

Short sales constraints and financial stability: evidence from the Spanish 2011 ban

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Abstract

This paper studies the impact of the short sales ban implemented in August 2011 in the Spanish stock market along two dimensions: financial stability and market performance. Regarding the former, we show that before the ban short positions were a significant determinant of the probability of default of medium-sized banks, typically more reliant on the domestic market and on public support than their larger competitors. We find that, by weakening the contagion effect coming from the sovereign risk, the ban helped stabilise the probability of default of medium-sized banks, an effect which is not significant in the case of the largest banks and non-financials. Nonetheless, the stabilising power of the ban came at the cost of a large decline in the relative liquidity, trading volumes and price information efficiency of medium-sized banks' stocks.

Keywords: Short-sales constraints; financial stability; financial institutions; credit default swap; contagion.

JEL Codes: G01, G12, G14, G18.

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1. Introduction

In August 2011, the Spanish Securities Market Commission (CNMV) together with three other European supervisors announced a precautionary ban on establishing or increasing short positions in financial institutions' stocks. This exceptional measure was adopted at a time of heightened aggregate uncertainty and volatility in financial markets, coinciding with the resurgence of tensions in the sovereign debt markets of various euro area countries.

Although that volatility spike had several roots, the direct cause seems closely linked to the problems of fiscal management on both sides of the Atlantic. On the one hand, delays in securing political agreement to raise the U.S. federal deficit ceiling fuelled the climate of uncertainty in financial markets worldwide. In the European markets, the spark was the failure of the European Council of 21 July 2011 to arrive at a clearly articulated and adequately funded agreement to guarantee the viability of Greece's public finances. As in earlier phases of the European sovereign debt market crisis, the expectation of a restructuring of Greece's public debt had a knock-on effect on other economies in the region. This time, however, the fallout was greater and more widely distributed than in earlier episodes. In particular, from the end of July 2011, a number of euro area countries like Italy, Spain, Belgium and France, which had safely ridden out the previous crisis episodes, began to suffer growing pressure on their sovereign bonds that manifested through the widening of their sovereign CDS spreads to then record highs. The fact that several of the region's largest economies were pulled into the debt storm heightened the perception of systemic risk for the whole area.

As it has been apparent since the initial stages of the current global financial crisis, the close linkages between the sovereign and the financial sector facilitated at that time the rapid spread of public debt market tensions to some key wholesale markets. In fact, the

intensification of the Greek crisis, through its spillovers into other euro-area sovereign debt markets, soon pushed up the credit risk indicators of most European banks.

In this context, on 11 August 2011, the European Securities and Markets Authority (ESMA) published a communication calling for harmonised regulatory action on short selling in the EU.¹ On this basis, ESMA justified its support for the national securities authorities of Belgium, Spain, Italy and France, all of whom announced that same day the imminent implementation of bans on the short selling of financial stocks. The Spanish regulator, in its own communication enacting the ban, characterised the extreme volatility then affecting European securities markets in general, and bank sector stocks in particular, as posing a material threat to the stability and the orderly functioning of markets. Therefore, the upkeep of financial stability was the leading motive for its decision to prohibit temporarily short positions in the shares of 16 Spanish financial corporations.

In this paper, we analyse some of the main implications of this ban along two dimensions: financial stability, which we proxy by means of some indicators of banks' default-risk, and market performance, including the impact of the ban on liquidity, stock prices and price discovery. In order to approximate the ban's power to support financial stability, we first offer empirical evidence that shows a positive causal relation between short positions and the credit risk indicators of medium-sized Spanish banks in the weeks preceding the ban, which subsequently eased. We then find that restricting short sales on these stocks had some stabilising influence on the market perception of the banks' solvency, a result that is found to be insignificant for the country's two largest banks. Indeed, following

¹ The persistent rumours sweeping the markets in the days before ESMA's communication about the supposedly perilous state of various European banks had particularly targeted some French institutions (see, for instance, the article "Panic in Paris", in *The Economist*, 20 August 2011). On 10th August, Société Générale lost as much as 20% of its stock exchange value at some points in trading before closing 15% down, while BNP Paribas and Crédit Agricole posted losses of 9.5% and 12%, respectively. The Banque de France had to issue a communication the following day underscoring the financial soundness of these institutions (available at: http://www.banque-france.fr/uploads/tx_bdfgrandesdates/press-release-Noyer-11-08-2011.pdf).

the inception of the ban, we document a decoupling between the default risk of the sovereign debt and that of the debt issued by medium-sized banks. Such a significant decoupling, that came hand in hand with a sizeable moderation in the path of the probability of default indicators of this last group of banks, is not found in either the largest banks or the non-financials. The asymmetric impact of short positions on the default risk indicators of medium-sized banks relative to the two largest banks of the country could well reflect the higher reliance of the former on the domestic market and on the implicit and explicit support of the Spanish government. In fact, the two largest banks are internationally diversified and perceived as much better capitalized than most of the medium-sized banks and, hence, less dependent on implicit or explicit public guarantees.

The stabilising effect on medium-sized banks is reflected not only in their credit risk indicators but also in the volatility of their stock returns, which fell significantly both in absolute terms and in comparison with that of the large banks and the non-financial companies. We also document an interesting collateral effect of the ban on short sales in financial stocks: the short positions in non-financial companies rose significantly following the ban. This reflected, in part, the existence of a common market-wide risk factor that is present in the sovereign CDS.

Nonetheless, the previous stabilisation effects of the ban on the credit risk indicators of medium-sized banks came at the cost of a significant and lasting decline in the liquidity of the medium-sized banks' stocks, which also suffered a more severe decline in their exchange trading volumes and a slowdown in the process of price formation than the other two groups of firms. This suggests that the possible benefits of a restriction on short-selling, in terms of its potential for stabilising the financial conditions of certain firms, must be carefully weighed against some significant losses in the quality of market functioning.

The paper is organised as follows. Section 2 discusses the related literature. Section 3 describes the data set. Section 4 analyses the effects of short sales on the probability of default of companies listed in the Spanish market before the ban, and studies its effects on corporate credit risk. Section 5 examines the effect of the ban on the Spanish stock market in terms of liquidity, volatility, excess returns and price discovery. Section 6 contains some final remarks.

2. Related literature

Our paper is directly connected to two different strands of the literature on short-selling constraints. The first is focused on the role of these constraints as devices to maintain financial stability. The second is mainly concerned with the impact of short sales constraints on several dimensions of stock market performance. In what follows we consider how our analysis is shaped by and contributes to some earlier analyses within these two areas.

Short sales constraints and financial stability

In a context dominated by aggregate economic and financial uncertainty, an increase in short positions in financial stocks can be understood as a natural market response reflecting heightened investor pessimism. There are, in fact, several channels through which an upturn in aggregate risk and market volatility can dull the outlook for present and future financial sector earnings. For example, a decline in the value of government bond portfolios will typically damage banks' balance sheets given their hefty exposure to this type of asset. Also, the fact that in the current financial crisis many governments are lending direct or indirect support to their domestic financial sector through guarantees, liquidity lines, etc., may cause the transmission of sovereign debt market tensions to banks' stocks. Likewise, a worsening of the macroeconomic outlook will tend to depress expectations for banks' net interest income and loan book quality, again bearing down on stock market quotes.

At the same time, it is conceivable that, in some circumstances, an increase in short positions in financial stocks could magnify the effect of any perturbations hitting the sector, in which case curtailing this kind of trade should, in theory, support the stability of affected institutions. The potential linkage between short selling of financial stocks and banking stability has been recently explored from a theoretical standpoint in several articles.

Brunnermeier and Pedersen (2005) analyse how short-selling may exacerbate fluctuations in assets prices, in the context of a model of fire sales under liquidity frictions. In particular, these authors explore the case in which short-sellers anticipate the need of one trader to sell an asset and, strategically, add some selling pressure on that asset upon the expectation of buying it back later at a lower price. Key to the logic of the underlying mechanism that renders such strategies optimal is the idea that short sales may trigger an undershooting effect on assets prices by allowing opportunistic investors to mimic the moves of other traders under pressure to liquidate their investments. A direct testable implication, which in fact we test with our data, is that a short-selling constraint, given everything else, should contribute to reduce the volatility of the price of the stocks affected by it.

Liu (2010) examines a mechanism through which short sales of a bank's stock can cause its failure. At the core of this mechanism is the idea that risk-averse creditors, who extract information from the shares price about the firm's underlying fundamentals, become increasingly unsure about the true fundamentals as the shares prices turn more volatile. Thus, they become less willing to maintain their exposure to the bank in question. As this happens at a sufficiently large scale, "too noisy" stock prices may end up triggering a bankruptcy. In anticipation of this, speculators find it then optimal to short-sell the stock beforehand to

amplify price volatility. As a result, they crowd out lenders and, therefore, increase the likelihood of benefitting from the firm's collapse.²

Venter (2011) also discusses the link between short sales and bank solvency in a context of equilibrium multiplicity in which different agents access asymmetric information sets. Specifically, he shows that under some conditions short-selling constraints may help avoid a run on a bank by providing the right incentives for the participation of the most active investors in the market, even though this entails some informational efficiency costs.

