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AND FINANCIAL STABILITY: EVIDENCE
FROM THE SPANISH MARKET**

2014

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**Documentos de Trabajo
N.º 1410**

BANCO DE ESPAÑA
Eurosistema



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ISSN: 1579-8666 (on line)

Abstract

We examine the effect of the short-selling ban in 2011 on Spanish stocks on the level of risk in the banking sector. Before the ban, short positions were found to be positive and significantly related to the creditworthiness of medium-sized banks, these being generally less internationally diversified and more reliant on official support. We show that the ban helped stabilise the credit risk of medium-sized banks, especially those more exposed to short-sellers' activity, but not that of large banks and non-financial corporations. This stabilising effect came at the cost of a significantly sharp decline in liquidity, trading and price efficiency of medium-sized banks' stocks relative to other stocks.

Keywords: short-sales constraints, financial stability, financial institutions, credit default swap, contagion.

JEL Classification: G01, G12, G14, G18.

Resumen

En este artículo se analiza el efecto que tuvo sobre el riesgo de crédito la restricción a la operativa de ventas en corto aplicada en 2011 a las acciones de un conjunto de entidades financieras españolas. En el periodo que precedió a la entrada en vigor de la prohibición sobre las ventas en corto, las posiciones cortas sobre estos valores estaban correlacionadas, significativamente, con los indicadores de solvencia financiera de las entidades de crédito de tamaño medio, que, en general, estaban menos diversificadas geográficamente y presentaban un mayor grado de dependencia, explícita o implícita, del apoyo público. En el artículo se muestra que la prohibición de ventas en corto contribuyó a moderar el riesgo de crédito de las entidades de tamaño medio, especialmente de aquellas más expuestas a este tipo de práctica, pero no el de aquellas otras entidades financieras de gran tamaño ni tampoco el relativo a las sociedades no financieras. El efecto estabilizador del riesgo de crédito de la prohibición, sin embargo, vino de la mano de una cierta pérdida de calidad de funcionamiento del mercado de valores, de modo que la liquidez, el volumen de negociación y la eficiencia de los precios de las acciones de las entidades de crédito de tamaño medio se resintieron con motivo de la aplicación de esta medida excepcional.

Palabras claves: restricciones a las ventas en corto, estabilidad financiera, instituciones de crédito, *credit default swaps*, contagio financiero.

Códigos JEL: G01, G12, G14, G18.

1 Introduction

On 11 August 2011, four European supervisors (in Belgium, France, Italy and Spain) announced a ban on establishing or increasing short positions in financial institutions' stocks. This measure was adopted at a time of heightened aggregate uncertainty and volatility in financial markets that coincided with the resurgence of tensions in the sovereign debt markets of various euro area countries and a rapid spread of public debt market tensions to some key wholesale markets. The Spanish regulator (CNMV), 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, safeguarding financial stability was the leading motive behind its decision to prohibit temporarily short positions in the shares of 16 Spanish financial corporations.

In practice, however, little is known about the impact of restrictions on short positions on the stability of the financial sector. In this paper, we analyse the implications of the ban imposed by the Spanish regulator in 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 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 and idiosyncratic default risks 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 offering 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 of banks' net interest income and loan book quality, again bearing down on their stock market prices.

It is also conceivable that, in some circumstances, an increase in short positions in financial stocks could magnify the effect of any disturbances hitting the sector, in which case curtailing this kind of trade could, in theory, support the stability of affected institutions. Additionally, the increase in short positions could be reflecting strategic trading by some traders that induce or amplify stocks' adverse price movements in financial firms that are more financially constrained to gain from their subsequent re-sales (Brunnermeier and Pedersen, 2005). As a consequence, these effects would hamper the banks' access to financing, thus raising their financial vulnerability. This argument (or a variation of it) is often invoked by regulators to justify restrictions on short positions; yet, in practice, the impact of short-sale bans on financial stability remains largely unexplored.

In order to approximate the ban's power to support financial stability, we first offer empirical evidence that shows a positive relationship between short positions and the credit risk indicators of medium-sized Spanish banks in the weeks preceding the ban. We then find that restricting short sales on these stocks had some stabilising influence on the market perception of the banks' credit risk, a result that is found to be insignificant for the country's

two largest banks and for non-financial firms. In particular, the ban then helped stabilise the market perception of medium-sized banks' credit risk levels, despite sovereign credit risk heading higher. 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. 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 reflect the higher reliance of the former on the implicit and explicit official support and the higher relative holdings of government bonds that would damage their balance sheets. Additionally, the two largest banks are internationally diversified and perceived as much better capitalised than most of the medium-sized banks and, hence, less dependent on implicit or explicit public guarantees and with a higher reliance on the domestic market and, therefore, on domestic risk. Interestingly, we find that the volume of short sales accumulated on medium-sized banks before the ban is a significant explanatory factor of the stabilising effect of the ban on CDS, i.e. short positions seemed to cause *per se* a rise in risk indicators. We also document a 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.

The previous stabilisation effects of the ban on the credit risk indicators of medium-sized banks, however, 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 more of 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 relationship of short sales to the creditworthiness of companies listed in the Spanish market before the ban. Section 5 examines the effects of the ban on the market perception of corporate credit risk and on the short-sales activity of non-financial firms. Section 6 examines the effect of the ban on the Spanish stock market in terms of liquidity, volatility, excess returns and price discovery. Section 7 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-sale 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.

2.1 Short-sale constraints and financial stability

The potential linkage between short-selling of financial stocks and banking stability has recently been explored from a theoretical standpoint in several articles. Brunnermeier and Pedersen (2005) analyse how short-selling may exacerbate fluctuations in asset 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 asset prices by allowing opportunistic investors to mimic the moves of other traders under pressure to liquidate their investments. A direct testable implication, which we do in fact test with our data, is that a short-selling constraint, given everything else, should contribute to reducing the volatility of the price of the stocks affected by it.

Brunnermeier and Oehmke (2013) explore a setting where financial institutions are subject to leverage constraints. In this environment, it is shown that shorting aggressively a financial stock may lead to an inefficient equilibrium in which the firm is liquidated, a result that is interpreted as a rationale for temporary bans on short positions in stocks of vulnerable financial institutions.

