JEL classification: G14; G18
Keywords: short-selling ban, credit default swap, information flow

Did Investors Seeking Short Exposure Move to the CDS Market after the 2011 Short-Sale Bans in European Financial Stocks?*

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Abstract

This paper addresses the effects that the August 2011 ban on covered short selling has had on the trading behavior of credit default swap (CDS) contracts. Overall, the results obtained suggest that CDS protection buying is not viewed by investors as a viable substitute for taking short interest in stocks. An initial analysis demonstrates that the CDS open interest of firms included in the ban lists actually declined after the ban. There is also no evidence of price pressure on CDSs written on firms on the ban lists. Surprisingly, CDSs on European financial firms and, in particular, on firms subject to the ban saw a decline in volatility after the ban. CDS bid-ask spreads evolved positively after the ban, but that increase was economically modest. Finally, there is evidence that the ability of CDS spreads to predict stock prices fell with the ban, which is inconsistent with the migration of informed traders to the CDS markets.

1. Introduction

In the aftermath of the Great Recession, European banks witnessed an increase of non-performing loans and impaired assets, which negatively affected their earnings expectations and profitability. Not yet recovered from that event, the banking sector found its resilience again challenged in 2010 with the onset of a new crisis affecting the valuation and risk of peripheral European government bonds. Indeed, rumors of the unsustainability of peripheral European countries' debt severely affected the banking sector, which regarded these securities as risk-free assets. These concerns applied not only to banks with significant exposure to that debt, but also to other banks with strong financial links to the former built on derivative and interbank markets.

The political discussion of a possible restructuring of Greek sovereign debt fueled the uncertainty of the banking sector during the second and third quarters of 2011. That period was characterized by repeated ratings downgrades, widening funding spreads, falling stock prices and increasing volatility in the banking sector, affecting primarily financial institutions with strong connections to peripheral countries' debt. As a consequence, the financial market authorities of several European countries temporarily banned covered short-selling activity on financial stocks (banks and insurance companies) on 11 August 2011. This public intervention was intended to reduce volatility and concerns about possible financial panic and a downward spiral in prices. The financial market authorities of France, Italy, Belgium and Spain

* The author is pleased to acknowledge financial support from Fundação para a Ciência e a Tecnologia and FEDER/COMPETE (grant UID/ECO/04007/2013) and FEDER/ COMPETE (grant POCI-01-0145-FEDER-007659). The author would also like to thank the editor and two anonymous referees for their valuable comments and suggestions.

officially justified these bans as having the purpose of restricting the benefits obtained by spreading false rumors or misleading information.[1] Although these measures were expected to be in place for only 15 days (with the exception of Belgium, which announced that the ban would remain in effect indefinitely), they remained in effect until February 2012.[2]

This paper addresses the effect of this ban on the behavior of credit default swap markets, namely on open interest, prices, volatility, liquidity and price discovery. The underlying rationale is that even though short selling was prohibited, informed investors could still profit from their superior knowledge through the purchase of credit protection. According to structural models of corporate capital structure (e.g. Merton, 1974), the price of a firm's debt and equity instruments and, consequently, their returns depend on the same company-specific information. In effect, these models price these instruments as contingent claims on a firm's assets, so that in the absence of any frictions, debt and stock markets should be perfectly integrated.[3] This no-arbitrage pricing relationship between equity prices and credit spreads also applies to equity prices and CDS spreads, given the close relationship between credit spreads and CDS premiums (Duffie, 1999). On the one hand, good (bad) news about a firm's fundamentals should translate into higher (lower) stock prices. Conversely, good (bad) news about a firm's fundamentals should drive down (up) the probability of default and, therefore, the CDS spreads. For that reason, CDS spreads and stock prices should move in opposite directions in the absence of capital structure changes or asset substitution.[4]

Investors holding private information about a firm may choose to trade in the stock market, the bond market or the CDS market. In the model developed by Easley *et al.* (1998), the decision as to which market to invest in depends on the leverage of the financial instrument and the liquidity of the markets. It is clear that CDSs have greater leverage as compared to stocks and bonds. However, the liquidity of the CDS markets is lower than that of stock markets. In addition, it has also been shown that when end-clients are on the buy side, they may pay a significant mark-up premium[5] due to the existence of a reduced number of sellers in the CDS market (Foley-Fisher, 2010).

[1] According to the ESMA's statement on 11 August, 2011, "European financial markets have been very volatile over recent weeks. The developments have raised concerns for securities markets regulators across the European Union. [...] While short-selling can be a valid trading strategy, when used in combination with spreading false market rumors, this is clearly abusive. [...]"

[2] On 13 February, 2012, FSMA and AMF announced the termination of the ban. This measure was also announced on 15 February 2012, by the CNMV. Finally, on 24 February 2012, the CONSOB ban also expired.

[3] Schaefer and Strebulaev (2008) demonstrate that the sensitivities of corporate debt returns to the underlying equity and riskless debt are significant, which is consistent with the idea that equity price changes subsume the fundamental information related to default risk. Several studies, including Aunon-Nerin *et al.* (2002), Zhang *et al.* (2009), Das *et al.* (2009), Tang and Yan (2007) and Galil *et al.* (2014), also use stock returns as an explanatory variable of CDS returns to capture changes in the fundamentals of a firm. Acharya and Johnson (2007) posit that stock market prices should reflect all available public information about a firm's business and, therefore, about its credit quality. Narayan *et al.* (2014) find strong evidence of co-integration between log stock prices and log CDS spreads.

[4] Asset substitution may affect the volatility of firms' business.

Indeed, most studies have shown that information is more likely to flow from the stock market to the CDS market than in the opposite direction. Hilscher *et al.* (2015) find that informed traders are active primarily in the equity market rather than the CDS market, an option that is determined partially by the existence of higher transaction costs in the CDS market. Marsh and Wagner (2016) and Wang and Bhar (2014) also confirm that information flows more strongly from the equity market to the CDS market than the other way around. The existence of larger bid-ask spreads in CDS markets than in stock markets is also confirmed in the present study. The average CDS bid-ask spread of the sample equals 5.5%, whereas the average stock bid-ask spread equals 0.6%.

Without prejudice of that, in the presence of short sale constraints, informed traders (not owning stock) holding negative information about a firm may not have a way to express their views in the stock market. As the ban did not apply to credit protection buying, a natural question that arises is whether CDSs helped relax binding short-sale constraints by allowing informed investors to speculate in the CDS market in lieu of the stock market. Evidently, they will use the CDS market only to the extent that the value of their information is sufficient to cover liquidity costs and mark-up premiums of CDS sellers.

To put this into context, while buying CDSs, an investor is speculating on the bad performance of a firm in a way similar to what a short seller does so that, at least theoretically, CDS protection buying is a viable substitute for short interest on stocks. This paper aims to determine whether the existence of CDSs on firms for which short selling became prohibited opened an alternative channel for pessimistic investors or traders in possession of superior negative information, but not owning the given stock, to indirectly build short interest. That presupposes the migration of these investors to CDS markets. In doing so, this paper adds to the bulk of the literature that investigates the impact of short-selling restrictions on financial markets. While the effects of such restrictions on stock markets have been widely discussed in recent years, there is still scarce evidence of their effects on trading activity in related markets such as the CDS and option markets, particularly in Europe.