Previous arguments underscore the idea that under some circumstances restricting short sales on banks' stocks may improve their financing conditions, by mitigating the incidence of imperfect knowledge and asymmetric information between the various classes of investors in a firm. We test this argument by examining the effects of short positions on some indicators of the default likelihood at the firm level in the months preceding the inception of the short-selling ban by the Spanish supervisor in August 2011, exploiting a wide sample of financial and non-financial companies. In particular, we take the CDS spreads and some model-based estimated probabilities of default as indicators of the firms' degree of financial vulnerability. A key piece of information used in this analysis is the series of short positions exceeding 0.2% of the total capital of every firm in the Spanish market. Then, based on event-study techniques, we assess the effect of the short-selling ban on the previous indicators of financial vulnerability, distinguishing by type of firm: medium-sized banks (almost exclusively oriented to the domestic market and more reliant on public support), large

² Standard & Poor's (2008) contains an interesting explanation of how this rating agency takes into account stock prices as signals in its credit risk assessments depending on the nature of the firm and the specific circumstances, arguing that under some conditions, especially in the case of *confidence-sensitive* companies, falls in stock prices may affect significantly the creditworthiness of a firm.

banks (well diversified internationally and less dependent on government aids) and non-financial firms (not subject to the short sales ban).³

The effects of short sales constraints on market performance

The effects of short sales on some basic dimensions of the functioning of the stock market, including liquidity, informational efficiency and the possibility of overvaluation, have received far more attention in the existing literature than the links between short-selling and financial stability described earlier. Diamond and Verrecchia (1987) developed a model with heterogeneous investors, in which, given the costs on short sales, some cannot take on short positions. The setup is then exploited to analyse how short-selling constraints affect the speed of adjustment of stock prices and also their level. A central conclusion of the analysis is that a prohibition on short sales, by crowding out some traders, reduces the speed of adjustment of prices to private information and, as a result, average bid-ask spreads rise.

The previous conclusion regarding the impact of short sales on liquidity and informational efficiency has received ample factual support in a number of recent empirical studies. For instance, as regards the effect of short-selling constraints on liquidity, Boehmer, Jones and Zhang (2011) – using stocks traded in U.S. markets – and Marsh and Payne (2012) – using references drawn for the U.K. market – find that restrictions on short positions damaged liquidity significantly, as proxied by bid-ask spreads. Beber and Pagano (2013) document a similar detrimental effect of shorting constraints on liquidity, exploiting a large panel of data from 30 countries over the period January 2008 to June 2009, when many of the countries enacted and lifted bans on short positions in the context of the global financial crisis. As regards price discovery, Bris, Goetzmann and Zhu (2007), Saffi and Sigurdsson

³ The group of medium-sized banks includes Banco Pastor, Banco Popular, Banco Sabadell, Banesto, Bankia, Bankinter and Caixabank. The two big banks are BBVA and Santander,

(2011), Boehmer and Wu (2012) and Pagano and Beber (2013) all find evidence suggesting that short sales increase the degree of informational efficiency of securities prices.

The impact of shorting bans on the shares price level is more controversial. At the theoretical level, models of disagreement, à la Miller (1977), typically prescribe that by limiting the participation of the most pessimistic traders, short sales constraints tend to bias prices upwards. By contrast, Diamond and Verrecchia (1987) argue that if market makers are risk-neutral, then prices will not be affected by the absence of short-sellers. This is because market makers do internalise the information that, in expectation, would be otherwise left out by the prohibition. This divergence of conclusions extends to the empiric front too. For instance, Boehmer, Jones and Zhang (2011) find little evidence that the 2008 ban on short sales in the U.S. caused a positive effect on prices. This finding, they argue, could reflect that the ban was expected to be short-lived since regulation in the US only allows the supervisor (the Securities and Exchange Commission, SEC) to maintain the ban for up to 30 days. Beber and Pagano (2013) also fail to find evidence in favour of the price-support hypothesis for most countries in their sample. On the other hand, Chang, Cheng and Yu (2007) find evidence that shorting constraints do convey an overvaluation effect in the Hong Kong market, which is stronger in those stocks with wider dispersion of investors' beliefs.

A similar lack of coincidence seems to exist around the conceivable effect of shorting constraints on the volatility of stock prices. For instance, while Chang, Cheng and Yu (2007) find that stocks subject to the ban exhibit lower volatility, Ho (1996) notes that shorting restrictions in Singapore in 1985 came with more volatile stocks returns. Boehmer, Jones and Zhang (2011) report a significant increase in returns volatility following inception of the ban in the U.S. in 2008, and fail to identify any significant difference between the volatility of the stocks subject to the ban and those unaffected by it within the ban period.

Besides methodological differences with respect to some of the previous papers, the main result arising from our analysis is related to the marked divergence between the market performance indicators for medium-sized banks vis-à-vis large banks and non-financial firms following the introduction of the shorting ban. While some of these papers have identified differences along previous dimensions of market performance across different groups of firms, based on certain indicators of size or trading activity (see e.g. Bohemer et al. 2011), interestingly, we find that such differences extend to every dimension analysed here (liquidity, returns volatility and price discovery), except relative stocks returns. In the latter case, we find no significant differences between the three groups considered following the ban on short sales of financial shares.

Taken together, the results obtained in this paper, concerning both the effect of the ban on the indicators of firms' solvency and on market performance, offer some interesting new insights when considered alongside previous literature. First, based on our results, we would argue that the "success" of a ban on short sales aimed at supporting financial stability should be better assessed against its power to stabilise the financial resilience of the firms targeted by the ban, rather than merely on the basis of its effect on relative stock prices. Second, the recent experience in the Spanish stock market reveals in a neat way the existence of a sort of trade-off between the effectiveness and the efficiency costs of a short-selling constraint, in the sense that the higher power of the ban to support the financial strength of a subset of banks unleashed important damages in the liquidity, volume of trading and price discovery of their stocks.

3. Data

We exploit daily information on the stock of short positions and stocks lending of shares trading on the Spanish continuous market (SIBE) from March 2011 to December 2011. These series come from a proprietary CNMV data set. Regarding the stock of short positions,

investors are obliged to notify the authorities of any net short position exceeding 0.2% of the issued share capital of the company concerned and each 0.1% above that.⁴ Further, the threshold for public disclosure of net short positions is set at 0.5% and each 0.1% above that, thus the information is available to all market participants.

The data on CDS spreads of the sovereign CDS and the Spanish corporations are obtained from Credit Market Analysis (CMA). This information is available for only 17 of the Spanish firms in the continuous stock market.⁵ The data are employed as indicators of the corporate risk premiums. CDS spreads are also used to compute the spillovers in credit risk between the sovereign and the financial sectors in Section 4. Due to the lack of CDS for most of the firms forming the Spanish market, we use alternatively default probabilities and implied ratings that are obtained from the StarMine Structural Credit Risk model through Reuters platform. Based on the input provided by equity market participants, this model produces an estimate of the probability that a company will go bankrupt or default on its debt obligations over the next one-year period.

The information on daily stocks prices, including the bid and ask prices, and the trade volume are obtained from Datastream. The information on the index Ibex 35, the EuroStoxx 50 and the volatility indexes VIX and VSTOXX also come from Datastream. Finally, the information regarding the European Central Bank (ECB) bond purchases was obtained from the ECB webpage.

⁴ Uncovered short positions, in which the seller has not ensured the availability of the securities, are prohibited in Spain.

⁵ The 17 firms include all banks in the sample with the exception of Banesto for which bid and ask CDS prices are not available and some non-financials (Abertis, ArcelorMittal, Endesa, Gas Natural, IAG, Iberdrola, Melia Hoteles, Repsol, Telefonica).

4. The effects of the ban on the credit risk of financial institutions

Towards the end of July 2011, many European banks suffered a sharp run-down in their shares prices and a parallel rise in their default-risk indicators (see left panel of figure 1). The right panel in figure 1 shows that the surge in financial sector risk coincided with a jump in the indicators of the risk-contagion effect running from the sovereign to the financial sector in Italy and Spain, precisely at a time when the two countries' sovereign risk indicators were deteriorating at a very fast rate.⁶ This last effect added to the deterioration of quality of the non-sovereign assets held by many European banks, putting further pressure on their capital ratios, hampering their access to financing and, thus, increasing their financial vulnerability.

< Insert Figure 1 here >

In the remainder of this section, we first analyse the impact of accumulated short positions on the several credit risk indicators of the companies listed in the Spanish market (section 4.1) and then provide some estimates of the effect of the ban on those indicators and also on the strength of the sovereign-corporate risk-contagion effect (section 4.2). The idea in the last case is to assess the extent to which the ban on short sales was effective in isolating the financial companies from the dominant aggregate source of risk at that time, i.e. sovereign risk.