Liu (2014) also 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 share's price about the firm's underlying fundamentals, become increasingly unsure about the true fundamentals as the share price turns more volatile. Thus, they become less willing to maintain their exposure to the bank in question. As this happens on a sufficiently large scale, "too noisy" stock prices may end up triggering a bankruptcy. In anticipation of this, speculators thus find it optimal to short-sell the stock beforehand to amplify price volatility. As a result, they crowd out lenders and, therefore, increase the likelihood of benefiting from the firm's collapse.¹

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.

1. 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.

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. In this vein, our paper aims to provide empirical evidence on the effect of short-sale constraints on the market perception of the banks' credit risk and, therefore, on financial stability. To this end, we first document the strong correlation between short positions and some indicators of the default likelihood of medium-sized banks in the months preceding the inception of the short-selling ban by the Spanish supervisor in August 2011. A key piece of information used in this analysis to measure this short-sales activity is a proprietary CNMV dataset that consists of the series of short positions exceeding 0.2% of the total capital of every firm in the Spanish market. Next, we assess the effect of the short-selling ban on the previous indicators of financial vulnerability, distinguishing by type of firm: medium-sized banks, large banks, and non-financial firms.

2.2 The effects of short-sale 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 January 2008-June 2009 period, 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 (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 share 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-sale 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 internalise the information that, in expectation, would otherwise be left out by the prohibition. This divergence of conclusions extends to the empirical 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 the fact 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 actually 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 the volatility of returns 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 concerns 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 in previous dimensions of market performance across different groups of firms, based on certain indicators of size or trading activity (see e.g. Boehmer et al. 2011), we find, interestingly, that such differences extend to every dimension analysed here (liquidity, volatility of returns and price discovery), except relative stock 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 neatly 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 greater power of the ban to support the financial strength of a sub-set of banks unleashed major damages in the liquidity, volume of trading and price discovery of their stocks.

3 Data

We exploit daily information on short positions and stock lending trading on the Spanish continuous market (SIBE) from 1 March 2011 to 31 December 2011 (i.e. around five months before and after the ban). 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.² The threshold for public disclosure of net short positions is set at 0.5% and each 0.1% above that.

The data on CDS spreads on sovereign and corporate debt are obtained from Credit Market Analysis (CMA). This information is available for 17 of the Spanish firms in the continuous stock market, which include medium-sized banks (Banco Pastor, Banco Popular, Banco Sabadell, Bankia, Bankinter and Caixabank),³ the two largest banks (BBVA and Santander) and non-financial firms (Abertis, ArcelorMittal, Endesa, Gas Natural, IAG, Iberdrola, Melia Hoteles, Repsol, and Telefonica). These 17 firms represent most of the trading volume in the Spanish stock market (almost 80% of the total daily volume of trade over the sample period). CDS spreads are used as indicators of the corporate risk premiums and as an input to compute the spillovers of credit risk between the sovereign and the financial sectors. In addition, in the interest of robustness, we use alternatively default probabilities that are obtained from the StarMine Structural Credit Risk model through the Reuters platform for a wider group of firms. 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 stock prices, including the bid and ask prices, and the trade volume are obtained from Datastream. The information on the Ibex 35, EuroStoxx 50 and the volatility indexes VIX and VSTOXX also come from Datastream. Finally, the information on European Central Bank (ECB) bond purchases was obtained from the ECB webpage.

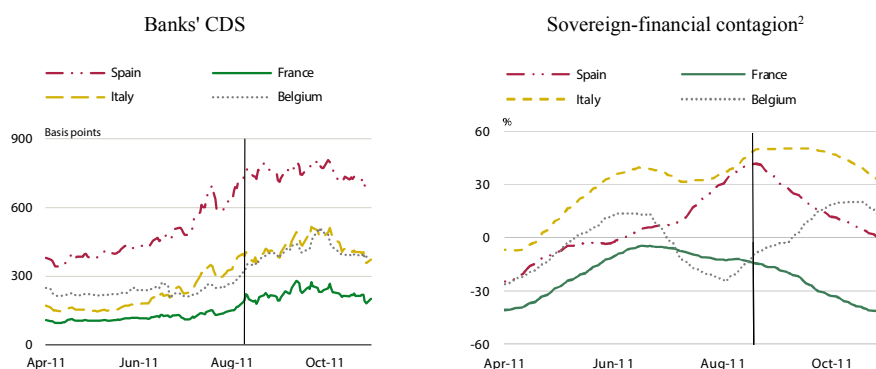
2. Uncovered short positions, in which the seller has not ensured the availability of the securities, are prohibited in Spain.

3. Banesto is not included among the medium-sized banks because bid and ask CDS prices are not available.

4 Short sales and credit risk in the pre-ban period

Towards the end of July 2011, many European banks suffered a sharp run-down in their share prices and a parallel rise in their credit risk indicators (see left panel of Figure 1). The surge in financial sector risk coincided with a jump in the indicators of risk-contagion 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 (see right panel of Figure 1).⁴ This last effect added to the deterioration in the quality of the non-sovereign assets held by many European banks, putting further pressure on their capital ratios and hampering their access to financing. At the same time, short sales of financial stocks built up faster than was the case for non-financial firms. In particular, the average ratio of short positions on medium-sized banks increased 32.4% from 1 March 2011 (the beginning of our sample) to 8 August 2011 (i.e. from 1.08% to 1.43% as a proportion of capital).

Figure 1 : Bank s' CDS and contagion between sovereign and financial risk¹



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

(2) The figure shows the difference between the net percentage change in the banking sector CDS index (see below) that is not attributable to the related 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, and vice versa. Each country's banking sector CDS index responds to the average of its banks' CDS. 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.

Based on the previous observations, we next analyse in detail the relationship between short positions and several credit risk indicators of the companies listed in the Spanish market. In principle, a positive relationship between short sales and credit risk could be indicative of the fact that short sellers are informed traders who may anticipate rating downgrades by exploiting market-based credit-risk indicators, as argued by Henry et al. (2014). Alternatively, a positive link between both variables could reflect some form of causality from short sales to credit risk, in line with the arguments of Brunnermeier and Oehmke (2013). Our aim here is to shed light on the existence of such potential relationship, without adopting a causality approach, an issue which is addressed in the next section. To that end, we estimate several regressions to shed light on the relationship of short positions (relative to the share capital) to the log-return of CDS for the three groups of firms considered (medium-sized banks, large banks and non-financial companies) before the ban.⁵ The

4. The measure of risk contagion in Figure 1 corresponds to the net spillover effect between the sovereign and the bank's CDS spreads obtained on the basis of Diebold and Yilmaz's (2012) methodology, which is described in the Appendix.