The migration of informed traders from stock to CDS markets in the aftermath of a short-selling ban may have several implications. First, it may induce changes in the price discovery process and, in particular, on cross-market information transmission between stocks and CDSs. Second, it may produce price pressure and order imbalance on CDSs written on firms subject to short-selling bans. Third, as more informed traders may be willing to participate in the CDS market, there is the expectation of rising CDS bid-ask spreads due to greater exposure of market makers to adverse selection risk. As a result, hedgers demanding credit insurance may face greater transaction costs in the wake of a stock short-selling ban. These potential

[5] Tang and Yan (2010) find that net buying interest, a measure of latent trade imbalance between consecutive trades, significantly affects CDS price changes, whereas net selling interest has only a moderate effect. Similarly, Gündüz *et al.* (2013) report that CDS traders use their market power to charge significantly higher premiums in transactions with buy-side investors. This means that if short-selling operations are allowed, they should be preferred in most cases to buying CDS protection, given that the latter entails higher liquidity costs and paying a large premium to the CDS seller.

implications are examined in detail in this study and may provide new insights and guidance to regulators on how to act in the future in the face of similar situations.

The consensus of the financial literature is that short-selling restrictions may produce effects on the prices, liquidity, volatility and price efficiency of stocks. From a theoretical perspective, Diamond and Verrecchia (1987) claim that if informed traders not owning the stock are prohibited from selling short, prices should take longer to adjust to new negative information, which in turn hinders price discovery. Miller (1977) predicts that short-sale constraints bring about overpricing due to the exclusion of the views of pessimistic investors not owning the stock. Harrison and Kreps (1978) and Duffie *et al.* (2002) also show that prices can be above fundamental values when short selling is constrained.

Bai *et al.* (2007) argue that short-sale constraints elevate the risk to uninformed market participants by excluding informed investors with negative information and lessening the informativeness of market prices. Hong and Stein (2003) relate short-selling restrictions with market crashes by showing that accumulated unrevealed negative information is unveiled at once in the form of a market crash. Brunnermeier and Oehmke (2014) develop a theoretical model predicting that predatory short selling can emerge in an equilibrium when a financial institution (i) is close to its capital constraint (the vulnerability region) or (ii) violates its capital constraint even in the absence of short selling (the constrained region). Their model provides a potential justification for temporary restrictions on short selling for vulnerable institutions.

In general, empirical evidence supports the hypothesis that overpricing is reduced when short-selling operations are allowed—e.g. Danielsen and Sorescu (2001), Jones and Lamont (2002), Cohen *et al.* (2007) and Harris *et al.* (2013)—and that short interest is driven by informed investors—e.g. Asquith *et al.* (2005), Desai *et al.* (2002), and Boehmer *et al.* (2008). However, Shkilko *et al.* (2012) document that short sales may also induce downward pressure on prices even in the absence of negative information. Several authors provide evidence consistent with the hypothesis that short selling constraints lessen price efficiency—e.g. Saffi and Sigurdsson (2011), Boehmer and Wu (2013), Beber and Pagano (2013) and Bris *et al.* (2007). Boehmer *et al.* (2013) find that stocks subject to the 2008 US ban faced a severe decline of market quality (larger spreads, higher price impacts and increased intraday volatility). Beber and Pagano (2013) report a disruption in the liquidity of the banned stocks, particularly small-cap ones. Kolasinksi *et al.* (2009, 2013) show that the negative association between the percentage volume of short sales and stock returns became stronger during the 2008 US ban.

So far, the investigation of the effects of such bans on related markets, such as the option market and, in particular, the CDS market, has received little attention in the financial literature. Blau and Wade (2013) report that when short-sale constraints become more binding, informed traders migrate from stock to option markets. In contrast, Battalio and Schultz (2011) find no evidence of migration of informed traders from stock to option markets during bans. The results of Grundy *et al.* (2012) suggest that short selling was not replaced by bearish put-option strategies after the 2008 US ban. Battalio and Schultz (2011) and Grundy *et al.* (2012) argue that substitution did not occur by virtue of a surge in the bid-ask spreads of derivatives on banned stocks. The wider spreads resulted from an attempt on the part of market makers to protect themselves from adverse selection risk and to keep their inventories unchanged. In addition, hedging long (short) calls (puts) became more difficult with the ban, dissuading non-registered market makers from providing liquidity.

Ni and Pan (2011) investigate information flows between the US equity, option, and CDS markets during the 2008 short-sale ban. They document an increment of cross-sectional predictability between stock and derivative prices. Their results suggest that during the ban the assimilation of negative information occurred first in derivative markets in accordance with the notion that the adjustment of equity prices to negative information became slower. Courtney (2010) shows that investors migrated to the CDS market to take short positions in banned stocks. As a consequence, banned firms experienced significant CDS price pressure that reverted following the ban, with those effects being more pronounced for those with greater pre-ban short interest.

In general, the results of this study do not support the hypothesis of a migration of informed traders from stock to CDS markets. An initial analysis shows that the net open interest—a measure of the willingness of traders to risk capital—of CDS contracts written on financial firms subject to the ban (treatment group) actually declined a week after short selling was prohibited. More importantly, the net open interest of those firms fell by a greater amount than did that of other European financial firms not subject to the ban (control group). Until the end of the ban, firms subject to the regulatory measure and control-group firms recorded similar patterns of open interest growth. These results contradict the idea that investors moved from stock to CDS markets to gain short exposure in financial stocks.

An analysis of the pattern of CDS spreads is also conducted in parallel. If informed investors suddenly moved from stock to CDS markets, a rise of CDS spreads due to greater short-term price pressure is expected. However, the spreads of CDSs written on firms that were subject to the prohibition actually fell in the five trading days after the ban announcement. A decline of CDS spreads is also observed for other financial firms not included in the ban lists, but the mean change difference of the treatment and control groups is negative and statistically significant. It is important to note that at least part of this downtrend in both groups is explained by country-specific factors. After excluding country-specific factors, the CDS spread changes of the groups in the five trading days after the ban announcement are not statistically different. Longer time horizons are also examined. However, in neither of the other time horizons is the hypothesis that the two groups recorded similar performance in CDS markets rejected.

To better understand the reason why migration did not occur, the pattern of CDS liquidity is investigated. If informed traders intend to migrate to the CDS market, dealers may respond with larger bid-ask spreads to prevent those traders from exploiting their private information in transactions and to ensure a positive profit in their activities. The results obtained corroborate this hypothesis; CDSs written on banned firms recorded higher bid-ask spread changes in comparison to other financial firms not included in the ban lists. A complementary analysis also shows that CDS volatility and kurtosis declined by a greater extent for firms included in the ban lists than for other firms, suggesting that the inventory risk of CDS market makers did not drive up bid-ask spreads. The skewness of CDS returns on firms included in the ban lists does not appear to have been affected by the ban.

A final investigation tackles the hypothesis of rising cross-predictability between CDS innovations and stock returns. Indeed, if informed investors migrated to the CDS market in the aftermath of the ban implementation, an increment of the predictability of stock returns based on past (positive) CDS innovations should occur. The principal findings of that analysis show that stock returns' cross-predictability existed prior to the short-selling prohibition for financial stocks included in the ban lists. However, the predictability diminished after that event. These findings also apply to the cross-predictability based solely on positive and large positive CDS innovations. Taken together, these results support previous conclusions that a migration of informed traders to the CDS market did not occur.

The remainder of this paper is organized as follows: Section 2 develops the hypotheses under investigation. Section 3 describes the data and variables employed in the study. Section 4 presents the empirical findings. Section 5 presents the final conclusions and puts forward a possible explanation for the results.

2. Development of Hypotheses

Given that the same company-specific information affects stock prices and credit spreads in opposite directions (Merton, 1974), investors holding private information about a firm may choose to trade in the stock market, the bond market or the CDS market. The decision as to which market to invest in depends on the leverage of the financial instrument and the liquidity of the markets (Easley *et al.*, 1998). Even though CDSs have greater leverage *vis-à-vis* stocks and bonds, they present the disadvantage of displaying higher transaction costs than stocks, which is why, under normal conditions, informed traders use stock markets to express their views (Hilscher *et al.*, 2015; Marsh and Wagner, 2016; and Wang and Bhar, 2014). Notwithstanding that, an interesting question that arises is whether the introduction of short-sale constraints in the stock market induces informed traders (not owning the stock) to exploit their negative information about a firm in the CDS market rather than the stock market.