4.1 Short positions and credit risk

We next construct an empirical model aimed at identifying the effect of short positions (relative to the share capital) on the log-return of CDS for the three groups of firms considered (medium-sized banks, large banks and non-financial companies) before the ban. The model is estimated using data from March 2011 to December 2011 on the basis of the

⁶ The details about the estimation of the contagion (spillover) effects can be found in the Appendix.

following fixed-effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation:

$$\begin{aligned}
\Delta \log(CDS_{i,t}) = & \alpha + \beta_1 Short_{i,t-1} + \beta_2 Short_{i,t-1} x Dum.PreBan_t x Dum.LargeBank_i \\
& + \beta_3 Short_{i,t-1} x Dum.PreBan_t x Dum.MedBank_i \\
& + \beta_4 Short_{i,t-1} x Dum.PreBan_t x Dum.NonFin_i \\
& + \beta_5 Dum.PreBan_t + \beta_6 Dum.MedBank_i + \beta_7 Dum.LargeBank_i \\
& + \zeta Controls_{i,t} + \varepsilon_{i,t} \quad (1)
\end{aligned}$$

where $\Delta \log(CDS_{i,t})$ denotes the log return of the CDS spread of firm i at date t and $Short_{i,t-1}$ denotes the ratio of short positions relative to the share capital of firm i at date $t-1$.⁷⁸ To study the incremental effect of the short positions before the ban for each type of firm, we include an interaction variable which is obtained as the product of the short positions in each firm lagged one period and two dummy variables: one dummy for each type of firm ($Dum.LargeBank$, $Dum.MedBank$, and $Dum.NonFin$) and another that takes the value of one before the ban (March 2011–August 2011) and zero otherwise ($Dum.PreBan$). We also use the individual dummies referred to the pre-ban period and the types of institutions with no interaction terms.

The regression also includes controls as possible additional explanatory factors for the percentage changes in the CDS spreads: the percentage changes in the CDS liquidity (bid-ask spread), the lagged stock returns, a measure for the stock return volatility lagged one day,

⁷ Excluded from the sample are short positions in the shares of non-banking financial corporations affected by the ban, mainly, insurance companies and providers of financial services other than banking ones (Grupo Catalana Occidente, Mapfre, Bolsas y Mercados Españoles, and Renta 4 Servicios de Inversión) and Caja de Ahorros del Mediterráneo.

⁸ We use the information referred to the short-positions related variables in levels because we are interested in both the trend and the level of the variables. We perform three panel unit root tests – Levin, Lin, Chu’s (2002) test, Modified Augmented Dickey-Fuller test (Taylor and Sarno, 1998) and Fisher-ADF type unit root test (Maddala and Wu, 1999) - and find no evidence of unit root in the panel data. Note also that the short-positions variables are defined as a ratio and so, their values are bounded between zero and one.

lagged percentage changes in the VIX index, percentage changes in the Spanish sovereign CDS lagged one day and percentage changes in the level of CDS spreads lagged one period. All these variables allow us to control for a number of potential explanatory factors of the credit risk of each group.

We also take into account that some days before the ban, and with effect from 8th August 2011 onwards, the European Central Bank (ECB) issued a communication welcoming the new measures and reforms announced by the Italian and Spanish governments. The ECB also restated its availability to implement actively its Securities Markets Programme (SMP). Under this programme, designed to support the orderly functioning of key financial markets, the ECB was authorised to make purchases of euro-area sovereign bonds in the secondary market. This declaration may have had a calming effect on the sovereign risk and, indirectly, on the private sector risk too. For this reason, as an additional control we consider the ECB's purchases of government bonds, in billions of euros, in the secondary market from 8th August onwards. The results are reported in column 1 of table 1.

From these results we can single out that before the ban the volume of short positions in medium-sized banks acted as a significant determinant of changes in their average CDS, such that the greater the volume of short positions, the higher the level of this credit risk indicator. In particular, a change of one standard deviation in the ratio of short-positions of medium-sized banks relative to their share capital during the pre-ban period (1.16%) would cause an average increase of 0.26% in their CDS premiums, which represents 37% of the average daily variation of the CDS premiums. By way of contrast, the pre-ban volume of short positions over the capital of the other two groups of firms does not seem to cause significant changes in their CDS.

In the case of the control variables considered, the direction of the effect is as expected. In particular, percentage changes in the sovereign CDS spreads have a significant

positive effect on average changes in the CDS of listed companies. As expected, increases in the CDS bid-ask spread, which signal a lower liquidity of the CDS contracts, have a positive and significant effect on the dependent variable. Conversely, increases in the average firm's stock returns of the previous trading session have a significant negative effect on the firm's CDS. This last result points towards the relatively high information content of share prices regarding company financial soundness in a sample period as convulsive as the one in question. We also note the negative and significant effect from the lagged changes in the CDS spreads, whose scale, low in absolute terms, reveals the persistent nature of this variable's deviations from its long-term average. The effect of the two volatility variables, the global indicator of market volatility (VIX Index) and the firm specific equity volatility (squared of stock returns), is not significant at 5%.

Finally, the ECB's purchases of government bonds in the secondary market have a negative effect on CDS, as expected, but this effect is not significant at any standard significance level. This lack of significance of the ECB purchases of debt after 8th August 2011 could be due to the seemingly short-lived impact of this announcement on the Spanish sovereign debt risk, arguably the variable that should have reacted more directly to that announcement. Indeed, as shown in figure 2 the Spanish sovereign risk surged again a few days after a sharp fall following the announcement. More formally, we regress the first difference of the sovereign CDS spread on a dummy variable that takes the value of one during the one and a half month after 8th August and zero during the same period before that date, and we find a non-significant effect of the dummy coefficient at any standard level of significance. Hence, we interpret the absence of a significant effect of the ECB's announcement on the Spanish sovereign CDS as the most plausible explanation for its low impact on the indicators of credit risk of the Spanish listed firms found in the estimates of table 1.

< Insert Figure 2 here >

One may argue that short-sellers are able to anticipate an increase in credit risk as opposed to causing the increase in such a risk. Henry, Kisgen, and Wu (2010) analyze whether short-sellers anticipate bond rating downgrades and show that daily data do not indicate significant increases in short selling prior to (or following) a downgrade. Based on this possibility, we re-estimate equation (1) using our daily short-positions variables, referred to the total sample and the three types of institutions before the inception of the ban, lagged 5, 10, and 15 days, and find that the short positions in medium-sized banks still have a significant effect on their levels of credit risk.

We next estimate a variation of the model in equation (1) in which we introduce new interaction variables that are not interacted with the short positions:⁹

$$\begin{aligned} \Delta \log(CDS_{i,t}) = & \alpha + \beta_1 Short_{i,t-1} + \beta_2 Dum.PreBan_t x Dum.LargeBank_i \\ & + \beta_3 Dum.PreBan_t x Dum.MedBank_i \\ & + \beta_4 Dum.PreBan_t x Dum.NonFin_i + \beta_5 Dum.MedBank_i \\ & + \beta_6 Dum.LargeBank_i + \zeta Controls_{i,t} + \varepsilon_{i,t} \quad (2) \end{aligned}$$

This specification enables us to test whether the rise in the levels of the CDS spreads before the ban was mainly due to any other distinctive group feature such that the surge in the CDS of medium-sized banks could have taken place independently of the short positions taken on these banks. Thus, we check whether the dummy for medium-sized banks is significantly different from zero before the ban. If that were the case, then short positions may not contain valuable additional information as determinants of the firms' credit risk. The second column of table 1 reports the incremental effects on the CDS spreads for the three groups of firms

⁹ The pre-ban dummy employed in equation (1) is not employed now given that the joint use of the constant term would lead to problems of collinearity.

before the ban. These results show that this interaction variable is not significant, which suggests that the stock of short positions does add informational value to explain the credit risk of medium-sized institutions. The remaining controls have a similar magnitude and degree of significance as the ones reported in column 1.

We next test the robustness of the previous results by estimating equation (1) using the volume of stocks lending instead of that of short positions reported to the supervisor, as a proxy of short sales activity. The third column of table 1 reports the effect of the volume of stocks lending relative to total capital and its incremental effect before the ban.¹⁰ As in the baseline, we find that before the ban the adjusted volume of stocks lending for the case of medium-sized banks was a significant driver of the changes in their CDS, while this was not the case for the other two groups of firms. The R-square and the magnitude and significance of the controls are similar to those obtained in the baseline estimation reported in the first column of table 1.

< Insert Table 1 here >

We next estimate the effect of the short positions restricting the sample to the pre-ban period (March 2011 – August 2011) on the basis of a fixed-effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation:

$$\begin{aligned} \Delta \log(CDS_{i,t}) = & \alpha + \beta_1 Short_{i,t-1} x Dum. LargeBank_i + \beta_2 Short_{i,t-1} x Dum. MedBank_i \\ & + \beta_3 Short_{i,t-1} Dum. NonFin_i + \beta_4 Dum. MedBank_i + \beta_5 Dum. LargeBank_i \\ & + \zeta Controls_{i,t} + \varepsilon_{i,t} \quad (3) \end{aligned}$$

¹⁰ The volume of lent shares is typically influenced by dividend payments, increasing significantly before the date of the dividend and converging later on to the previous level. To control for this, we filter the variable by regressing it on a dummy that takes the value of one ten days before and after the dividend payment date and on a time dummy. We then use the residuals and the constant of this filtering regression to construct the explanatory variables referred to the securities lending activity that are used in the regression whose results are reported in column 3 of table 1.

