Table 8 - Bank Risk Taking and Z-Score Components

Variables	**ROA**		**Capital to Assets Ratio**		**Standard Deviation of ROA**	
	(6.1)	**(6.2)**	**(6.3)**	**(6.4)**	**(6.5)**	**(6.6)**
CDS	-0,095***	-0,079***	-0,012***	-0,011***	0,039***	0,040***
	(0,025)	(0,021)	(0,003)	(0,003)	(0,003)	(0,003)
Crisis	-0,243***	-0,313***	-0,141***	-0,226***	0,636***	0,629***
	(0,060)	(0,062)	(0,021)	(0,027)	(0,089)	(0,084)
CDS*Crisis	0,108***	0,098***	0,011	0,014*	-0,107***	-0,119***
	(0,018)	(0,018)	(0,007)	(0,007)	(0,017)	(0,016)
Assets	-0,224***	-0,070***	-0,283***	-0,098***	0,144**	-0,099**
	(0,040)	(0,007)	(0,022)	(0,005)	(0,058)	(0,010)
Prob. Sup.	-0,096	0,268***	-0,000	-0,104**	0,343***	0,699***
	(0,068)	(0,064)	(0,033)	(0,046)	(0,127)	(0,081)
Liquidity	-0,018	-0,020	-0,021***	-0,009	0,090***	0,156***
	(0,012)	(0,014)	(0,005)	(0,010)	(0,019)	(0,020)
Herfindahl Ind.	1,215***	1,161***	0,093	0,304	-0,617**	-0,156
	(0,234)	(0,264)	(0,104)	(0,101)	(0,263)	(0,268)
GDP per capita	-0,313***	-0,451***	0,776***	0,551***	-0,118	0,165
	(0,098)	(0,073)	(0,037)	(0,028)	(0,128)	(0,111)
GDP Growth $_{t-1}$	0,001	0,004**	-0,009***	-0,007***	-0,026***	-0,030***
	(0,001)	(0,001)	(0,000)	(0,000)	(0,002)	(0,002)
GDP Growth $_{t-2}$	-0,001	-0,001	0,005***	0,004***	0,041***	0,043***
	(0,002)	(0,002)	(0,000)	(0,000)	(0,002)	(0,002)
Banking Credit (%GDP)	-0,004***	-0,006***	-0,004***	-0,005***	-0,002*	-0,004*
	(0,001)	(0,001)	(0,000)	(0,000)	(0,001)	(0,001)
MSIC	-0,277***	-0,420***	-0,149***	-0,181***	2,104***	1,706***
	(0,074)	(0,074)	(0,034)	(0,039)	(0,143)	(0,132)
MSIC*Crisis	0,183*	0,345*	0,096***	0,255***	-0,537***	-0,531***
	(0,102)	(0,106)	(0,036)	(0,047)	(0,141)	(0,134)
Constant	6,041***	5,505***	-1,612***	-1,888***	-4,697***	-4,095***
	(0,975)	(0,846)	(0,322)	(0,340)	(1,276)	(1,195)
Nº Clusters	4673	4673	4758	4758	4758	4758
Nº Fixed Effects	4673	31	4758	31	4758	31
Nº Observations	29.200	29.200	32.061	32.061	32.061	32.061
F	46,190	40,370	136,270	136,270	90,460	109,350
R^2	0,730	0,264	0,921	0,290	0,674	0,315
R^2 Adjusted	0,679	0,263	0,908	0,289	0,617	0,314

Column (6.1) , (6.3) and (6.5) report results for FI fixed effects regressions. Column (6.2) , (6.4) and (6.6) report results for country fixed effects regressions. Standard errors are reported in parentheses below their coefficient estimates and are adjusted for both heteroskedasticity and within correlation clustered at the FI level. ***, **, * indicate 1%, 5%, and 10% significance, respectively

The coefficients of the variables MSIC, CDS, and the interaction of both with the crisis dummy have an expected sign and a significant relationship with the 3 regressands. More specifically, in column 1 of Table 8, the results show that the ROA is 24.3 percentage points (pp) smaller during the crisis (results statistically significant at 1%) for a bank located in a country with a CDS spread and MSIC equal to zero. An increase of 100 bps in the CDS spread is associated to an average decrease of 9.5 pp in the ROA in normal times. In turn, this relationship is considerably changed in times of crisis. That is, every increase of 100 bp in the CDS represents an increase of 1.30 pp (100 * (-0.095 + 0.108)) in the ROA.

Concerning the MSIC, there is an alignment with Gropp, Hakenes, and Schnabel (2011), which suggests that increases in this variable are related to decreases in profitability. For each increment of 0.1 in the MSIC, the ROA is reduced by 2.77 pp (-0.277 * 0.1). However, the MSIC negative effect on profitability is substantially mitigated during the crisis. In this case, for increases of 0.1 in the MSIC, a reduction of only 0.94 pp is observed ((-0.183 + 0.277) * 0.1) in the profitability indicator.