This study exploits a unique event that occurred in European markets: the 2011 short-sale ban on European financial stocks. Under the coordination of the European Securities and Market Authority[6] (ESMA), financial market authorities in Belgium, France, Italy and Spain announced the prohibition of covered short-selling operations on financial stocks on 11 August 2011. The ban became applicable on the following trading day, i.e. 12 August 2011.[7] It is important to realize that this ban applied to stock markets but not to CDS markets, so that, in the face of this prohibition, investors could turn to the CDS market to bet on their negative information.

[6] See the public statement issued by ESMA (reference ESMA/2011/266).

[7] It is important to note that naked short-selling operations had already been banned in these European markets as of 2008. The current ban on covered short selling not only prohibited the creation of new net short positions, but also limited increases in existing ones, including intra-day operations. Nonetheless, positions resulting from formal market-making activities were exempted from the ban. The ban targeted both public and OTC markets.

To investigate whether informed traders migrated to the CDS market after the prohibition of short-selling operations in the stock markets, I start by evaluating whether CDS open interest soared by virtue of a substitution of credit protection buying in lieu of short interest in stocks.

H1: The short-selling ban raised the demand for credit risk protection in the CDS market.

Open interest measures are used herein with the objective of capturing the willingness of traders, in aggregate, to risk capital and the presence of informed trading. Indeed, these measures were used with the same goals previously in the literature. Oehmke and Zawadowski (2015) document that proxies for speculative trading motives are associated with larger net CDS positions. Launois and van Oppens (2003) show that open interest data is superior to volumes when the objective is to detect informed trading. In their study of takeover announcements, they note the advantages of utilizing open interest rather than volume as an indicator of market activity. First, open interest is less volatile than volume and, second, open interest is not affected by very short-term intraday noise trading. Focusing on futures, Bessembinder and Seguin (1993) use unexpected open interest as a proxy for the willingness of traders, in aggregate, to risk capital. Fodor *et al.* (2011) point out that changes in call and put open interest levels have predictive power for future equity returns. Finally, Silva (2015) shows that CDS open interest innovations help predict CDS rate changes and stock returns ahead, which conforms to the notion that unexpected open interest contains information not embedded in prices.

The migration of informed traders holding negative information to CDS markets would imply higher buying pressure and order imbalance, thereby forcing market makers to raise CDS premiums in order to limit their growing inventory and associated risk. In addition, if order imbalance emanates from the fact that incoming investors have superior information, dealers may increase bid-ask spreads to cover potential losses arising from adverse selection risk and information asymmetry. Increasing bid-ask spreads may also stem from synchronization risk (as dealers will also find more difficulties in hedging CDS protection by selling through short positions in stock markets[8]) and higher inventory risk if the ban drives up CDS spread volatility.

H2: Short-selling bans were followed by rising CDS premiums.

H3: By shifting informed trading from stock to CDS markets, short-selling bans raised transaction costs of CDSs.

Finally, if informed traders migrate from the stock to CDS market after the introduction of a short-selling ban, negative information should be incorporated first into CDS rates (and only thereafter into stock prices). As a result, CDS rates are expected to predict stock prices ahead and lead the price discovery process in the adjustment to negative information.

H4: By shifting informed trading from stock to CDS markets, short-selling bans raised the information flow from CDS to stock markets. The increase in the infor-

[8] Indeed, Das and Hanouna (2009) establish an association between CDS spreads and equity market liquidity. They argue that, since CDS contracts are actively hedged and hedging costs are incurred whether or not liquidity risk is systematic, illiquidity costs from the equity markets are transmitted into CDS spreads.

mation flow fueled the predictability of stock returns based on past information on CDS returns.

The next section presents a description of the sample and the definition of the variables analyzed in the study.

3. Sample Selection, Definition of Variables and Descriptive Statistics

The data used in the analysis consists of CDS (bid, ask and mid) quotes of five-year contracts[9] on senior unsecured debt, daily stock (bid, ask and closing) prices, total assets, total debt, the price-earnings ratio, price-to-book ratio and stock market capitalization for a sample of 56 European financial firms retrieved from Bloomberg. Weekly data on open interest are gathered from the Depository Trust & Clearing Corporation (DTCC).[10] Three measures of open interest are examined: the net notional amount, the gross notional amount and the number of contracts. The former stands for the sum of net protection bought (sold) by counterparties that are net buyers (sellers) of protection for a particular obligor and captures the stock of credit risk transferred in the CDS market. The gross notional amount denotes the aggregate notional amount of all the contracts written on an entity open in the market and is thereby driven by operations related to the management of counterparty exposure (e.g. portfolio compression cycles and novation). On balance, net notional amount variations uncover the hedging and speculative demand, whereas gross notional amount variations are a measure close to the traded volume.

After obtaining the aforementioned data, daily CDS and stock returns and bid-ask spreads are computed. The log change of CDS mid-rates is utilized as a proxy for CDS "returns". In effect, the percentage change in the credit spread well approximates the return on holding credit protection, as shown by Hilscher *et al.* (2015) and Wang and Bhar (2014). Stock returns are computed as the log change of daily closing prices. The percentage bid-ask spread is calculated as the difference between ask and bid prices divided by the mid-price, whereas the absolute bid-ask spread is calculated as the difference between ask and bid prices.

The ban on covered short-selling became applicable on the same day—12 August 2011—in France, Italy, Belgium and Spain. It affected 58 companies, of which 19 had active CDS contracts at that time. To measure the effect of the ban, it is important to draw counterfactuals, which requires conjecture about the behavior of the CDS trading activity of banned firms in the absence of the prohibition. That may be achieved through the identification of a group of firms (control group) that mimic the behavior of those included in the ban lists (treatment group) if the prohibition had not existed. This means that, until the ban, treatment- and control-group firms are expected to share common trends in CDS liquidity, price performance, trading activity and price formation. Comparing the behavior of these groups of firms allows us to gauge the effect of the ban because, if relevant, the ban should trigger a deviation from the common path. This technique makes it possible to draw counterfactuals and control for unobservable factors.

[9] The five-year tenor of CDS contracts is used because it is a common benchmark for practitioners and academics.

[10] DTCC covers 95% and 99% of the trading of CDS contracts on single-name references in terms of the number of contracts and total notional amounts, respectively (Gündüz *et al.*, 2013).

Table 1 List of Financial Firms Included in the Analysis

Name	Industry Group Name	Ban
Credit Agricole Sa	Banks	Yes
Ace Ltd	Insurance	No
Aegon Nv	Insurance	No
Allied Irish Banks Plc	Banks	No
Allianz Se-Reg	Insurance	No
Aon Plc	Insurance	No
Aviva Plc	Insurance	No
Barclays Plc	Banks	No
Banco Bilbao Vizcaya Argenta	Banks	Yes
Banco Comercial Portugues-R	Banks	No
Banco Espirito Santo-Reg	Banks	No
Bank of Ireland	Banks	No
Banca Monte Dei Paschi Siena	Banks	Yes
Bnp Paribas	Banks	Yes
Banco Popolare Sc	Banks	Yes
Commerzbank Ag	Banks	No
Axa Sa	Insurance	Yes
Credit Suisse Group Ag-Reg	Diversified Financials	No
Danske Bank A/S	Banks	No
Deutsche Bank Ag-Registered	Diversified Financials	No
Dexia Sa	Banks	Yes
Dnb Asa	Banks	No
Erste Group Bank Ag	Banks	No
Assicurazioni Generali	Insurance	Yes
Societe Generale Sa	Banks	Yes
Hannover Rueck Se	Insurance	No
Hsbc Holdings Plc	Banks	No
Ikb Deut Industriebank Ag	Banks	No
Permanent Tsb Group Holdings	Banks	No
Ing Groep Nv-Cva	Banks	No
Intesa Sanpaolo	Banks	Yes
Natixis	Banks	Yes
Legal & General Group Plc	Insurance	No
Lloyds Banking Group Plc	Banks	No
Mediobanca Spa	Diversified Financials	Yes
Muenchener Rueckver Ag-Reg	Insurance	No
Nordea Bank Ab	Banks	No
Old Mutual Plc	Insurance	No
Banca Popolare Di Milano	Banks	Yes
Prudential Plc	Insurance	No