The estimated effects of short positions are contained in column 4 of table 1. These estimates confirm that the volume of short positions in medium-sized banks was affecting significantly their average CDS before the ban. This last effect is not significant for the non-financial corporations and the largest banks in any of the four months preceding the ban. Again, we obtain a significant effect of the sovereign CDS premium.

The availability of CDS data for the firms in the Spanish stock market limits the previous analysis to 17 single-names, although these firms represent most of the trading volume in the Spanish stock market. Specifically, on average, they represent almost 80% of the total daily volume of trade over the sample period. To enlarge the number of firms in our analysis, we next exploit data on default probabilities and implied ratings of the listed companies, obtained from StarMine, instead of CDS spreads. This enables us to extend our analysis to 105 firms for which such information exists (out of 135 firms in the market), that together account for 93% of the average total daily trading volume.¹¹

The default probabilities (dp) series are logit transformed as in Altman and Rijken (2004), according to the formula $\ln(dp/(1-dp))$. We use the first difference of the transformed default probabilities as the dependent variable to be consistent with the previous baseline regression analysis. The model and the estimation methodology are similar to the ones summarized in equation (1) with the standard errors also clustered by firm to produce unbiased errors. Given that the dependent variable is the change in the logit transformed default probability, we do not use the percentage change of the CDS premium lagged one period as a regressor but the lagged change in the logit transformed default probabilities. Moreover, we do not use the information referring to the market value of equity and volatility in our regression because the default probabilities and ratings reported by StarMine are based

¹¹ Based on market sources, the StarMine Structural Credit Risk model evaluates the probability that a company will default on its debt obligations over the next one-year period.

on a model that is an extension of the structural default prediction framework introduced by Merton. Thus, this model uses the equity market value, its volatility and liability structure as inputs. As the ban affected not only big and medium-sized banks, but also other financials (insurance and financial services firms) that are now included in the sample, we also control for the short positions (lagged one period) on them before the ban.

The results, reported in column 1 of table 2, confirm that before the ban, short positions in medium-sized banks had a positive and significant effect on the changes of default probabilities. A change of one standard deviation in the variable referred ratio of short-positions of medium-sized banks relative to their share capital during the pre-ban period (0.49%) would cause an average increase of 0.13% in the their logit transformed default probabilities (around 0.53% in the non-transformed default probability) for this type of firms during that period. Additionally, we find a positive and significant effect of short positions in the other financial firms on their default probabilities. This result reinforces the one obtained for the medium-sized banks as the last group of institutions consists of relatively small firms, as compared with the big firms included in the Ibox 35. We do not find any significant effect for the other two groups of firms considered (big banks and non-financials). The negative and relatively low scale, in absolute terms, of the effect from the lagged changes in default probabilities reveals the persistent and significant nature of this variable's deviations from its long-term average. We find positive and significant effects of the changes in the VIX Index and the Spanish sovereign CDS, thus, suggesting that default probabilities were also being positively affected by global uncertainty and risk. As in the case in which we use the log returns of CDS as the dependent variable, the ECB's bond purchases in the secondary market do not have a significant effect at any standard significance level.

Besides the first difference in default probabilities, we use the first difference in the implied ratings reported by StarMine. In particular, we assign value of 1 to rating category

AAA and so successively until the last rating category in our sample, which is CC and takes a value of 21. The use of first differences means that an upgrade in the rating category will take value +1, while a downgrade will take value -1. As this implies that we are imposing a linear relation in the rating values, we repeat the analysis using two transformations of the rating category. First, we assign values that increase in one unit for the investment grade category and two units for the junk grade category. Thus, we assign values from 1 to 10 to ratings ranging from AAA to BBB- and from 12 to 32 for the ratings ranging from BB+ to CC. This implies that upgrades or downgrades in the investment grade (junk grade) category are assigned values of +1 or -1 (+2 or -2), respectively. Second, we assign value 1 for the ratings in the investment grade category and ratings ranging from 2 to 12 for the BB+ to CC in the junk grade category. That is, the dependent variable will take value 0 independently of the existence of downgrades or upgrades in the investment grade category and +1 or -1 depending on whether there is a downgrade or upgrade in the junk category. The results for these three variations of the dependent variable based on the implied rating are reported in columns 2 to 4 of table 2. Results are consistent with those reported in column 1. In particular, an increase of 1% in the level of short positions in medium-sized banks relative to their capital before the ban would raise the probability of a downgrade by 63.5%.

< Insert Table 2 here >

The results of this last exercise support and, in some cases, refine and supplement the main arguments and observations put forward earlier. Specifically, the evidence supports the presence of a singular effect on the perceived credit risk of medium-sized Spanish banks in response to the aggregate macro-financial turmoil unleashed in July 2011. As conjectured before, this result would be consistent with a the relatively higher exposure of medium-sized banks vis-à-vis the big banks to domestic macroeconomic risk (lower geographical

diversification) and Spanish sovereign risk (stronger perception of their eventual reliance on public aids).

4.2 The effect of the ban on credit risk

A first notable observation is that the average spreads of the CDS of the medium-sized banks eased appreciably following the ban, after several weeks of sustained increases (see figure 3). In effect, although their CDS continued trading higher than those of larger banks and non-financial listed companies, their trend clearly stabilised after the ban. Indeed, the CDS spreads of both sets of financial institutions headed lower in the first few weeks after the ban, while those of non-financial corporations went on accelerating for some weeks more.

< Insert Figure 3 here >

Indicators of spillover effects between the CDS of Spanish government bonds and those of the previous three groups of firms elucidate this incremental easing of medium-sized banks CDS in the wake of the ban. In particular, from figure 2 it can be seen that while the sovereign CDS headed higher with some interruptions until November, risk contagion from the sovereign to the medium-sized bank wore off appreciably once restrictions on short sales began. The indicator of contagion had been growing in intensity since approximately one month before, driving the CDS of medium-sized banks sharply higher in the intervening period. Conversely, the contagion from government bonds to the largest banks, which had likewise intensified a few weeks earlier, appeared largely unaffected by the ban and the same was true of the group of non-financial corporations.

In order to test statistically the hypothesis about the stabilising power of the ban, we next perform an estimation based on an event-study methodology of the behaviour of the CDS and the indicators of sovereign-corporate risk contagion for the three groups of firms before and after the ban. A first important question concerns the length of the sample period.

The net impact of short-selling bans may be clouded by the inclusion of observations that are far away from the inception date of the ban. Based on this reflection, we assess the impact of the ban by using information over a three-month period, centred on the date of its inception. Specifically, we regress the first difference of the indicator of sovereign-corporate risk contagion on three firm-group dummies that take the value of one during the post-ban period (one month and a half after the ban) and zero during the pre-ban period (one month and a half before the ban). We then run a similar regression replacing the contagion indicator by the first difference in the spreads of the corporate CDS as the dependent variable. The estimates are obtained using OLS fixed-effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation. Panel A of table 3 reports the coefficients of this estimation, while panel B contains the estimated differences between these coefficients, indicating whether the incremental effect of the ban for the different firm groups is significantly different from zero. The null hypothesis to test in panel B is whether the effects of the ban on the changes of the indicator of contagion and the CDS spreads for the first firm in the cells of the first column are stronger (more negative coefficients) than the effect for the second firm.

< Insert Table 3 here >

The main result of the first regression (column 1 of panel A) is that after the ban there was a significant (at a 5% significance level) decrease in the credit risk contagion indicator running from the sovereign to the group of medium-sized banks. This effect is also negative, but not significant, for the largest banks and positive, not significant either, for non-financials. The figures show that the average decrease of the daily variation in the sovereign risk contagion indicator corresponding to the medium-sized banks following the ban was around 1.5 percentage points. The comparison of the changes in the contagion indicator after the ban for the medium-sized banks with respect to the other two firm groups reveals that the

fall in the intensity of the contagion effect was significantly higher in the first group than in the other two (see column 1 of panel B). Thus, these estimates corroborate the intuition outlined before coming from the visual inspection of figure 2.

Regarding the effect of the ban on the CDS spreads, we find a negative impact for the three types of firms (see column 2 of panel A), which is, however, only significant (at a 5% confidence level) in the case of the medium-sized banks. The average decrease of the daily change in the CDS spreads of medium-sized banks after the inception of the ban was around 10 basis points. This decrease is found to be significantly higher, in absolute terms, than the ones observed for large banks and non-financial firms (see column 2 of panel B).

We finally analyse the effect of the ban on the short positions in non-financial firms, i.e. those not directly targeted by this measure. To this aim, we perform some additional exercises using information on short positions one quarter around the inception date. First, we regress the first difference in the short positions of non-financial firms on a dummy that takes the value of one during the post-ban period. The estimates are obtained using an OLS fixed-effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation. We find a positive and significant (at a 5% significance level) difference, or around 0.004 percentage points in the ratio of short positions over capital ratio, between short positions after and before the ban. Given that there are no short positions in many of the non-financial firms in the sample, we next restrict our sample to the firms for which there are short positions. In this subsample, the previous positive differential effect rises up to 0.01 percentage points.