The negative and significant coefficient of the Assets variable points to a decrease in profitability, whereas the total assets increase. Thus, an increase of 1% in the balance of assets of the FI corresponds to a decrease of 0.224 pp in the ROA. By reexamining the coefficient of the Herfindahl variable, the charter value hypothesis is confirmed. That is, it is found that increases of 10 bp in the concentration index result in increases of 12.15 pp in the ROA.

The coefficient of the GDP per capita variable, which is negative and significant, provides evidence that the average profitability of FIs operating in developed countries is lower than that of FIs operating in developing countries. Along these lines, it is found that, for each 1% increase in the GDP per capita, the ROA is reduced by 0.313 pp.

In column 3 of Table 8, we run an analogous regression using the Capital to Assets Ratio as the dependent variable. In line with what was indicated above, increases in the CDS spread rate are related to decreases in the dependent variable, i.e., greater risk taking. However, there is an inversion in the sign of the coefficient in periods of crisis. Thus, while in normal times, an increase of 100 bps in the CDS corresponds to a 1.2 pp decrease (100 * -0.012 * 1) in the index, during crises, the same variation is linked to an increase of 0.8 pp (100 * (-0.012 + 0.020) * 1) in the ratio.

Regarding the impact of the MSIC on the Capital to Assets ratio, the hypothesis that, in normal times, increases in this variable are related to greater risk taking, is again corroborated. Thus, for each increase of 10 bps in the MSIC, the Capital to Assets ratio is reduced by 1.49 pp (100 * -0.149 * 0.1). However, during the crisis, the negative effect of the MSIC on profitability is substantially mitigated: a 10 bp increase in the MSIC is associated to a reduction in the indicator of only 0.53 pp (100 * (-0.149 + 0.096) * 0.1).

Finally, in Table 9 it is confirmed that FIs located in countries with high CDS spreads are less risky than FIs located in countries with low CDS spreads in normal times. During the crisis however, banks in high-CDS countries increase their risk more than banks in low CDS countries. This result suggests that, during the crisis, FIs acting in countries with smaller bailout capacity (high CDS) increase their risk taking less than their counterparties in the low credit risk countries, possibly because they are less likely to be bailed out.

Table 9 - Difference - in - Differences Model

Variables	Z-score	ROA	Capital to Assets Ratio	Standard Deviation of ROA
LCDS	0,840***	-0,690***	-0,310***	-1,155***
	(0,021)	(0,014)	(0,008)	(0,022)
Crisis	-0,256***	-0,199***	0,033**	0,281***
	(0,035)	(0,024)	(0,014)	(0,036)
LCDS * Crisis	-0,098**	0,212***	0,012	0,115***
	(0,040)	(0,027)	(0,016)	(0,042)
Constant	-0,519***	-0,534***	2,226***	-1,814***
	(0,018)	(0,012)	(0,007)	(0,019)
Nº observations	41.632	38.002	41632	41632
F	782,320	878,630	596,460	128,590
R^2	0,050	0,060	0,041	0,085

Standard errors are reported in parentheses below their coefficient estimates are adjusted for both heteroskedasticity and within correlation clustered at the FI level. ***, **, * indicate 1%, 5%, and 10% significance, respectively.

5 Conclusion

The goal of this article is to assess the impact of government rescue expectation on bank risk in periods of crisis and normality. Additionally, it is tested whether the countries' financial capacity, measured by the Credit Default Swap (CDS) spread, can contribute to changing this effect.

In line with the moral hazard hypothesis, it is found that increases in the bailout expectation and the total assets of an institution are associated with greater risk taking. In addition, the market share of insured competitors (MSIC) or the rescue expectation of the competitors of an institution also influences its risk taking. However, depending on the period of analysis, this link can be different; that is, in normal periods, there is a predominance of the channel through which increases in the MSIC distort competition, reduce profit margins, and increase the risk taking of small institutions. During crises, however, increases in the MSIC indicate lower risk taking. It is assumed that the reason is, the higher the MSIC of an institution, the lower its relative importance in the system and thus the lower its prospect of eventually being rescued. Therefore, there is greater conservatism.

Similarly, we observe that the association between sovereign CDS spread and banking risk is modified during the crisis. Thus, while in normal times, increases in the sovereign CDS are linked to higher levels of banking risk, in the crisis, this effect is reversed. That is, in line with the moral hazard hypothesis, by raising the CDS and thus reducing the financial ability of a country to undertake bank bailouts during the crisis, there is a decrease in the appetite for risk. Finally, corroborating previous results, indications are found that, during the crisis, institutions located in countries with high CDS spreads have, on average, an increase in the risk level that is lower than that in countries with low CDS spreads.

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