Royal Bank of Scotland Group	Banks	No
Banco De Sabadell Sa	Banks	Yes
Banco Santander Sa	Banks	Yes
Scor Se	Insurance	Yes
Skandinaviska Enskilda Ban-A	Banks	No
Svenska Handelsbanken-A Shs	Banks	No
Standard Chartered Plc	Banks	No
Swedbank Ab—A Shares	Banks	No
Ubi Banca Spa	Banks	Yes
Ubs Group Ag-Reg	Diversified Financials	No
Unicredit Spa	Banks	Yes
Xl Group Plc	Insurance	No
Nn Group Nv	Insurance	No
Talanx Ag	Insurance	No
Banco Popular Espanol	Banks	Yes
Raiffeisen Bank International	Banks	No

Note: This table lists the financial entities covered in the analysis along with their GICS classification and a binary variable indicating which ones were subject to the ban.

The control group consists of 37 European financial firms (other banks and insurance companies) that were not subject to the covered short-selling ban on 12 August 2011, and whose stocks and CDS contracts were concomitantly active. There are various reasons justifying the inclusion of these firms in the control group. The first is that they belong to the same industry and operate in the same economic region as the firms covered by the ban. Importantly, these firms are subject to the same regulatory constraints, even though they are not domiciled in the same country as the firms covered by the ban.[11] They also share the same market and business cycle, interest rates and funding constraints. Drawing counterfactuals from firms of the same industry mitigates any bias resulting from differentials in liquidity dynamics, price performance or price formation arising from industry factors (see *Table 1*).

There is substantial empirical research suggesting that the CDS spreads of European banks are partially determined by the same common factors and it is thus expected that they, on average, follow the same long-term trend. Annaert *et al.* (2010) show that market-related and business-cycle-related variables are relevant determinants of the CDS spreads of eurozone banks.[12] Likewise, Samaniego-Medina *et al.* (2016) highlight the relevance of market returns and volatility and sovereign interest rates as determinants of CDS spread dynamics of European banks. Ötker-Robe *et al.* (2010) investigate the fundamental determinants of credit default risk for European large complex financial institutions and find that economic uncertainty along with business climate and earnings potential are among the most significant determinants of credit risk.

[11] It is also clear that some of these firms operate in more than one jurisdiction.

[12] In fact, their results reveal that the market and business cycle account for a greater adjusted-R^2 in explaining CDS spreads than do credit-risk-related variables.

Table 2 Descriptive Statistics

		Mean	Std. Dev.	Range	Percentiles		
					25	50	75
Banned Firms	CDS spread	279.8	129.4	447.1	172.6	250.0	353.2
	CDS bid-ask spread	5.2%	1.2%	4.0%	4.3%	4.6%	5.9%
	CDS cumulative change	47.9%	19.3%	64.0%	26.3%	51.2%	65.5%
	Net Notional Amount	1943.0	1368.5	4726.6	562.0	2084.2	2974.7
	Gross Notional Amount	28468.3	19546.4	54920.0	5566.5	32284.9	46373.6
	Number of contracts	3245.4	1972.3	5412.0	964.3	3736.6	5007.4
	Stock bid-ask spread	0.3%	0.4%	1.0%	0.0%	0.1%	0.6%
	Stock cumulative return	-30.4%	9.0%	33.0%	-36.2%	-27.7%	-24.1%
	Vol. 360 days	39.5%	6.7%	26.7%	37.0%	40.7%	43.8%
	Vol. 30 days	48.5%	14.3%	48.9%	38.0%	43.6%	61.1%
	D/A	40.1%	19.0%	69.9%	35.9%	42.1%	54.1%
	Market Cap.	11548.1	16199.0	61803.0	0.0	3483.5	22365.7
	Total Assets	384828.1	510450.5	1593815.0	0.0	102997.4	713197.7
	P/B	58.1%	29.0%	102.0%	32.1%	57.8%	82.3%
	P/E	8.9	3.1	11.3	6.6	7.9	11.8
Control Group Firms	CDS spread	291.5	385.9	1314.9	102.0	141.1	189.9
	CDS bid-ask spread	5.7%	1.5%	6.0%	4.5%	5.2%	7.1%
	CDS cumulative change	90.2%	378.3%	2231.0%	11.9%	28.1%	36.9%
	Net Notional Amount	1228.3	1094.7	4894.1	453.7	912.2	1577.6
	Gross Notional Amount	18782.8	16724.9	59558.5	4210.2	9476.5	32735.2
	Number of contracts	2464.5	1874.6	5568.7	655.1	1740.8	4285.0
	Stock bid-ask spread	0.8%	2.5%	14.0%	0.1%	0.1%	0.3%
	Stock cumulative return	-26.2%	11.2%	43.0%	-35.8%	-25.7%	-16.5%
	Vol. 360 days	39.3%	27.5%	133.6%	28.2%	31.6%	39.2%
	Vol. 30 days	42.4%	40.1%	236.4%	26.4%	32.5%	44.3%
	D/A	27.3%	18.7%	62.8%	5.1%	29.6%	41.0%
	Market Cap.	20090.9	22762.1	121697.3	5913.8	12983.5	29703.2
	Total Assets	509874.5	660107.4	3054862.3	72308.4	237110.4	709072.2
	P/B	86.7%	44.7%	205.0%	54.9%	90.3%	109.2%
	P/E	11.2	4.5	18.5	8.2	10.3	14.3
All Firms	CDS spread	287.5	319.4	1314.9	113.2	164.3	291.7
	CDS bid-ask spread	5.5%	1.4%	6.0%	4.4%	4.9%	6.8%
	CDS cumulative change	75.6%	305.2%	2231.0%	19.6%	33.6%	49.6%
	Net Notional Amount	1507.2	1243.3	4894.1	530.5	1211.3	2236.6
	Gross Notional Amount	22562.5	18275.3	59558.5	4733.0	17223.4	35658.0
	Number of contracts	2769.2	1927.7	5612.0	760.1	2709.8	4606.5
	Stock bid-ask spread	0.6%	2.0%	14.0%	0.1%	0.1%	0.4%
	Stock cumulative return	-27.7%	10.6%	46.0%	-35.6%	-27.3%	-18.9%
	Vol. 360 days	39.4%	22.5%	133.6%	29.3%	36.4%	41.8%

Vol. 30 days	44.5%	33.4%	236.4%	29.0%	37.0%	52.7%
D/A	31.7%	19.6%	71.2%	8.0%	35.9%	47.7%
Market Cap.	466589.2	610255.9	3054862.3	51479.1	224862.6	644622.6
Total Assets	17133.8	20965.7	121697.3	1824.9	12304.1	25378.3
P/B	76.8%	42.0%	205.0%	51.6%	73.8%	99.2%
P/E	10.4	4.2	18.5	7.1	10.0	12.7

Notes: This table reports sample statistics (mean, standard deviation, range and percentile 25, 50 and 75) for several variables. Descriptive statistics are presented for the entire sample of financial firms and disaggregated by treatment and control groups. The CDS spread, net notional amount (USD million), gross notional amount (USD million), number of contracts, historical volatility (estimated using a 30-day and a 360-day historical time horizon), debt-to-assets ratio (D/A), market capitalization (EUR million), total assets (EUR million), price-to-book ratio (P/B) and price-earnings ratio (P/E) are analyzed with respect to the latest information available as at the ban announcement date (11 August 2011). CDS bid-ask spreads and stock bid-ask spreads are averaged for the time window that spans the period from 1 July 2011 to 11 August 2011. CDS cumulative change (stock cumulative return) corresponds to the appreciation of the CDS contract (stock) in the time window spanning the period from 1 July 2011 to 11 August 2011.