The previous results suggest the existence of a shift-effect in short positions that were initially built up on the capital of financial entities to non-financial companies following the ban on the former. Such a substitution effect could be consistent with the idea that to some extent there was common underlying source of pessimism of short-sellers affecting both

financial and non-financial firms. According to this view, short positions in companies of either group would be seen somewhat as substitutes, as a bet against such a common aggregate risk factor. A natural candidate to proxy the latter is the sovereign risk. Hence, we next check whether there was a stronger increase in short-selling activity for certain types of non-financial firms according to their exposure to sovereign credit risk. With this goal in mind, we split the post-ban period dummy into two variables depending on whether the contagion indicator is above or below the median level of the non-financial firms during the quarter around the time of the ban. The corresponding regression reveals that the variation in short positions is larger (almost twice) in the firms that were more exposed to the sovereign risk. Actually, the dummy for the firms below the median level of contagion is not significant, while the one for the firms above the median level of contagion is significant at any standard level of significance. In particular, the average change in the levels of short positions in the subset of non-financial companies more heavily exposed to the sovereign risk increased by 0.012 percentage points after the ban.

From the previous results we learn that the ban on short-selling seemed to have unleashed a singular stabilising effect on the CDS of medium-sized banks that would have taken place by partially isolating these banks from the pressure coming from the public debt market. Similar qualitative results are found for the case of the two largest Spanish banks, although the economic and statistical significance of the previous stabilising effect is significantly weaker than in the case of the smaller banks.

5. Market performance under the short-selling ban

In this section we supplement our previous analysis of the effects of the ban on short-selling by providing some estimates of the impact of this measure on several dimensions of market performance, including liquidity, trading volume, returns volatility, price discovery and excess returns. As a benchmark against which to assess the impact of the ban on the previous

variables, we take the group of companies included in the Ibex 35 that were not subject to the ban. As in section 4.2, we exploit information from three months centred on the inception date (except for the case of cumulative returns).

The panels of figure 4 contain the proxies used to capture the several market variables of interest. To test for the statistical significance of the patterns shown therein, we perform several event-study regressions. In particular, we regress the proxies for volatility, liquidity, volume and returns on three dummies that refer to three different groups of firms considered (large banks, medium-sized banks and non-financial firms included in the Ibex 35). These dummies take the value of one during the post-ban period and zero during the pre-ban period (i.e. one month and a half after the ban in each case). These estimates are obtained using OLS with robust standard errors and including fixed effects for each firm and day effects using a time trend variable. The coefficients of this estimation are shown in panel A of table 4, while panel B reports the differences for such coefficients and the signs that indicate whether the incremental effect of the ban is significantly different from zero. The null hypothesis to test in panel B in the case of volatility and volume (relative bid-ask spreads and returns) is whether the effect of the ban obtained for the first firm in each cell of column 1 is significantly higher (lower) than the effect for the second firm in the cell.

< Insert Figure 4 here >

Returns. The first apparent effect of the ban is the outperformance of financial corporations' shares relative to those of non-financial firms included in the Ibex 35, in terms of the accumulated returns from the date of the ban (see top-left panel of figure 4). Also, prices seem to react differently depending on the size of the banks in the sample, with a longer-lived boost effect in the case of the medium-sized banks (see top left panel of figure 4). However, in any of the cases the positive effect was temporary and had faded considerably around two weeks after the inception of the ban. Furthermore, as revealed by the differences-in-

differences estimates contained in panel B of table 4, the hypothesis that the ban had a significantly stronger effect on the banks, independently of their size, than on the non-financial firms can be rejected at a 1% significance level. This result echoes those of Beber and Pagano (2013) and Boehmer, Jones and Zhang (2011), who fail to find a significant distinctive effect of short-selling constraints on the price of the stocks affected by them vis-à-vis other stocks unaffected.

Volatility. The disparity found in the preceding section in terms of the unequal impact of the ban on the various indicators of default risk between large and medium-sized banks carries over in terms of their volatility. Specifically, while the average relative returns volatility of the two largest banks was apparently unaffected, that of medium-sized banks died down considerably. As illustrated in the top right panel of figure 4, the ban had a strongly moderating impact on the price fluctuations of medium-sized banks' shares, whose volatility readings dropped below those of the companies in the Ibex 35 not covered by the ban. Conversely, the relative volatility of the largest cap banks showed little variation. In fact, it increased two weeks after the ban and reached higher levels than those observed around the inception of the ban one month after this date. The estimates in column 1 of panel A in table 4 show that following the ban there was a significant decrease in the volatility of the returns of the medium-sized banks, that was also perceived in the non-financial firms but not in the largest banks. The estimates in column 1 of panel B in table 4 confirm that the ban contributed to dampen significantly the fluctuations in the returns of medium-sized banks relative to the other two groups of firms considered.

Liquidity. One of the most visible consequences of the ban was the intense worsening of the liquidity conditions of medium-sized banks. Specifically, the bid/ask spreads of firms in this group widened persistently after the onset of the ban compared with those of non-financial Ibex 35 members (see lower left panel of figure 4). Concretely, the ban led to an average

increase of around 0.17% and 0.04% of the relative bid-ask spread of the medium-sized banks and non-banned firms, respectively. In the case of the medium-sized banks, the average relative bid-ask spread was 0.34% at the end of June 2011 (i.e. at the beginning of the sample period) and so, the increase following the ban was around 50% of the level prevailing well before the ban. When compared with the other two groups, the previous drop in liquidity was statistically significant at a 1% confidence level (see panel B of table 4).

Trading volume. The previous fall in the liquidity of the shares of medium-sized banks seems linked to a slump in these stocks' trading volumes, which fell to nearly 70% in the weeks following the ban relative to the flow of trades at the time of its inception (see lower right panel of figure 4).¹² The maximum drop in trading volumes from immediate pre-ban levels was just over 40% in the case of the banking majors and 20% among non-financial corporations. These post-ban differences in trading volumes between firm groups were statistically significant (see panel B of table 4).

Price discovery. In tune with their deteriorating liquidity and trading conditions, the speed of the price adjustment of medium-sized bank stocks was curtailed following the entry of the ban. To estimate this effect, we calculate the speed of share price adjustment as the average difference (in percentage) between the first-order correlations of the component of daily share returns that is not explained by overall market performance before and after the start of the ban. Concretely, we compare the first-order correlations for the 35 trading days before and after the ban. Thus, a positive value indicates deterioration in the price discovery process in each category. In the case of medium-sized banks the percentage change in the speed of share price adjustment after the ban is 8.4%. In the case of the largest banks and non-financial

¹² The average value of the standardised turnover volume in medium-sized banks in the quarter around the time of the ban was 0.9 units; being equal to 1.25 units before the inception of the ban and 0.60 after it. Nevertheless, the variation in the volume of the medium-sized banks shares was remarkable in the two previous periods. The volatility and the maximum value for the indexed volume were 2.35 and 30.73 (0.94 and 10.42) units, respectively, before (after) the short sales constraint.

corporations, the apparent loss of price information efficiency was considerably less severe. The percentage change in the speed of share price adjustment was 1.6% and 2.4% for the large banks and the non-financial corporations, respectively. Thus, beyond these differences across firm groups, we find support for the hypothesis that short sales constraints hinder price discovery, which is a common result in the previous empirical studies on this issue.

6. Conclusions

This paper analysed the main effects of the short sales ban implemented in August 2011 in the Spanish stock market along two dimensions: financial stability and market performance.

As regards the stabilising effects of the ban, we found that short positions were a significant determinant of the spreads of the CDS of medium-sized banks before the ban, whose relatively low degree of geographical diversification and higher perceived implicit and explicit reliance on public aid rendered them more vulnerable to surges in domestic macroeconomic and sovereign risk. Subsequently, by weakening the contagion effect coming from the sovereign risk, the ban helped stabilise the probability of default of this group of banks. By way of contrast, we failed to find evidence of a significant effect of short sales on the indicators of credit risk for the case of the largest banks and the non-financial firms before the ban, arguably less so exposed in both cases to sovereign risk than the medium-sized banks. Likewise, the ban did not convey a significant stabilising effect on the credit risk indicators of the largest banks. An interesting finding is that following the ban on short sales in financial stocks the short positions in non-financial companies rose significantly, which reflected the presence of a common market-wide risk factor that is embedded in the sovereign CDS.

Interestingly, the previous asymmetry regarding the effect of the ban on the credit risk indicators of the different firm groups does not extend to the case of stocks returns, in the

sense that the ban did not give rise to significant differences in the returns of the shares between the medium-sized banks, the large banks and the non-financial companies. Based on these findings, we argue that the effectiveness of a ban on short sales aimed at supporting financial stability should be better assessed against its power to stabilise the financial resilience of the firms targeted by the ban (proxied here through their CDS spreads), rather than merely on the basis of its effect on relative stock prices.