5. We use the first difference in the logarithm of the CDS spread because the CDS spread series are integrated of order one, so as to avoid any bias in the results due to the dependence on the level of the CDS premium or the degree of risk.

baseline regression is estimated using daily data from March 2011 to December 2011 on the basis of the following Ordinary Least Squares (OLS) fixed-effects regression with the standard errors clustered by day and robust to heteroskedasticity:

$$\begin{aligned} \Delta \log(CDS_{i,t}) = & \alpha + \beta_1 Short_{i,t-1} \times LargeBank_i + \beta_2 Short_{i,t-1} \times MedBank_i \\ & + \beta_3 Short_{i,t-1} \times NonFin_i + \beta_4 Short_{i,t-1} \times PreBan_t \times LargeBank_i \\ & + \beta_5 Short_{i,t-1} \times PreBan_t \times MedBank_i + \beta_6 Short_{i,t-1} \times PreBan_t \times NonFin_i \\ & + \beta_7 PreBan_t + \beta_8 MedBank_i + \beta_9 LargeBank_i + \zeta Controls_{i,t} + \varepsilon_{i,t} \end{aligned} \quad (1)$$

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 $t-1$.⁶ The dummy variables *LargeBank*, *MedBank*, and *NonFin* denote the type of firm: large banks, medium-sized banks and non-financial firms, respectively. 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 and another that takes the value of one before the ban and zero otherwise (*PreBan*). We also use the individual dummies relating 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, including the following: the percentage changes in the CDS liquidity (bid-ask spread), the lagged stock returns, a measure for the stock return volatility lagged one day, 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. We also take into account that some days before the ban, and with effect from 8 August 2011 onwards, the European Central Bank (ECB) restated its readiness 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 sovereign risk and, indirectly, on private sector risk too. Hence, we consider the ECB's weekly purchases of government bonds (in billions of euro) in the secondary market from 8 August onwards.

The results are reported in column 1 of Table 1. From these results we can single out that, before the ban, the ratio of short positions relative to share capital in medium-sized banks moved in tandem with their average CDS. In particular, a change equal to the average standard deviation of the ratio of short-positions in medium-sized banks relative to their share capital during the pre-ban period (0.31%) goes hand in hand with an average daily percentage increase of 0.21% in their CDS premiums, which amounts to 33.6% of the average daily percentage change of the CDS premiums of medium-sized banks during the pre-ban period. By way of contrast, the pre-ban volume of short positions relative to the capital of the other two groups of firms does not seem to be significantly related to changes in their CDS. The asymmetric intensity in the ratio of short positions and the default risk

6. We use the information that refers 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) – on the ratio of short positions and find no evidence of unit root in the panel data. Note also that the short-positions variables are defined as a ratio and, therefore, their values are bounded between zero and one.

indicators of medium-sized banks relative to the two largest banks could be related to the greater exposure of the former group to domestic macroeconomic risk, this being due to their lesser geographical diversification and their greater direct exposure to sovereign risk. In particular, according to the European Banking Authority (EBA), the average ratio of medium-sized banks' gross direct exposures to the Spanish sovereign relative to their total assets stood at 8.5% by December 2010 while the relative exposure of big banks was 6.9%. The average percentage of loans granted by medium-sized banks to domestic borrowers was 94.8%, which more than doubled the corresponding ratio for the two big banks (45.6%).⁷

Finally, we observe that none of the interactions between the short positions using the whole sample and the three groups of firms have a significant coefficient, indicating that there were no significant relationships between the levels of short positions and credit risk for any of the three types of firms apart from those found in the pre-ban period.⁸

The relationship of the remaining control variables to the dependent variable is as expected. In particular, percentage changes in the sovereign CDS spreads are positive and significantly related to the average changes in the CDS of listed companies. As expected, increases in the CDS bid-ask spread, which signal lower liquidity of the CDS contracts, have a positive and significant relationship to the dependent variable. Conversely, increases in the average firm's stock returns of the previous trading session have a significant negative coefficient. 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. The negative and significant coefficient for the lagged changes in the CDS spreads, whose scale is low in absolute terms, reveals the persistent nature of deviations of this variable from its long-term average. The estimated coefficients for the two volatility variables, the global indicator of market volatility (VIX Index) and the firm-specific equity volatility (squared of stock returns) are not significant at 5%.

Finally, the ECB's purchases of government bonds in the secondary market have a negative relationship to CDS, as expected, but this relationship is not significant at any standard significance level. This lack of significance of ECB purchases of debt after 8 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, we find a non-significant effect of the ECB's announcement on the Spanish sovereign CDS in the weeks after 8 August.⁹ More to the contrary, as shown in Figure 2, the Spanish sovereign risk surged again a few days after a sharp fall following the announcement.

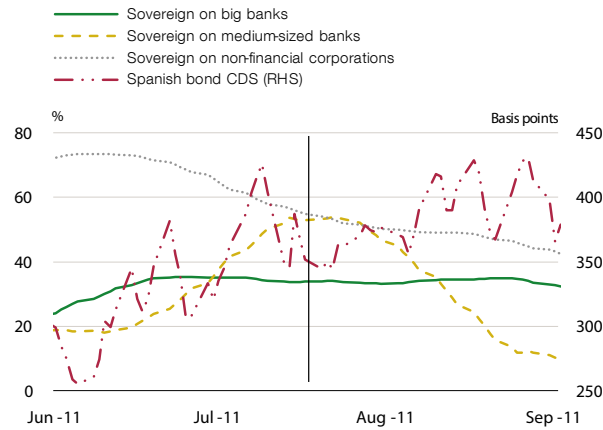
Column 2 contains the estimated coefficients of short positions for the pre-ban period (March 2011 – August 2011). These estimates confirm that the ratio of short positions to share capital in medium-sized banks was significantly related to their average CDS before the ban, whereas the closure of such positions afterwards would have not had a significant coefficient. The previous significant pre-ban link found in the medium-sized banks is not present either in the non-financial corporations or in the largest banks. Again, we obtain a significant coefficient of the sovereign CDS premium.

7. These data are obtained from SNL DataSource.

8. As an alternative proxy for short-sale activity we also use the volume of stock lending relative to share capital adjusted by dividend payment dates and other calendar and find similar results.