Consistent with the results of the aforementioned studies, an exploratory analysis corroborates the view that both groups shared the same common trend between 4 January 2011 and 10 August 2011 in terms of open interest growth (the correlation of the average growth of net open interest of the two groups equals 0.78), the bid-ask spread (the correlation of the average bid-ask spread of the two groups equals 0.88) and returns (the correlation between the average CDS returns of the two groups equals 0.92).

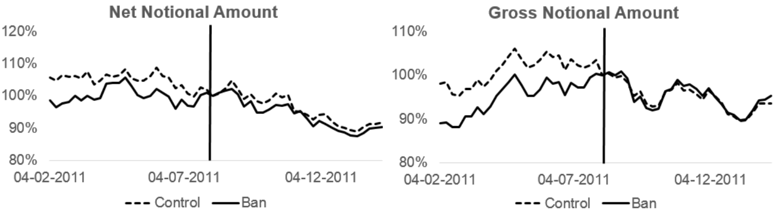
Table 2 outlines the sample of firms attending to the information available between 1 July 2011 and 11 August 2011. In that time window, the stocks of these firms depreciated 27.7% on average, whereas the CDS spreads climbed 75.6%. Interestingly, the control group outperformed the treatment group in the stock market in that period (-26.2% vs. -30.4%), but the CDSs of the treatment group saw lower appreciation than did those of the control group (47.9% vs. 90.2%).

The treatment- and control-group firms displayed similar median stock bid-ask spreads prior to the ban (0.1%), but the average bid-ask spreads differed substantially due to the existence of an outlier in the treatment group (with a bid-ask spread of 14.0%). With respect to transaction costs in the CDS market, firms in the control group exhibited, on average, greater percentage CDS bid-ask spreads (5.7% vs. 5.2% for control- and treatment-group firms, respectively).

One of the objectives of the ban was to reduce stock price volatility, so it is interesting to compare the pre-ban stock volatility of the firms subject to the ban with that of the control group. Indeed, while the stock price volatility of the two groups over a 360-day span is very close (39.5% and 39.3% for the treatment- and control-group firms, respectively), the stock price volatility over a 30-day span is clearly greater among banned firms (48.5% and 42.4% for the treatment- and control-group firms, respectively). This result is in line with the argument presented by the financial authorities of France, Belgium, Italy and Spain that volatility of financial stocks under their supervision was too high in comparison with historical standards.

The average net notional amount was greater for banned firms than for the control-group firms (USD 1,943 million vs. USD 1,228 million). Regarding size, the treatment group presents, on average, lower market capitalization (EUR 11,548 mil-

Figure 1 The Pattern of Open Interest



Notes: The figure on the LHS (RHS) plots the average accumulated growth of the net (gross) notional amount of treatment- and control-group firms for the six-month periods that preceded and followed the ban. To ease the comparison of the path of the two groups after introduction of the ban, the growth is evaluated having as a reference the net (gross) notional amount on 5 August 2011, which is to say that the series corresponds to an index with a value of 100% on that date.

lion vs. EUR 20,090 million for the treatment- and control-group firms, respectively) and a lower book value of assets (EUR 384,828 million vs. EUR 509,874 million for the treatment- and control-group firms, respectively). Finally, it is also noteworthy that both the price-to-book ratio (*P/B*) and price-earnings ratio (*P/E*) were higher for control-group firms than for banned firms.

4. Results

4.1 The Effect of the Ban on CDS Open Interest

The starting point of the analysis concerns the effect of the ban on covered short selling over the notional amounts outstanding of CDS contracts, i.e. open interest. H1 postulates that the ban should elevate the demand for protection buying in CDS markets as a consequence of informed traders' migration from stock to CDS markets. A direct way to examine this hypothesis is to assess whether open interest changed significantly with the implementation of the ban. As other systematic factors may also affect the open interest dynamics (Silva *et al.*, 2015), I examine whether the equally weighted average growth rate of the open interest of entities subject to the ban is higher than that of other financial firms not subject to the ban. *Figure 1* plots the average cumulative growth of the gross and net notional amount of the treatment and control groups in the six-month periods that preceded and followed the ban. To ease the comparison between the two groups, the series are scaled with respect to the latest information disclosed by DTCC before the ban announcement date (5 August 2011), where the indices take the value of one. A visual inspection of *Figure 1* shows that the net notional amount trended downwards after the ban announcement in both groups.

In what follows, the open interest percentage growth of the treatment and control groups is compared in several time horizons having as a baseline the level of open interest on 5 August 2011, i.e. the last disclosure of open interest by DTCC prior to the ban announcement. The analysis of the alternative spans is undertaken with the goal of identifying and distinguishing short-term and long-lasting effects and providing insights for the full time frame of the ban. Columns [1] and [2] of *Table 3* report the difference of the average and median open interest growth of treatment- and control-group entities along with the corresponding statistical significance. The net open interest of treatment-group entities presents, on average, a lower growth

Table 3 Open Interest and Implementation of the Ban

	Panel A		Panel B	
	Full sample		Matched-firms sample	
	Mean diff	**Median diff**	**Mean diff**	**Median diff**
GNA growth t_{-6}, t_{+8}	0.5% (./.)	1.1% (./.)	0.8% (.)	1.4% (./.)
GNA growth t_{-6}, t_{+15}	1.2% (./.)	2.1% (./.)	1.3% (.)	1.2% (./.)
GNA growth t_{-6}, t_{+50}	0.2% (./.)	-1.1% (./.)	-0.3% (.)	0.5% (./.)
GNA growth t_{-6}, t_{+78}	-1.3% (./.)	-2.3% (./.)	0.7% (.)	0.0% (./.)
GNA growth t_{-6}, t_{+141}	-0.2% (./.)	-2.1% (./.)	-0.2% (.)	2.7% (./.)
GNA growth t_{-6}, t_{+183}	1.2% (./.)	0.6% (./.)	0.4% (.)	0.7% (./.)
NNA growth t_{-6}, t_{+8}	-0.6% (./.)	1.0% (./.)	-1.1% (.)	-1.3% (./.)
NNA growth t_{-6}, t_{+15}	-2.4% (*/*)	-2.7% (./.)	-3.0% (*)	-2.1% (*/**)
NNA growth t_{-6}, t_{+50}	-2.8% (./.)	-1.4% (./.)	-5.0% (**)	-7.3% (./*)
NNA growth t_{-6}, t_{+78}	-2.5% (./.)	-2.9% (./.)	-6.2% (**)	-6.6% (./*)
NNA growth t_{-6}, t_{+141}	-1.6% (./.)	-1.0% (./.)	-6.0% (.)	-2.1% (./.)
NNA growth t_{-6}, t_{+183}	-1.1% (./.)	-0.4% (./.)	-8.9% (.)	0.9% (./.)
Num. contracts growth t_{-6}, t_{+8}	1.1% (./.)	1.1% (./.)	1.4% (.)	1.4% (./.)
Num. contracts growth t_{-6}, t_{+15}	1.8% (./.)	2.3% (*/*)	2.1% (.)	1.9% (./.)
Num. contracts growth t_{-6}, t_{+50}	0.4% (./.)	-0.6% (./.)	0.4% (.)	-1.6% (./.)
Num. contracts growth t_{-6}, t_{+78}	-0.8% (./.)	-2.7% (./.)	0.7% (.)	-2.2% (./.)
Num. contracts growth t_{-6}, t_{+141}	0.3% (./.)	0.0% (./.)	-1.0% (.)	0.5% (./.)
Num. contracts growth t_{-6}, t_{+183}	2.0% (./.)	1.3% (./.)	0.2% (.)	1.9% (./.)