Nonetheless, the previous stabilising power of the ban came at the cost of a significant decline in the liquidity, trading volumes and price information efficiency of medium-sized banks' stocks. In short, such deterioration was significantly more intense than the one caused in the rest of the stocks. This last result suggests a trade-off between the effectiveness and the efficiency costs of this short-selling constraint, in the sense that the power of the ban to support the financial strength of a subset of banks unleashed important damages in the liquidity, volume of trading and price discovery of their stocks.

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Appendix

The spillover or contagion effects are obtained from a variance decomposition associated with an N -variable vector autoregression (VAR) following the methodology employed in Diebold and Yilmaz (2012). These authors measure directional spillovers in a generalised VAR framework that eliminates the possible dependencies of results on ordering. In particular, we first consider a covariance stationary N -variable VAR (p):

$$X_t = \sum_{i=1}^p \Phi_i X_{t-i} + \varepsilon_t \quad (A.1)$$

where X_t denotes a vector of stationary changes in the CDS spreads and $\varepsilon \sim (0, \Sigma)$ is a vector of independently and identically distributed disturbances such that the moving average

representation is $X_t = \sum_{i=0}^{\infty} A_i \varepsilon_{t-i}$, where the $N \times N$ coefficient matrices A_i obey the recursion

$$A_i = \Phi_1 A_{i-1} + \Phi_2 A_{i-2} + \dots + \Phi_p A_{i-p}, \text{ with } A_0 \text{ an } N \times N \text{ identity matrix and } A_i = 0 \text{ for } i < 0. \text{ Thus,}$$

the error from the forecast of X_t at the H -step-ahead horizon, conditional on information

available at $t-1$, can be expressed as $\xi_{t,H} = \sum_{h=0}^H A_h \varepsilon_{t+H-h}$, and the variance-covariance matrix of

the total forecasting error is computed as $Cov(\xi_{t,H}) = \sum_{h=0}^H A_h \Sigma A_h'$, where Σ is the variance-

covariance matrix of the error term in equation (A.1), ε_t .

We rely on variance decomposition of the moving-average coefficients, which allows us to parse the forecast error variances of each variable into parts attributable to the various

system shocks.¹³ By means of this decomposition we can obtain the proportion of the H -step-ahead error variance in forecasting X_i that is due to shocks to X_j , $\forall j \neq i$, for each i .

We first compute the variance shares which are defined as the fractions of the H -step-ahead error variances in forecasting X_i due to shocks to X_i , for $i= 1, 2, \dots, N$, and cross variance shares, or spillovers as the fractions of the H -step-ahead error variances in forecasting X_i due to shocks to X_j , for $i, j = 1, 2, \dots, N$ such that $i \neq j$. The H -step-ahead forecast error variance decompositions are denoted by $\theta_{ij}^g(H)$, for $H = 1, 2, \dots$:

$$\theta_{ij}^g(H) = \frac{\sigma_{ii}^{-1} \sum_{h=0}^{H-1} (e_i' A_h \Sigma e_j)^2}{\sum_{h=0}^{H-1} (e_i' A_h \Sigma A_h' e_i)} \quad (A.2)$$

where Σ is the variance matrix for the error vector ε , σ_{ii} is the standard deviation of the error term for the i^{th} equation and e_i is the selection vector with one as the i^{th} element and zeros elsewhere. The sum of the elements of each row of the variance decomposition table is not equal to 1, i.e. $\sum_{j=1}^N \theta_{ij}^g(H) \neq 1$. Each entry of the variance decomposition matrix can be

normalized such that the elements of each row sum 1 as:

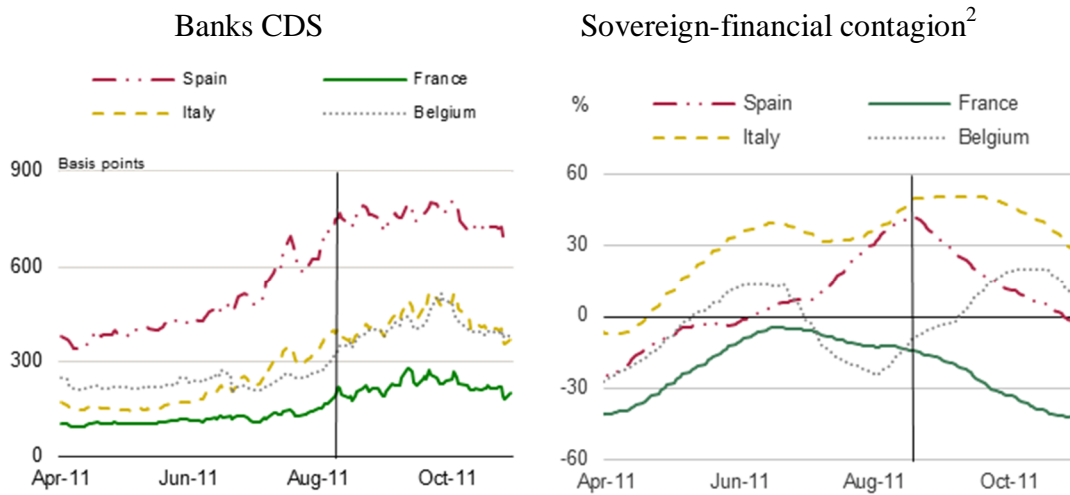
$$\tilde{\theta}_{ij}^g(H) = \frac{\theta_{ij}^g(H)}{\sum_{j=1}^N \theta_{ij}^g(H)} \quad (A.3)$$

This methodology is used to estimate the spillovers between the sovereign and the corporate CDS spreads. The spillovers show the degree of variation in the changes of the corporate (sovereign) CDS spreads, which is not due to the historical information of the changes in the

¹³ The variance decomposition requires orthogonal innovations, but the VAR innovations are generally contemporaneously correlated. The use of the generalised VAR instead of the Cholesky factorisation avoids using arbitrary ordering assumptions.

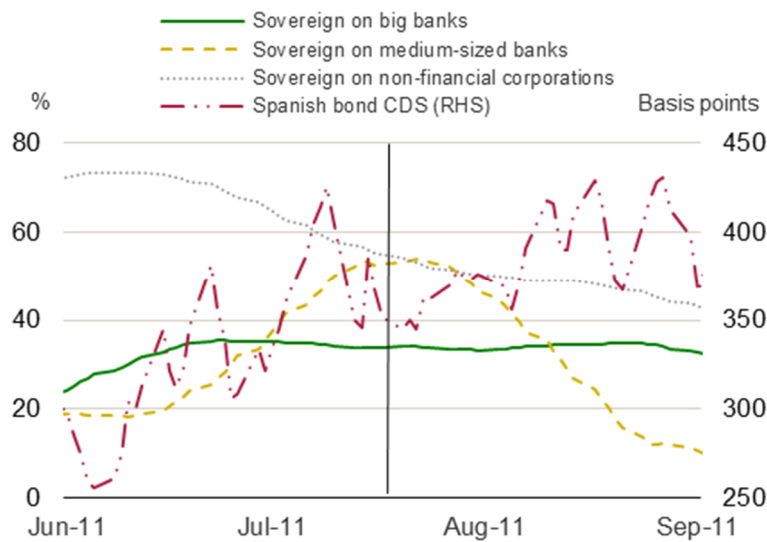
sovereign and the corporate CDS spreads but to shocks (innovations) in the changes of the sovereign (corporate) CDS spreads. This indicator takes higher values as the intensity of the contagion effect, caused by the specific shocks of the corporate (sovereign) CDS premium increases. In the extreme case in which there are no spillovers from one CDS premium to the other, the indicator series is equal to zero.

Figure 1: Banks CDS and contagion between sovereign and financial risk¹



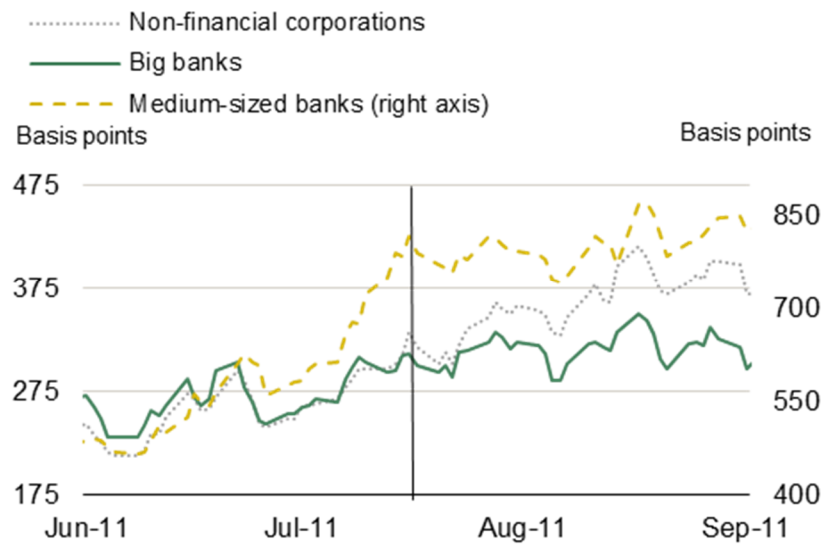
(1) The vertical line marks the start date of the short-selling ban in Spain (12 August 2011).
 (2) The figure shows the difference between the net percentage change in banking sector CDS index (see below) that is not attributable to their historical information but to contemporaneous shocks in sovereign credit risk, and the net change in the opposite direction. The indicator is positive when the impact of sovereign risk shocks on financial risk indicators is higher than vice versa. The banking sector CDS index of each country responds to the average of the CDS of its banks. The spillover on a given date is calculated from available data for the 60 preceding days. In addition, the series are filtered with a 30-day moving average. See Appendix for further details on the methodology.