9. We regress the first difference of the logarithm of the sovereign CDS spread on a dummy variable that takes the value of one for the month after 8 August and zero for the month before that date, and we find a non-significant effect of the dummy coefficient at any standard level of significance.

Figure 2: Sovereign risk contagion in Spain²



(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, and 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.

To enlarge the number of firms in our analysis, which is limited in the baseline due to lack of CDS references for many listed firms, we next exploit data on default probabilities of the listed companies, obtained from StarMine.¹⁰ 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 almost all trading volume. We use the first difference of the logarithm of model-based default probabilities as the dependent variable to be consistent with the previous baseline regression analysis, and also because the series for default probabilities are not stationary. The model and the estimation methodology are similar to those summarised in equation (1). Given that the dependent variable is the change in the log of the default probability, we use the lagged change in the logarithm of default probabilities as a control. However, we do not employ information referring to the stock returns and their volatility in our regression because the default probabilities reported by StarMine are based on a model that exploits the equity market value, its volatility and liability structure as inputs. As the ban also affected financial firms other than banks (insurance and financial services firms) that are now included in the sample, we also control for the short positions on them (lagged one period).

The results reported in column 3 of Table 1 confirm that, before the ban, short positions in medium-sized banks had a positive and significant relationship to the changes in default probabilities. A change equal to the average standard deviation in the ratio of short-positions of medium-sized banks relative to their share capital during the pre-ban period (0.31%) goes hand in hand with an average daily percentage increase of 0.24% in their default probabilities for this type of firm during that period. This increase represents 36.6% of the average daily percentage change in the CDS premiums of medium-sized banks during the pre-ban period. As in the baseline, we do not find any significant effect at the 5% significance level for the other three groups of firms considered (big banks, other financials and non-financials). We find positive but not significant coefficients for the changes in the VIX Index and the Spanish sovereign CDS. Interestingly, if we restrict the analysis to firms for which we have CDS information or to the 50 largest firms, the effect

¹⁰Based 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.

of the Spanish sovereign CDS is positive and significant. This suggests that default probabilities were also positively affected by global uncertainty but the effect of the Spanish sovereign CDS was only significant for the largest firms.

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

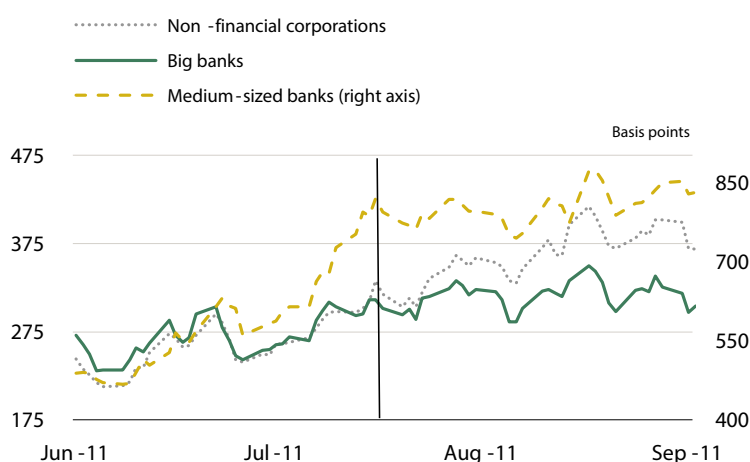
| | (1) | (2) | (3) |
|--|--------------------|-------------------|--------------------|
| Short positions (t-1), non-financials | 0.179 (0.34) | | 0.054 (0.09) |
| Short positions (t-1), medium-sized banks | 0.446 (0.27) | | 0.487 (0.30) |
| Short positions (t-1), large banks | -0.089 (2.30) | | -3.053 (4.06) |
| Short positions (t-1), other financials | | | 0.417 (0.52) |
| Short positions (t-1), non-financials before ban | -0.022 (0.13) | -1.219 (1.69) | 0.018 (0.06) |
| Short positions (t-1), medium-sized banks before ban | 0.237* (0.10) | 0.900* (0.42) | 0.291* (0.13) |
| Short positions (t-1), large banks before ban | -7.552 (5.87) | -9.130 (5.62) | 4.161 (13.27) |
| Short positions (t-1), other financials before ban | | | 0.437 (0.22) |
| Dummy before ban | 0.008 (0.00) | | 0.003 (0.00) |
| Dummy medium-sized banks | -0.024 (0.01) | -0.031 (0.02) | -0.025 (0.02) |
| Dummy big banks | 0.001 (0.00) | 0.003 (0.00) | 0.003 (0.01) |
| $\Delta \log(\text{CDS}(t-1))$ | -0.108** (0.04) | -0.073 (0.06) | -0.122** (0.02) |
| Stock returns (t-1) | -0.151* (0.07) | -0.215 (0.13) | |
| Squared of stock returns (t-1) | 0.047 (0.08) | 0.018 (0.16) | |
| $\Delta \log(\text{VIX}(t-1))$ | 0.033 (0.02) | 0.000 (0.03) | 0.037 (0.02) |
| $\Delta \log(\text{Spanish sovereign CDS}(t-1))$ | 0.238** (0.04) | 0.212** (0.06) | 0.036 (0.03) |
| ECB Bond purchases since 8 August 2011 | 0.001 (0.00) | | 0.000 (0.00) |
| $\Delta \log(\text{CDS bid-ask spread})$ | 0.008* (0.00) | 0.008 (0.01) | |
| Constant | -0.006 (0.00) | 0.000 (0.00) | -0.006 (0.01) |
| Fixed effects | YES | YES | YES |
| Observations | 3,767 | 1,790 | 23,093 |
| Number of companies | 17 | 17 | 105 |
| Adj. R-squared | 0.11 | 0.07 | 0.03 |

This table reports the effects of the short positions in Spanish listed companies on the changes in their default risk by firm type (medium-sized banks, large banks, and non-financials) before the ban. The coefficients are estimated using daily data from March 2011 to December 2011 on the basis of an OLS fixed effects regression with the standard errors clustered by day and robust to heteroskedasticity. The first column reports the effect of the short positions relative to capital and their incremental effect before the ban for each firm group on the log-change of CDS prices. The second column reports the results for a sample restricted to the pre-ban period (March 2011 – August 2011). Column 3 reports the results obtained when the measure of credit risk is the first difference of the logarithm of default probabilities. All the variables are defined as 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.