Notes: Panel A, presents differences in the average and median open interest growth of CDSs written on firms on the ban lists and firms in the control group after implementation of the ban. Panel B, presents differences in the average and median open interest growth of CDSs written on firms on the ban lists and matched firms after implementation of the ban. Having the open interest on 5 August 2011 as a baseline (t_{-6}, with t_0 as the announcement date), the growth rate is calculated for each firm and for six alternative dates (19 August 2011 [t_{+8}], 26 August 2011 [t_{+15}], 30 September 2011 [t_{+50}], 28 October 2011 [t_{+78}], 30 December 2011 [t_{+141}] and 10 February 2012 [t_{+183}]). Three measures of open interest are considered: gross notional amount (GNA), net notional amount (NNA) and number of contracts. Average and median growth are calculated for each group (treatment and control groups) and statistical inference is conducted by means of parametric tests (ordinary *t*-test and *t*-test corrected for unequal variances) and non-parametric tests (Wilcoxon/Mann-Whitney and Kruskal-Wallis tests). To save space, only the statistical significance of the average and median differences is reported. Significance at the 10%, 5% and 1% levels is indicated by *, ** and ***, respectively.

rate than their peers in the control group during the ban period. For instance, in the span [t_{-6},t_{+50}] (with t_0 as the announcement day), net open interest growth was, on average, 2.8 percentage points lower for the treatment-group entities *vis-à-vis* control-group entities. Qualitatively similar results are found for the remaining analyzed time frames. The median difference of the net notional amount growth is also negative in the various time windows apart from the span [t_{-6},t_{+8}].

Statistical inference is conducted by means of *t*-tests on the difference of the average and median net open interest growth of the groups. The results from parametric tests show that the differences in the patterns of the treatment and control groups are not statistically significant during the ban period, with the exception of the window $[t_{-6},t_{+15}]$, where the difference is negative and statistically significant (at the 10% significance level). Regarding non-parametric tests, Wilcoxon/Mann-Whitney and Kruskal-Wallis tests do not reject the hypothesis that the median net notional amount growth is equal for the two groups in the ban period.

A similar approach is conducted to evaluate whether the growth rates of the gross notional amount and the number of contracts outstanding differ substantially for firms in the two groups. Notably, the results of the parametric and non-parametric tests indicate that the treatment and control groups experienced similar gross notional amount growth trends during the ban. As for growth of the number of contracts, the mean and median differences are not statistically different except for one case (in the window $[t_{-6},t_{+15}]$, Wilcoxon/Mann-Whitney and Kruskal-Wallis tests indicate that the median growth of the number of contracts of firms subject to the ban is higher than that of control-group firms). Taken jointly, these results suggest that the open interest pattern of the groups did not deviate significantly after implementation of the ban. These findings cast some doubt on the existence of a migration of informed investors from stock to CDS markets and do not lend support to H1.

4.1.1 Robustness Check

To strengthen the conclusions of the analysis, an alternative approach based on a matched sample and difference-in-differences analysis is also undertaken. Brunnermeier and Oehmke (2014) developed a theoretical model predicting that predatory short selling is more likely when a financial institution is close to its capital constraint or violates its capital constraint even in the absence of short selling. This signifies that the level of distress of a financial institution turns it into a target for predatory short selling, thus leading to its selection by regulators for inclusion in a ban lists. Consistent with that reasoning, Alves *et al.* (2016) show that the path of short-run volatility of financial stocks *vis-à-vis* the long-run volatility and the level of default risk at the time of the ban help explain why some countries banned short-selling activity in 2011 while others did not.

Certainly, the perceived level of distress of financial institutions can be captured by stock volatility variations and the CDS spread level. These variables are used herein as instrument variables to find pairs of banned and not-banned financial stocks with the same level of distress around the ban event and thus with the same likelihood of being selected by regulators for inclusion in a ban lists. Each financial firm subject to the ban is matched with another financial firm not subject to the ban with similar characteristics in terms of those variables using an approach similar to Battalio and Stulz (2011). In doing so, each banned financial firm is matched with the control-group firm presenting the smallest sum of the squared percentage difference in CDS spread and volatility variation[13] prior to the ban. No control-group firm is used twice.[14] CDS spreads and volatility variation are measured on the day prior to the ban announcement date, i.e. 10 August 2011.

[13] The volatility variation is defined as the difference between the 30-day volatility and 360-day volatility on the day prior to the ban announcement date.

[14] However, non-tabulated results show that the conclusions survive if the control-group firms are used more than once.

Matching firms subject to the ban with other financial firms displaying a similar likelihood of being included in the ban reduces concerns of endogeneity and reverse causality. The results of the difference-in-differences analysis are displayed in *Table 3*, columns [3] and [4]. This alternative approach reinforces the previous conclusions by showing that the net open interest of banned firms declined at a faster pace than that of the matched firms after implementation of the ban. In effect, it can be seen that the treatment-group firms saw lower net notional amount growth than their peers in the ban period, with that difference being statistically significant in three of the analyzed periods: $[t_{-6},t_{+15}]$, $[t_{-6},t_{+50}]$, and $[t_{-6},t_{+78}]$.

4.2 The Effect of the Ban on CDS Spreads

The previous sub-section showed that the open interest of European financial firms did not change significantly with the ban. Still, a simple attempted migration of informed traders could have induced major CDS dealers to raise bid-ask spreads and CDS premiums for all end-users, deterring participants from purchasing CDSs and leaving open interest virtually unaltered. This is because such migration would bring about higher order imbalance with negative effects on the inventory risk of major dealers. In addition, since some of those investors that attempt to move to the CDS market are likely to possess better information than dealers, the latter would face greater adverse selection risk in trades with end-clients in general. A straightforward way to deal with greater net demand for insurance is to increase CDS premiums.

This prediction is examined herein by means of an event study analysis. In the first pass, the average cumulative CDS returns of the treatment and control groups are compared. In the second pass, excess cumulative returns with respect to the iTraxx Europe Senior Financials[15] and sovereign CDSs of the country where the financial company is domiciled are calculated. Industry (country) excess returns are computed as the difference between the CDS returns of financial firms and the returns of the industry index (sovereign CDSs).

When analyzing industry excess returns, the influence of industry-wide news on the price performance of the firms is removed. Thereby, idiosyncratic shocks to each firm are isolated from the presence of sectoral or/and macroeconomic information. When computing country excess returns, the influence of sovereign risk on the financial sector is eliminated. Sovereign CDS spreads constitute, in most cases, a ceiling for the spreads of financial firms insofar as there are implicit or explicit government guaranties for the deposits and debt of those firms. As a result, sovereign CDS spreads and spreads of financial firms tend to move in parallel. Country excess returns translate changes of the CDS spreads of financial firms that are not explained by movements in the credit risk of sovereigns.[16] (*Table 4*)

The left-hand side of *Figure 2* plots the path of the CDS spreads of the two groups between 4 January 2011 and 15 February 2012. In the first pass, daily CDS

[15] The Markit iTraxx Europe Senior Financial index includes 25 equally weighted CDSs on investment-grade European entities.