Figure 2: Sovereign risk contagion in Spain^{1,2}



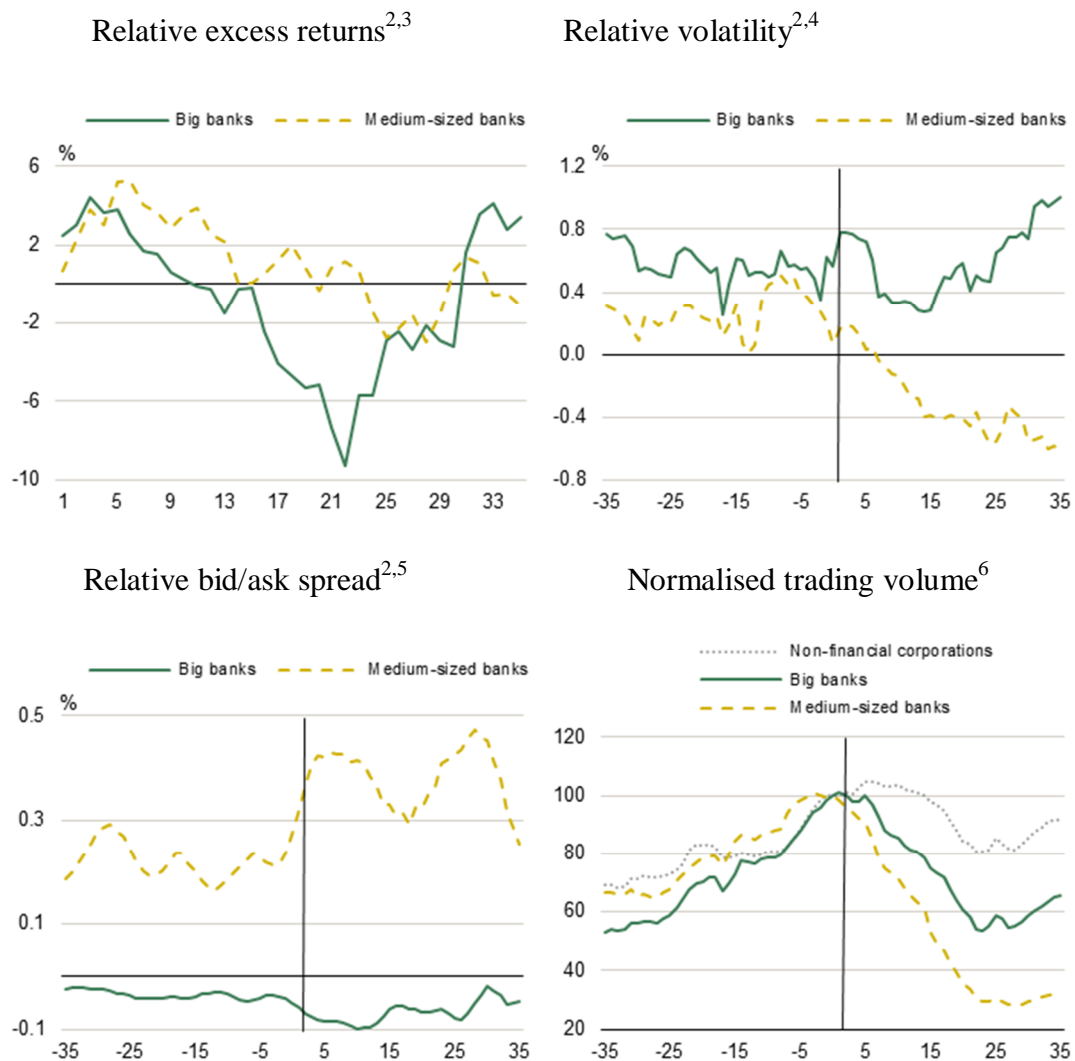
(1) The vertical line marks the date the ban was introduced in Spain (12 August 2011).
 (2) The figure shows the difference between the net percentage change in the CDS of the three groups of firms that is not attributable to their historical information but to contemporaneous shocks in sovereign credit risk, and the net change in the opposite direction. The indicator is positive when the impact of sovereign risk shocks on companies risk indicators is higher than vice versa. The CDS index for each of the three groups of companies is obtained as the average of the CDS spreads of the firms in each group. The contagion indicator on a given date is calculated from available data for the 60 preceding days. In addition, the series is filtered with a 30-day moving average. See Appendix for further details on the methodology.

Figure 3: CDS spreads of Spanish listed companies¹



(1) The vertical line marks the start date of the short-selling ban in Spain (12 August 2011).

Figure 4: Liquidity, trading volume, volatility and stock prices¹



- (1) The X-axis is a time scale in which 0 marks the start date of the prohibition.
- (2) Relative to the set of companies in the Ibex 35 not subject to the ban.
- (3) The relative excess return of each group on each date is the average of the cumulative spread to that date in the daily returns of component corporations versus the average return of the Ibex 35.
- (4) The volatility of each group on each date is calculated by reference to the average standard deviation of the share returns of component corporations in the preceding 22 trading days.
- (5) Bid/ask spreads are calculated for each group as the average ratio of the bid/ask spreads of the share prices of component corporations to the average bid/ask spread of the group. Series are also filtered with the moving average of the last five trading sessions.
- (6) This indicator represents the cumulative sum of the 22 previous average daily trading volumes of the shares of firms in each group as normalised to 100 at the time of the ban.

Table 1: Effect of short positions on the CDS of Spanish listed companies before the ban

This table reports the effects of the short positions in the Spanish listed companies on the changes in their CDS by the type of firm (medium-sized banks, large banks and non-financials) before the ban. The coefficients for the effects of the volume of short positions and the additional controls are estimated using data from March 2011 to December 2011 on the basis of a fixed effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation. The first column reports the effect of the short positions relative to capital and their incremental effect before the ban for each firm group. The second column reports the differential effects for the three firm groups for the period before the ban without being interacted with the short positions. The third column reports the effect of the volume of stocks lending relative to capital and their incremental effect before the ban. The fourth column reports the effect of the short position for a sample restricted to the pre-ban period (March 2011 – August 2011). All the variables are defined in percentages with the exception of the debt purchased by the ECB which is defined in billions of euros. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors are reported between brackets.

Dependent variable: $\Delta\log(\text{CDS}(t))$	(1)	(2)	(3)	(4)
Short positions (t-1), total sample	0.407 (0.24)	0.458* (0.23)		
Stocks lending (t-1), total sample			0.104 (0.06)	
Short positions (t-1), non-financials before ban	-0.071 (0.11)			-0.757 (1.76)
Dummy non-financials before the ban		-0.001 (0.01)		
Stocks lending (t-1), non-financials before the ban			0.011 (0.03)	
Short positions (t-1), medium-sized banks before ban	0.226** (0.09)			0.809* (0.40)
Dummy medium-sized banks before ban		0.007 (0.01)		
Stocks lending (t-1), medium-sized banks before ban			0.065* (0.03)	
Short positions (t-1), large banks before ban	-8.016 (5.88)			-10.631 (5.85)
Dummy large banks before ban		-0.001 (0.01)		
Stocks lending (t-1), large banks before ban			-0.003 (0.04)	
Dummy before ban	0.000 (0.01)		-0.003 (0.01)	
Dummy medium-sized banks	-0.022 (0.01)	-0.024 (0.01)	0.001 (0.00)	-0.027 (0.02)
Dummy big banks	0.001 (0.00)	0.000 (0.00)	-0.003 (0.00)	0.003 (0.00)
$\Delta\log(\text{CDS}(t-1))$	-0.108** (0.04)	-0.107** (0.04)	-0.114** (0.04)	-0.084 (0.06)
Stock returns (t-1)	-0.156* (0.07)	-0.155* (0.07)	-0.169* (0.07)	-0.197 (0.12)
Squared of stock returns (t-1)	0.065 (0.08)	0.066 (0.08)	0.032 (0.08)	0.009 (0.15)
$\Delta\log(\text{VIX}(t-1))$	0.034 (0.02)	0.034 (0.02)	0.031 (0.02)	-0.013 (0.03)
$\Delta\log(\text{Spanish sovereign CDS}(t-1))$	0.236** (0.04)	0.237** (0.04)	0.239** (0.04)	0.244** (0.06)
ECB Bond purchases since 8 August 2011	-0.00004 (0.00)	-0.00004 (0.00)	-0.00003 (0.00)	0.001 (0.00)
$\Delta\log(\text{CDS bid-ask spread})$	0.008* (0.00)	0.008* (0.00)	0.009* (0.00)	0.008 (0.01)
Constant	0.001 (0.01)	0.002 (0.01)	-0.001 (0.01)	0.000 (0.00)
Fixed effects	YES	YES	YES	YES
Observations	3,767	3,767	3,547	1,790
Number of companies	17	17	17	17
Adj. R-squared	0.11	0.11	0.11	0.08