5 The effect of the ban on credit risk

After several weeks of sustained increases, the average spreads of the CDS of the medium-sized banks eased appreciably following the ban (see Figure 3). In effect, the CDS of both sets of financial institutions stabilised in the first few weeks after the ban, while those of non-financial corporations went on accelerating for some weeks more. In order to test statistically the hypothesis about the stabilising effect of the ban, we first perform several estimations aimed at identifying a potentially differential impact of this policy measure on the CDS spreads across the three firm groups (section 5.1). We then examine potential substitution effects in short positions across financial and non-financial firms triggered by the ban (section 5.2).

Figure 3: CDS spreads of Spanish listed companies



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

5.1 The impact of the ban on the credit risk of financial firms

One important first question concerns the length of the sample period over which the short-selling ban may exert a first-order effect on asset prices. In order to limit the possibility that observations far away from the date of the ban might cloud its effects, we assess the impact of the ban using information over a two-month period, centred on the date of its inception. Specifically, we regress the first difference of the logarithm of the CDS spread on several dummy variables relating to the two bank groups and on the date of the ban:

$$\begin{aligned} \Delta \log(CDS_{i,t}) = & \alpha + \beta_1 MedBank_i + \beta_2 LargeBank_i + \beta_3 After_t + \\ & + \beta_4 After_t * MedBank_i + \beta_5 After_t * LargeBank_i + \zeta_t + \varepsilon_{i,t} \end{aligned} \quad (2)$$

where *MedBank* (*LargeBank*) is equal to 1 if firm *i* is a medium-sized (large) bank and zero otherwise. The dummy *After* takes value 1 during the post-ban period (one month) for all the firms in the sample while the interaction of the dummy *After* and the two bank groups takes value 1 for the corresponding group after the ban.

The estimates are obtained using OLS day fixed-effects regression with the standard errors robust to heteroskedasticity and clustered by firm. Column 1 of Table 2 reports the coefficients of this estimation. The time fixed effects are included to account for a potential aggregate risk factor that may affect the CDS price changes of all the firms in the study. Column 2 contains the results when we also consider fixed firm-level effects to control for unobserved heterogeneity. As these regressions include dummy variables related to the firm type, we ease the computational task of estimating firm fixed effects by demeaning all the firm-specific variables at the firm level and then performing a regression similar to that reported in equation (2).

Table 2: Effects of the ban on CDS spreads

| | (1) | (2) | (3) |
|------------------------------------|--------------------|--------------------|--------------------|
| Dummy medium-sized banks after ban | -0.015** (0.01) | -0.015** (0.01) | -0.017** (0.01) |
| Dummy big banks after ban | -0.003 (0.00) | -0.003 (0.00) | 0.001 (0.00) |
| Dummy other financials after bank | | | 0.005 (0.01) |
| Dummy after ban | -0.079** (0.01) | -0.079** (0.01) | -0.021* (0.01) |
| Dummy medium-sized banks | 0.008 (0.01) | 0.008** (0.00) | 0.010** (0.00) |
| Dummy big banks | -0.001 (0.00) | 0.003 (0.00) | 0.002 (0.00) |
| Other financials | | | -0.002 (0.00) |
| Intercept | 0.032** (0.01) | 0.021* (0.01) | 0.052** (0.01) |
| Fixed effects | No | Yes | Yes |
| Time effects | Yes | Yes | Yes |
| Number of observations | 765 | 765 | 4,725 |
| Adjusted R-squared | 0.57 | 0.57 | 0.15 |

This table reports the estimation results of the effect of the short-sale ban on the changes in the logarithm of default risk indicators of the groups of firms affected by the ban relative to the effect of the firms unaffected by the ban (control group). We classify the firms affected by the ban in different groups depending on their size and the type of financial institution (big banks, medium-sized banks, and other financials). Specifically, we regress the first difference of the logarithm of the default risk indicator on (i) dummies for the types of firms affected by the ban that take value 1, if the firm belongs to such group, and zero otherwise, (ii) a dummy that is equal to 1 during the post-ban period for all the firms, and (iii) the interaction of the post-ban dummy with the type of financial institution. The coefficients for the last two dummies will indicate the differential effect of the ban on the type of financial institution relative to the non-financial firms. The estimates are obtained using OLS day and firm fixed-effects regression with the standard errors robust to heteroskedasticity and clustered at the firm level. The indicator of default risk in the first and second columns is the CDS price. The only difference between these two columns is that the first one does not include firm fixed effects while the second one does. The indicator of default risk in the third column is the default probability. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors are reported between brackets.

The first hypothesis to test is whether the ban contributed to a significant moderation in the CDS spreads for the two groups of banks relative to that of non-financial firms (Column 1). The coefficient β_4 (β_5) provides the differential effect of the ban on medium-sized (large) banks vis-à-vis non-financials. The main result is that, after the ban, there was a significant decrease (at 1% significance level) in the credit risk of medium-sized banks with respect to the non-financial firms. Specifically, the average decrease of the daily percentage change in the CDS spreads of medium-sized banks after the inception of the ban was around 1.5 percentage points (pp) with respect to the daily percentage changes of the non-financial firms. For the largest banks we observe a lower fall in the percentage change in the CDS spreads (0.3 pp) that is not statistically

different from zero. The results reported in column 2 confirm the differential significant effect observed on medium-sized banks after the inclusion of fixed firm-level effects.

To check the robustness of the previous results to changes in the credit risk metric and in the size of the firms' sample, we next perform a similar regression in which we use the log of changes in the default probabilities corresponding to 105 firms estimated by StarMine. As before, we consider both day- and firm-fixed effects. The results of this estimation (column 3 of Table 2) are similar to those obtained using CDS prices as the default risk measure.