[16] It is important to highlight the fact that the market turmoil after 2010 was fueled primarily by the uncertainty and speculation surrounding some sovereigns' ability to fulfill their obligations. Indeed, the association between the financial sector risk and the sovereign risk peaked after the 2008 financial crisis and remained at high levels during the sovereign debt crisis.

Figure 2 The Pattern of CDS Spreads

Notes: The figure on the LHS (RHS) plots the average cumulative CDS (country excess) returns of treatment- and control-group firms from 4 January 2011 until 10 February 2012, having the ban announcement date as a baseline.

Table 4 Computation of Returns and Excess Returns

Raw Returns	$rCDS_t = \ln\left(\frac{CDS_t}{CDS_{t-1}}\right)$, with $rCDS_t$ and CDS_t as the CDS raw returns and the CDS spread in *t*, respectively.
Industry Excess Returns (*IER*)	$IER_t = rCDS_t - rCDS_t^I$, with $rCDS_t^I$ as the return of the iTraxx Europe Senior Financials in *t*.
Country Excess Returns *(CER)*	$CER_t = rCDS_t - rCDS_t^S$, with $rCDS_t^S$ as the return of the CDS of the country (in which the financial firm is domiciled) in *t*.

returns are averaged for the entire sample and for the two groups under study separately. With the aim of capturing changes in the valuation of the baskets, price indices are formed for each of the groups, having an equal-weighted investment on 11 August 2011 as a baseline (where the value of the three indices is one). A visual evaluation of the left-hand side of *Figure 2* shows that the treatment- and control-group indices evolved similarly until the end of the third quarter of 2011. Then a positive gap emerges, becoming clearly wider during the last quarter of 2011. The right-hand side of *Figure 2* plots the path of country excess cumulative returns (indices) and shows that the control group was outperforming the treatment group by the end of October 2011, but an inversion of that trend occurred in the last quarter of 2011. Nevertheless, it is important to highlight that the aforementioned gaps almost vanished before the end of the ban.

To get a detailed picture of the effect of the ban on the performance of CDSs written on the entities in the treatment group, an event study analysis is conducted. CDS raw returns, industry excess returns and country excess returns are computed for the time horizon consisting of 130 trading sessions before and after the ban announcement. The performance of the treatment- and control-group firms is then compared in the following windows: $[t_1;t_5]$, $[t_1;t_{10}]$, $[t_1;t_{30}]$, $[t_1;t_{60}]$, $[t_1;t_{90}]$ and $[t_1;t_{130}]$, with t_0 as the announcement date. The analysis uses both short and long time windows for two reasons. The first is to distinguish short-term and long term effects and, more precisely, to unveil possible price reversals. In addition, CDSs are traded in OTC markets, which means that there are search frictions in that an investor may take time to find a counterparty to perform a trade (Duffie *et al.*, 2007). As a result, information flows more gradually in CDS markets than in stock exchanges.

The observations of the time window $[t_{-130};t_{-10}]$ are utilized to compute the expected standard error of (excess) CDS returns. Statistical inference is conducted by means of parametric and non-parametric tests.

An important issue of this analysis concerns the fact that the ban was disclosed simultaneously for the members of the ban lists. Due to event clustering, observations are not independent by virtue of the cross-dependence of returns. Therefore, ordinary *t*-tests are not robust and may understate the true standard errors. Two alternatives are employed to address this issue. The first is the "crude adjustment" described by Brown and Warner (1985). The second is the cross-dependency adjustment of standard errors developed by Kolari and Pynnönen (2010). With respect to the non-parametric approach, the GRANK-*t* test developed by Kolari and Pynnonen (2011) is carried out.[17] The authors show that the GRANK test outperforms previous rank tests and is robust to serial correlation and event-induced volatility. Moreover, the GRANK test exhibits superior empirical power relative to parametric tests such as those developed by Patell (1976) and Boehmer *et al.* (1991).

Panels A–C of *Table 5* report the results of parametric tests. Panel A tabulates various combinations of average cumulative returns by group and by event window (along with the corresponding *t*-statistics and level of statistical significance). Both groups recorded negative cumulative returns in the five trading days after the ban announcement. These returns are statistically significant for the treatment group, but not for the control group. Moreover, the difference between the average cumulative returns of treatment- and control-group firms is negative and statistically significant. The results are virtually identical when using industry excess returns instead of raw returns. Panel C shows that both groups witnessed positive country excess returns in the five trading days that followed the announcement, meaning that the perceived credit risk decreased to a lesser extent for financial firms than for the corresponding countries in which they are domiciled. Notably, the difference between the average country excess returns of the two groups is not statistically significant. Bearing these results in mind, it cannot be concluded that the ban on financial stocks contributed to short-term order imbalance or demand pressure in CDSs. Indeed, during the short time window analyzed, the CDS spreads of banned firms actually declined. Even though the difference between the spreads of financial firms and sovereigns mounted, the surge was more pronounced for CDSs written on obligors not subject to the ban.

The price performance in the spans $[t_1;t_{10}]$, $[t_1;t_{30}]$, $[t_1;t_{60}]$, $[t_1;t_{90}]$ and $[t_1;t_{130}]$ is also assessed. It is noteworthy that the difference in the average cumulative raw returns of the groups is positive in only two of the aforementioned windows: $[t_1;t_{30}]$ and $[t_1;t_{90}]$. In the first case, that difference (equal to 0.22%) is not statistically significant or economically meaningful. In the time frame $[t_1;t_{90}]$, both groups recorded positive CDS spreads changes, with the difference in their performance not being statistically significant. The examination of country excess returns suggests that price performance in the latter time span was driven mostly by sovereign risk, given that country cumulative excess returns are substantially smaller than cumulative raw returns in both groups. Importantly, the difference in the average country cumulative

[17] The CUMRANK-*t* test statistic corrected for event-induced volatility and cross-correlation using the approach described by Hagnas and Pynnönen (2014) is also computed and leads to qualitatively similar conclusions.

excess returns of the groups is also not statistically significant in any of the time frames analyzed.

Panel D of *Table 5* details the results of non-parametric tests (GRANK-*t* test). The null hypothesis of no abnormal performance is not rejected for the various combinations of event windows, groups and type of returns utilized in the assessment. The aforementioned conclusions are reinforced while using matched samples to test for differences in cumulative returns after implementation of the ban. In doing so, the procedure described in Subsection 4.1 is undertaken. It can be observed that the results are qualitatively similar to those depicted above. Combining all the results, the hypothesis of the inexistence of short-term price pressure or order imbalance after the ban announcement is not rejected. As for longer spans, it is true that average cumulative raw CDS returns of the treatment group are higher than those of the control group in the time frame $[t_1;t_{90}]$. Notwithstanding that, this surge of CDS rates is explained principally by the evolution of sovereign CDS spreads of the countries of domicile of banned financial firms. Not least important, the difference between the cumulative raw returns of the groups is not statistically significant. Overall, these results do not lend support to the hypothesis that informed investors migrated from stock to CDS markets after the ban.

4.3 The Effect of the Ban on CDS Transaction Costs, Volatility and Tail Risk

The results presented in previous subsections conform to the notion that informed investors were unable to migrate from stock to CDS markets. A possible explanation for the inexistence of such migration may be related to the relatively high bid-ask spread faced when trading CDSs. Not least important, the shift of informed traders from stock to CDS markets (and heightened counterparty risk) could itself trigger an upsurge of bid-ask spreads and a decline of liquidity provision. This prediction is supported by the notion that bid-ask spreads reflect information asymmetry in addition to search costs, inventory costs and processing costs. As argued in the microstructure literature—e.g. Copeland and Galai (1983) and Bagehot (1971)—liquidity suppliers tend to raise bid-ask spreads when they perceive that other traders hold superior information. So, if more informed traders move to the CDS market, dealers may respond with higher bid-ask spreads to reduce adverse selection risk.