Table 2: Effect of short positions on the default probabilities and implied ratings of Spanish listed companies before the inception of the ban

This table reports the effects of the short positions in the Spanish listed companies on the changes in their default probabilities and implied ratings distinguishing the type of firm (medium-sized banks, large banks and non-financials) before the inception of the ban. The coefficients for the effects of the volume of short positions and the additional controls are estimated using data from March 2011 to December 2001 on the basis of a fixed effects regression with the standard errors robust to heteroskedasticity and within panel correlation. The first column reports the effect of short positions relative to capital and their incremental effect on the first differences of the logit transformed default probabilities for all the firms with available information, before the ban for the following four groups of institutions: medium-sized banks, large banks, other financials and non-financials. The second, third and fourth columns report the differential effects on the first difference of the implied rating and two variations of the implied ratings before the ban for these four types of institutions. All the variables are defined in percentages with the exception of the debt purchased by the ECB, which is defined in billions of euros. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors are reported between brackets.

	Dep. var. Δ logit transformed PD	Dep. var. Δ implied ratings(t)		
	(1)	(2)	(3)	(4)
Short positions (t-1), total sample	-0.076 (0.26)	0.282 (0.32)	0.977 (1.29)	0.000 (0.00)
Short positions (t-1), non-financials before ban	0.008 (0.16)	0.373* (0.17)	1.114 (0.58)	0.001 (0.00)
Short positions (t-1), medium-sized banks before ban	0.254** (0.08)	0.635** (0.24)	3.025** (1.18)	0.006** (0.00)
Short positions (t-1), large banks before ban	2.750 (7.05)	-0.163 (2.70)	-1.105 (9.54)	0.002 (0.02)
Short positions (t-1), other financials before ban	0.628** (0.15)	1.656** (0.22)	6.546** (0.84)	0.017** (0.00)
Dummy before ban	0.000 (0.00)	-0.008 (0.01)	-0.022 (0.02)	-0.001 (0.00)
Dummy medium-sized banks	-0.008 (0.01)	0.063** (0.02)	0.300** (0.09)	-0.021* (0.01)
Dummy big banks	-0.006 (0.02)	0.035 (0.03)	0.116 (0.11)	0.007 (0.02)
Dummy other financials	-0.011 (0.02)	0.030 (0.03)	0.046 (0.11)	0.001 (0.02)
Δ Dependent variable (t-1)	-0.222** (0.04)	-0.258** (0.01)	-0.250** (0.02)	-0.258** (0.01)
Δ log(VIX (t-1))	0.057** (0.01)	0.164** (0.02)	0.595** (0.10)	0.094** (0.02)
Δ log(Spanish sovereign CDS (t-1))	0.054** (0.02)	0.098* (0.04)	0.425* (0.18)	0.095** (0.04)
ECB Bond purchases since 8 August 2011	1.00E-05 (0.00)	-7.00E-05 (0.00)	-1.80E-04 (0.00)	-1.00E-05 (0.00)
Constant	0.006 (0.02)	-0.023 (0.03)	-0.070 (0.11)	0.000 (0.02)
Fixed effects	YES	YES	YES	YES
Observations	24,037	23,999	23,999	23,999
Number of companies	105	105	105	105
Adj. Rsquared	0.05	0.06	0.06	0.06

Table 3: Effects of the ban on CDS spreads and sovereign-corporate credit risk contagion

This table reports the estimation of the effects of the ban on the variations in the sovereign-corporate credit risk contagion indicator for the three firm groups (large banks, medium-banks, and non-financial companies). The estimations exploit data from a three-month period centered at the time of the ban. The estimated parameters are shown in Panel A. The dependent variable of the regression corresponding to column 1 is the first difference in the contagion indicators (in percentages). Column 2 contains the results obtained when the dependent variable is the first difference of CDS spreads. The explanatory variables are the same in the two columns: three dummies for the type of firm that take the value of one during the post-ban period and zero before. The estimates are obtained using Ordinary Least Squares (OLS) fixed-effects regression with the standard errors robust to heteroskedasticity and cross-sectional correlation. Panel B reports the estimated difference for the coefficients reported in each column of panel A and whether the incremental effects of the ban for the three types of institutions are significantly different from zero. The null hypothesis to test in Panel B is whether the effect of the ban obtained for the first firm in each cell in the first column is significantly stronger (more negative coefficient) than the effect for the second firm. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors in Panel A and p-values in Panel B are reported between brackets.

Panel A	Δ Sovereign Credit Risk Spillovers (%)	Δ CDS Premium (basis points)
Dummy medium-sized banks after ban	-1.498* (0.76)	-9.750* (4.91)
Dummy large banks after ban	-0.485 (0.45)	-1.999 (3.05)
Dummy non-financials after ban	0.479 (0.33)	-2.103 (3.21)
Constant	0.037 -0.314	5.762** (2.08)
Fixed effects	YES	YES
Time effects	YES	YES
Observations	1,278	1,216
Number of companies	18	18
Adj. R-squared	0.01	0.02

Panel B	Δ Sovereign Credit Risk Spillovers (%)	Δ CDS Premium (basis points)
Diff of coefficients	H ₀ (Diff of coeff ≤ 0)	H ₀ (Diff of coeff ≤ 0)
Medium-sized banks - Big-sized banks	-1.013* (0.05)	-7.751* (0.03)
Medium-sized banks - Non-financial corporations	-1.976** (0.01)	-7.647* (0.02)
Big-sized banks - Non-financial corporations	-0.964* (0.02)	0.104 (0.52)

Table 4: Effects of the ban on the volatility, liquidity, volume and returns of the shares of large banks, medium-sized banks and non-financial firms

This table reports the estimated effects of the ban on the stock's returns, volatility, liquidity and volume for three different types of Ibox 35 firms (large banks, medium-sized banks and non-financials). The coefficients of the estimation are shown in Panel A. The dependent variable in column 1 is the firm's daily stock returns. Column 2 contains the results obtained when the dependent variable is the realised volatility of the stock prices that is proxied by means of the square of the firm's stock returns. The dependent variable of the regression corresponding to column 3 (liquidity) is the relative bid-ask spread. Column 4 contains the results obtained when the dependent variable is the trading volume normalised to 100 at the time of the inception of the ban for all the firms in the sample. The explanatory variables are the same in the four columns: three dummies for the different types of firms that take the value of one during the post-ban period (one month and a half after the ban) and zero before. The estimates are obtained using Ordinary Least Squares (OLS) with robust standard errors and including fixed effects for each firm and day effects using a time trend variable. Panel B of table 3 reports the difference for the coefficients reported in panel A and whether the incremental effects of the ban for the three types of institutions are significantly different from zero. The null hypothesis to test in panel B in the case of the volatility and volume (returns and relative bid-ask spreads) is whether the effect of the ban obtained for the first firm in each cell of the first column is significantly higher (lower) than the effect for the second firm. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors in panel A and p-values in panel B are reported between brackets.

Panel A	Returns	Squared returns	Relative bid-ask	Volume (1 = date of ban)
Dummy medium-sized banks after ban	0.01559** (0.003)	-0.00058** (0.000)	0.00165** (0.000)	-1.06664** (0.170)
Dummy large banks after ban	0.01786** (0.006)	-0.00014 (0.000)	0.00014 (0.000)	-0.49044** (0.048)
Dummy non-financials after ban	0.01486** (0.003)	-0.00033** (0.000)	0.00041* (0.000)	-0.40786** (0.041)
Constant	0.03920** (0.009)	-0.00174** (0.000)	0.00066 (0.001)	-0.43146* (0.172)
Fixed effects	YES	YES	YES	YES
Time effects	YES	YES	YES	YES
Observations	2308	2308	2308	2308
Number of companies	35	35	35	35
Adj. R-squared	0.02	0.03	0.06	0.06

Panel B	Returns	Squared returns	Relative bid-ask	Volume (1 = date of ban)
Diff of coefficients	H ₀ (Diff of coeff ≤ 0)	H ₀ (Diff of coeff ≥ 0)	H ₀ (Diff of coeff ≤ 0)	H ₀ (Diff of coeff ≥ 0)
Medium-sized banks - Large banks	-0.00227 (0.649)	-0.00044* (0.029)	0.00151** (0.000)	-0.57620** (0.000)
Medium-sized banks - Non-financials	0.00073 (0.386)	-0.00025** (0.006)	0.00124** (0.000)	-0.65878** (0.000)
Large banks - Non-financials	0.00300 (0.299)	0.00019 (0.802)	-0.00027 (0.995)	-0.08258** (0.008)