The previous results would be consistent with the existence of a causality relationship running from short positions to CDS spreads during the pre-ban period, in the spirit of the theoretical arguments of Brunnermeier and Oehmke (2013), Liu (2014) and Ventel (2011). Subsequently, the ban would have significantly eased this short-sale risk-channel. To gain further insight into this potential channel, we next extend the estimation in equation (2) with firm fixed-effects after classifying the set of medium-sized banks into two groups, depending on the average level of short positions on them before the ban (above and below the median ratio of short sales to capital for the entire group of medium-sized banks). Thus, we consider four different groups of firms: the control group (non-financial) and three different treatment groups (big-banks, medium-sized banks more exposed to short-sale activity and medium-sized banks less exposed). In order to better isolate the potential short-sale risk channel, we also control for individual and aggregate risk. The former is proxied by each bank's leverage, measured as the ratio of total debt relative to the firm market value.¹¹ As for aggregate risk, we consider the impact of a rise in sovereign risk on individual banks' risk indicators proxied by means of the spillover effect between sovereign and banks' CDS spreads described earlier. In addition, we control for the one-day lagged percentage change of CDS prices and the one-day lagged square stock return.

$$\begin{aligned} \Delta \log(CDS_{i,t}) = & \alpha + \beta_1 MedBank_i + \beta_2 LargeBank_i + \beta_3 After_t + \\ & + \beta_4 After_t * MedBank_i + \beta_5 After_t * LargeBank_i + \zeta_t + \delta RiskFactors_{i,t} + \varepsilon_{i,t} \end{aligned} \quad (3)$$

The estimation is conducted on the basis of an OLS day and firm fixed-effects regression with the standard errors robust to heteroskedasticity and clustered at the firm level. The estimation is implemented replacing dependent and independent firm-specific variables by their deviations from the respective firm-level average and including daily fixed effects in the regression. Results are shown in column 1 of Table 3.

We find that the ban was in fact effective (in a significant way) in reducing the CDS only in those medium-sized banks that had relatively high levels of short positions. This result suggests that short positions could be causing risk by themselves through, for example, the self-fulfilling mechanism highlighted by Brunnermeier and Oehmke (2013). A *t-statistic* test comparing the coefficients obtained for the two types of medium-sized banks depending on the level of short positions indicates, as expected in view of the previous result, that the decrease in the CDS prices of medium-sized banks whose stocks were more exposed to short-sales activity is significantly larger than that obtained for the other group of medium-sized banks. Regarding the risk variables employed as controls, the only one with a positive and significant effect is the sovereign risk spillover. Neither leverage nor the square stock return has a significant effect on the dependent variable.

¹¹. This definition of leverage is motivated by the need to have daily variation in the proxy for leverage to be employed as an additional regressor. We employ other proxies for leverage with a lower variation and find similar results.

Table 3: Effects of the ban on CDS spreads using different groups of medium-sized banks

| | (1) | (2) | (3) |
|---|--------------------|--------------------|--------------------|
| Dummy medium-sized banks (above median) after the ban | -0.022** (0.00) | -0.014* (0.01) | -0.016* (0.01) |
| Dummy medium-sized banks (below median) after the ban | -0.009 (0.01) | -0.016* (0.01) | -0.012* (0.01) |
| Dummy big banks after ban | -0.005 (0.00) | -0.004 (0.00) | -0.004 (0.00) |
| Dummy after ban | -0.045** (0.01) | -0.081** (0.02) | -0.044** (0.01) |
| Dummy medium-sized banks (above median) | 0.011** (0.00) | 0.007 (0.00) | 0.008* (0.00) |
| Dummy medium-sized banks (below median) | 0.005 (0.00) | 0.009** (0.00) | 0.006* (0.00) |
| Dummy big banks | 0.004* (0.00) | 0.003* (0.00) | 0.003 (0.00) |
| Intercept | 0.014* (0.01) | 0.022* (0.01) | 0.013* (0.01) |
| Fixed effects | Yes | Yes | Yes |
| Time effects | Yes | Yes | Yes |
| Idiosyncratic and aggregate risk controls | Yes | Yes | Yes |
| Number of observations | 765 | 765 | 765 |
| Adjusted R-squared | 0.57 | 0.57 | 0.57 |
| P-value: Coef (Dummy med-sized banks - above median - after ban) ≠ Coef (Dummy med-sized banks - below median - after ban) | 0.025 | 0.812 | 0.581 |

This table reports the estimation results of the effect of the short-sale ban on the changes in the logarithm of CDS prices of the groups of firms affected by the ban (big and medium-sized banks) relative to the effect of the firms unaffected by the ban (control group). We classify the medium-sized banks in different groups depending on their exposure to short-sale activity (column 1: above vs. below the median ratio of short sales to capital of medium sized-banks), their leverage (column 2: above vs. below the median ratio of debt to market value of medium sized-banks), and their exposure to sovereign risk (column 3: above vs. below the median net spillover effect between the sovereign and the group of medium-sized banks). Specifically, we regress the first difference of the logarithm of the CDS prices on (i) three dummies for the three types of firms affected by the ban that take value 1, if the firm belongs to such group, and zero otherwise, (ii) a dummy that is equal to 1 during the post-ban period for all the firms, (iii) the interaction of the post-ban dummy with the three bank groups, and (iv) idiosyncratic and global risk factors. The coefficients for the last three dummies will indicate the differential effect of the ban on the three bank groups relative to the non-financial firms. The estimates are obtained using OLS fixed- and time-effects regression with the standard errors robust to heteroskedasticity and clustered at firm level. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors are reported between brackets. The last row contains the *p-value* for a simple test comparing the coefficients for both types of medium-sized banks in each column. The *p-value* is for the null hypothesis indicating that the difference between the coefficients for the two types of medium-sized banks is not different from zero.

Finally, we analyse whether the medium-sized banks' exposures to idiosyncratic and aggregate risk factors also lead to an asymmetric effect on their CDS. In this connection, we classify medium-sized banks into two alternative categories attending to the level of leverage factor (above versus below the median ratio of debt to market value in medium-sized banks) and their exposure to sovereign risk (above versus below the median net spillover effect between the sovereign and the medium-sized banks). The estimated coefficients for the two alternative classifications are reported in columns 2 and 3 of Table 3, respectively. We find a similar significant differential effect on the changes in the CDS prices of the two types of medium-sized banks irrespective of whether the separation is based on the leverage or the exposure to sovereign risk. In fact, a simple *t-statistic* reveals that the difference across these two dimensions of risk is not significantly different from zero at any standard significance level.

In sum, we conclude that the ban led to a decrease in the CDS prices of medium-sized banks but this effect was not influenced by their leverage ratio or their exposure to sovereign risk of medium-sized banks. Only the level of short positions constitutes a differential factor within the category of medium-sized banks, as suggested by the significant relationship between short positions and CDS spreads for this group of banks before the ban.

5.2 The effect of the ban on non-financial firms

We next analyse the potential (indirect) effects of the ban on short positions in non-financial firms. To this end, we perform some additional exercises using information on short positions two months around the inception date. First, we regress the ratio of short positions relative to share capital 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. We find a positive and significant coefficient (at 1% significance level) indicating an increase of 0.04% in the ratio of short positions after the ban. This increase in short positions represents 9% of the average ratio of short positions over capital of non-financial firms during the month after 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 last sub-sample, the previous positive differential effect rises to 0.11% (i.e. 9.2% of the average ratio of short positions over capital for these firms during the month after the ban).

The previous result suggests 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 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 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 around the time of the ban. To compute the contagion indicator we apply Diebold and Yilmaz's (2012) methodology to the firm's stock returns and the sovereign CDS returns. The corresponding regression reveals that the variation in short positions is larger (more than double) in the firms that were more exposed to the sovereign risk. Indeed, the dummy for the firms below the median level of contagion is not significant, while that for the firms above the median is significant at any standard level of significance. In particular, the average increase after the ban in the levels of short positions in the sub – set of non – financial companies that are more heavily exposed to the sovereign risk is 0.16%. This increase represents 10% of the average ratio of short positions over capital for this sub-set of firms during the month after the ban.

6 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 using the same 17 firms as in the previous sections.¹² As in Section 5, we exploit information from a two-month period centred on the inception date of the ban.

The panels of Figure 4 contain the proxies used to capture several market variables of interest. To test for the statistical significance of the patterns shown therein, we perform several regressions on the basis of the following equation changing the dependent variable accordingly:

$$\begin{aligned} Dep Var_{i,t} = & \alpha + \beta_1 MedBank_i + \beta_2 LargeBank_i + \beta_3 After_t + \\ & + \beta_4 After_t * MedBank_i + \beta_5 After_t * LargeBank_i + \zeta_t + \mu_i + \varepsilon_{i,t} \end{aligned} \quad (4)$$

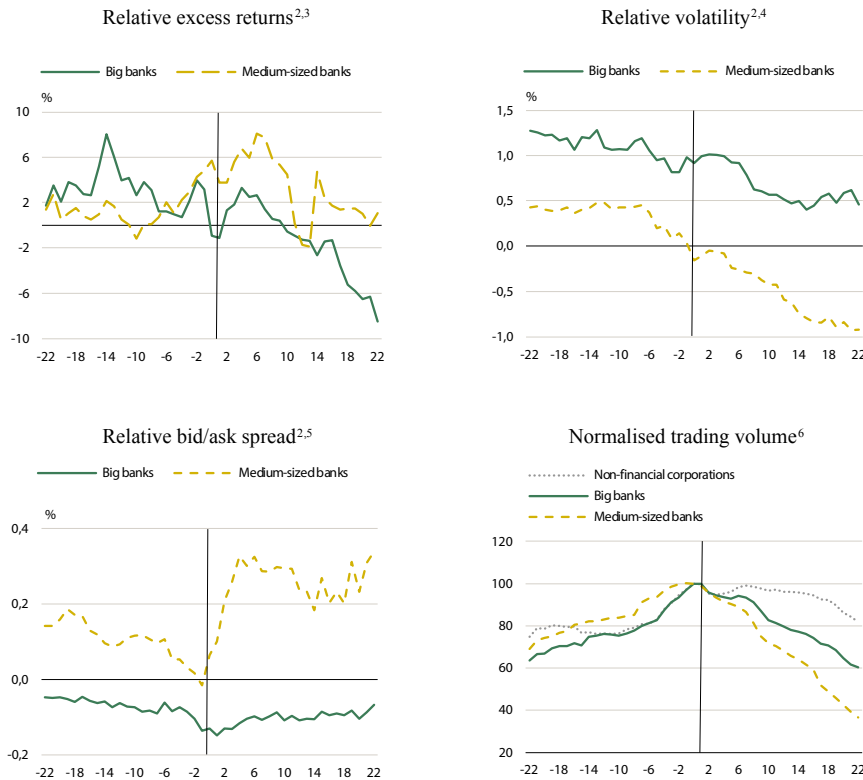
In particular, as dependent variables we employ proxies for volatility, liquidity, volume and returns. The estimates are obtained using OLS with robust standard errors to heteroskedasticity and including time (day) and firm fixed-effects. The coefficients of this estimation are shown in Table 4. The null hypotheses to test are whether the effect of the ban on the (i) returns, (ii) volatility, (iii) relative bid-ask spreads, and (iv) trading volume of financial firms was significantly higher than that observed for non-financial firms.

Returns. The first apparent effect of the ban is that financial corporations' shares outperformed those of non-financial firms, 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 some cases the positive effect was temporary and had faded considerably around one week after the inception of the ban. Furthermore, as revealed by the estimates contained in column 1 of Table 4, the hypothesis that the ban had a significantly stronger effect on the banks, irrespective of their size, than on the non-financial firms can be rejected at any standard 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 such constraints in comparison with other stocks unaffected by them.

Drawing on the theories that attribute a potential destabilising role to short sales based on the possibility of multiple equilibria, as in Brunnermeier and Pedersen (2005), Venter (2011), and Brunnermeier and Oehmke (2013); the lack of a significant differential effect of the ban on the prices of financial stocks with respect to non-financial ones could reflect the fact that the ban is effective in removing self-fulfilled "low valuation" equilibria and, hence, also in helping the market to clear at "normal" (fundamentals-based) equilibrium. Yet, once in the latter equilibrium, the ban does not, by itself, exert any overvaluation effect on the stocks affected vis-à-vis other stocks.

¹² As an additional benchmark for assessing the impact of the ban on the previous variables, we also take the group of companies included in the Ibex 35 that were not subject to the ban and find similar results to the ones presented in this section.

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



- (1) The X-axis is a time scale in which 0 marks the starting date of the ban.
- (2) Relative to the set of non-financial companies with CDS that were not subject to the ban.
- (3) The relative excess returns of each group on each date is the average of the cumulative spread to that date between the daily returns of component corporations and the average returns of the control group.
- (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 difference between the average relative bid/ask spreads of the share prices of component corporations and the average relative bid/ask spreads of the control group (non-financials). 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 normalised to 100 at the time of the ban.

Volatility. The disparity found in the preceding section in terms of the unequal impact of the ban on the indicators of default risk between large and medium-sized banks carries over in terms of their volatility. The top right panel of Figure 4 illustrates that while the average relative returns volatility of the two largest banks was apparently unaffected, that of medium-sized banks abated considerably. Thus, the ban had a strong moderating impact on the price fluctuations of medium-sized banks' shares, whose volatility readings dropped below those of companies not covered by the ban. Conversely, the relative volatility of the largest-cap banks showed little variation and, in fact, increased two weeks after the ban. The estimates in column 2 of Table 4 show the coefficients for the dummy variables of the regression in which volatility is measured as the absolute value of the stock returns for an easier interpretation of such parameters.¹³ We observe that following the ban there was a significant decrease in the volatility of the returns of the medium-sized banks with respect to that of the non-financial firms. This significant decrease was not perceived in the largest banks. The average decrease of the daily volatility of medium-sized banks following the ban was around 0.64 pp with respect to that of the non-financials group. In the case of the medium-sized banks, average volatility during the month before the ban was 1.93% and, therefore, the decrease following the ban was around 33% of the pre-ban level. This decrease was significantly larger than that obtained for the large banks.

¹³ Alternatively, we define the daily volatility from the coefficient of variation of the intraday stock prices and obtain similar results to those obtained for the absolute value of the stock returns.

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

| | Return | Volatility | Volume | Bid-ask |
|------------------------------------|-------------------|---------------------|---------------------|---------------------|
| Dummy medium-sized banks after ban | -0.0031 (0.00) | -0.0064** (0.00) | -0.6260** (0.10) | 0.0015** (0.00) |
| Dummy big banks after ban | -0.0044 (0.00) | -0.0024 (0.00) | -0.1967** (0.03) | -0.0001 (0.00) |
| Dummy after ban | -0.0043 (0.00) | 0.0258** (0.00) | 0.3343** (0.12) | -0.0003 (0.00) |
| Dummy medium-sized banks | 0.0072 (-0.01) | -0.0085* (0.00) | 0.5270* (0.27) | 0.0015** (0.00) |
| Dummy big banks | 0.0006 (0.00) | 0.0047 (0.00) | -0.1568** (0.05) | -0.0004** (0.00) |
| Intercept | 0.0043 (0.00) | 0.0086* (0.00) | 0.5629** (0.08) | 0.0012** (0.00) |
| Fixed effects | YES | YES | YES | YES |
| Time effects | YES | YES | YES | YES |
| Number of observations | 758 | 758 | 759 | 757 |
| Adjusted R-squared | 0.08 | 0.49 | 0.36 | 0.46 |

This table reports the estimation results of the effect of the short-sale ban on stocks' excess returns, volatility, liquidity, and volume of the two groups of firms affected by the ban (medium-sized and large banks) relative to a control group of non-financial firms that were not affected by the ban. Specifically, we regress the first difference of the four previous measures on (i) two dummies for the two bank groups (medium and large banks) that take value 1 if the firm belongs to such group and zero otherwise, (ii) a dummy that is equal to 1 during the post-ban period for all the firms, and (iii) the interaction of the post-ban dummy with the two bank groups. The coefficients for the last two dummies will indicate the differential effect of the ban on the two bank groups relative to the control group. The estimates are obtained using OLS fixed- and time-effects regression with the standard errors robust to heteroskedasticity. The dependent variable in column 1 is the firm's daily stock return in excess of the market. 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 absolute value 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 1 at the time of the inception of the ban for all the firms in the sample. The null hypothesis to test is whether the effect of the ban on the (i) returns, (ii) volatility, (iii) relative bid-ask spreads and (iv) volume of financial firms was significantly higher than that observed for non-financial firms that were not affected by the ban. ** and * indicate whether the coefficients are significant at a significance level of 1% and 5%, respectively. Standard errors are reported between brackets.

Liquidity. One of the most visible consequences of the ban was that the liquidity conditions of medium-sized banks deteriorated sharply. Specifically, the bid/ask spreads of firms in this group widened persistently after the onset of the ban compared with those of non-financial firms (see lower left panel of Figure 4). Concretely, the ban led to a significant average increase of around 0.15% of the relative bid-ask spread of the medium-sized banks with respect to the relative bid-ask spread of the non-financials. Nevertheless, we do not observe a differential effect significantly different from zero in the case of the largest banks. In the case of the medium-sized banks, the average relative bid-ask spread was 0.34% at the beginning of the sample period (June 2011) and, therefore, the increase following the ban was around 50% of the level prevailing well before. Compared to the differential effect on the big banks, the decrease in the relative bid-ask spread of medium-sized banks is significantly larger.

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 more than 60% 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). The maximum drop in trading volumes from levels immediately before the ban was over 40% in the case of the banking majors and 10% among non-financial corporations. The estimates in column 4 of Table 4 show that following the ban there was a significant decrease in the stock trading volume of the medium-sized and large banks with respect to the volume of the non-financial firms. The average differential decrease of medium-

sized and large banks' trading volumes with respect to that of non-banned firms was almost 63% and 20%, respectively.

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 commencement of the ban. To estimate this effect, we calculate the speed of share price adjustment as the average difference (as a 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 three months before and after the ban.¹⁴ Thus, a positive value (higher first-order correlations after the ban) 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 10%. In the case of the largest banks and non-financial corporations, the apparent loss of price information efficiency was considerably less severe. The percentage change in the speed of share price adjustment was 0.1% and 3% for the large banks and the non-financial corporations, respectively.

¹⁴ We extend our sample in order to obtain a larger number of observations to procure the correlations for computing the price discovery metrics. Using two months instead of three, the percentage change in the speed of share price adjustment is 13.6%, -1.1%, and 3.2% for the medium-sized banks, large banks, and non-financials, respectively.

7 Conclusions

This paper analyses the main effects of the short-sale 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 grew in parallel to the CDS spreads of medium-sized banks before the ban, whose relatively low degree of geographical diversification and higher perceived implicit and explicit reliance on State aid rendered them more vulnerable to surges in domestic macroeconomic and sovereign risk. Subsequently, the ban helped stabilise the risk indicators of medium-sized banks more exposed to short-sellers' activity, even while sovereign credit risk headed higher. Specifically, we find that the volume of short positions accumulated before the ban is a significant explanatory factor of the negative effect of the ban on CDS. This suggests the existence of an underlying short-sale risk-channel, through which increases in short positions cause *per se* a rise in risk indicators. By way of contrast, we failed to find evidence of a significant link between short sales and credit risk for the largest banks and the non-financial firms before the ban, both being arguably less exposed 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.

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 sharper than that of other stocks. This latter 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 caused significant damage to the liquidity, trading volume and price discovery of their stocks.

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