This hypothesis is examined herein through the lens of a two-way fixed effect model as in Boehmer *et al.* (2013).

$$BAS_{i,t} = \alpha_i + \gamma_t + \theta \times BAN_{i,t} + u_{i,t} \tag{1}$$

where $BAS_{i,t}$ stands for the bid-ask spread of a CDS on firm *i* on day *t*, α_i denotes the firm's fixed effect, γ_t stands for the calendar fixed effect, and BAN is an indicator variable for which the value of one is assigned if and only if the shorting ban is in effect for the stocks of firm *i* on *t* and zero otherwise.

θ represents the incremental change of the bid-ask spread due to the implementation of the ban. In effect, it measures the deviation of the liquidity pattern of CDSs on banned firms from the expected pattern they would have if the ban had not been introduced. The pattern of the control group replicates the expected behavior of the treatment group in the absence of the ban, allowing generation of the counterfactual and removing the influence of unobserved covariates. The identification of the marginal effect of the ban on the bid-ask spreads is undertaken using the cross-section of firms, i.e. comparing the pattern of banned firms to non-banned firms on the same day and through time by comparing trends before and after the ban announcement.

The output of the estimation of (1) is presented in columns [1] and [2] of *Table 6*, Panel A, using the two alternative definitions of bid-ask spreads. The estimated coefficient for *BAN* is positive and statistically significant in both cases. On average, the ban is associated with an increase of 0.1 percentage points in the percentage bid-ask spread. Although statistically significant, it represents a modest 3.6% of the average percentage bid-ask spread of CDSs on firms affected by the ban, so its economic relevance is questionable.

Next, attention is shifted to the effect of the ban on the volatility and tail risk of CDSs. In doing so, a procedure similar to the one described above is adopted. Equation (1) is estimated for five dependent variables that proxy the price risk of CDSs using monthly data: (i) volatility, defined as the standard deviation of the CDS returns in a one-month period; (ii) upside volatility, defined as the standard deviation of positive or null CDS returns in a one-month period; (iii) downside volatility, defined as the standard deviation of negative CDS returns in a one-month period; (iv) kurtosis, defined as the monthly kurtosis of CDS returns; and (v) skewness, defined as the monthly skew of CDS returns.

In a first pass, the effect of the ban on CDS return volatility is assessed. If informed traders migrated to the CDS market generating greater order imbalance, higher short-run volatility should emerge for CDSs on firms affected by the ban. If so, it is reasonable to assume that it would have affected upward volatility (linked to negative information about the firm and more difficult to utilize in the stock market due to the ban) to a greater extent than downward volatility. Columns [3] to [5] of *Table 6*, Panel A, present the results of regressing volatility, upside volatility and downside volatility against a dummy variable assuming the value of one after the introduction of the ban if the firm is included in the treatment group and fixed id effects and calendar effects. Contrary to what was anticipated, the regression results indicate that volatility actually declined for the treatment-group CDSs *vis-à-vis* the control-group CDSs. The estimated coefficient for *BAN* is negative and statistically significant in the equations having volatility, upward volatility and downward volatility as dependent variables.

A similar exercise is conducted for kurtosis and skewness with the aim of assessing the effect of the ban on the tail risk of the CDSs on firms affected by the prohibition. Intuitively, the ban may contribute to a decline of CDSs' skewness and kurtosis by virtue of a more continuous inflow of negative information brought by informed investors. Looking at columns [6] and [7] of *Table 6*, Panel A, $\hat{\theta}$ is negative for both cases. However, it is statistically significant only in the case of kurtosis.

As a robustness test, each banned firm is matched with a control firm with a similar likelihood of default using the approach employed in Subsection 4.1. Then, I take the difference of the value of each representative variable (bid-ask spread, volatility, upward volatility, downward volatility, kurtosis and skewness) for the banned

Table 5 Price Performance after Implementation of the Ban

		[1;5]	[1;10]	[1;30]	[1;60]	[1;90]	[1;130]
Panel A: Raw CDS Returns							
Entire Sample	CAR	-4.42%	10.07%(*/*)	18.20%(**/*)	9.57%	27.39%(*/*)	-11.28%
	t-stat	(-1.19/-1.14)	(1.92/1.84)	(2.00/1.92)	(0.73/0.71)	(1.70/1.67)	(-0.58/-0.57)
Ban list	CAR	-9.68%(*/*)	6.45%	18.34%	7.54%	37.75%(./*)	-12.50%
	t-stat	(-1.80/-1.89)	(0.85/0.89)	(1.38/1.46)	(0.40/0.42)	(1.62/1.73)	(-0.45/-0.48)
Control Group	CAR	-1.71%	11.93%(***/**)	18.13%(**/**)	10.62%	22.06%(*/.)	-10.65%
	t-stat	(-0.58/-0.51)	(2.86/2.50)	(2.50/2.20)	(1.02/0.91)	(1.72/1.54)	(-0.69/-0.62)
Ban list vs. Control Group	CAR	-7.97%(**/**)	-5.48%	0.22%	-3.08%	15.68%	-1.85%
	t-stat	(-2.74/-2.62)	(-1.33/-1.27)	(0.03/0.03)	(-0.31/-0.29)	(1.27/1.22)	(-0.12/-0.12)
Ban list—Matching Firms	CAR	-7.18%(**/**)	-5.61%	-0.98%	-0.44%	17.30%	2.88%
	t-stat	(-2.43/-2.38)	(-1.34/-1.31)	(-0.14/-0.13)	(-0.04/-0.04)	(1.38/1.35)	(0.19/0.19)
Panel B: Industry Excess CDS Returns							
Entire Sample	CAR	-5.12%	10.79%(*/*)	21.75%(**/**)	11.17%	32.07%(*/*)	-10.68%
	t-stat	(-1.24/-1.19)	(1.85/1.78)	(2.14/2.07)	(0.77/0.75)	(1.79/1.76)	(-0.50/-0.49)
Ban list	CAR	-10.38%(*/*)	7.18%	21.37%	8.61%	41.90%(./*)	-11.90%
	t-stat	(-1.79/-1.90)	(0.87/0.93)	(1.49/1.59)	(0.42/0.45)	(1.66/1.80)	(-0.40/-0.43)
Control Group	CAR	-2.41%	12.65%	21.95%(**/**)	12.49%	27.02%(*/*)	-10.05%
	t-stat	(-0.72/-0.64)	(2.66/2.37)	(2.66/2.38)	(1.05/0.96)	(1.84/1.69)	(-0.57/-0.52)
Ban list vs. Control Group	CAR	-7.97%(**/**)	-5.48%	-0.58%	-3.88%	14.88%	-1.85%
	t-stat	(-2.74/-2.46)	(-1.33/-1.19)	(-0.08/-0.07)	(-0.38/-0.35)	(1.20/1.08)	(-0.12/-0.11)

		[1;5]	[1;10]	[1;30]	[1;60]	[1;90]	[1;130]
Panel C: Country Excess CDS Returns							
Entire Sample	CAR	8.58%(**/**)	9.63%	4.81%	3.05%	2.72%	-4.82%
	t-stat	(2.05/2.04)	(1.61/1.62)	(0.46/0.47)	(0.21/0.21)	(0.15/0.15)	(-0.22/-0.23)
Ban list	CAR	6.74%	9.83%	-0.76%	-6.40%	5.65%	-5.51%
	t-stat	(1.32/1.43)	(1.36/1.47)	(-0.06/-0.07)	(-0.36/-0.39)	(0.26/0.28)	(-0.21/-0.23)
Control Group	CAR	9.52%(**/**)	9.53%	7.68%	7.91%	1.21%	-4.47%
	t-stat	(2.31/2.27)	(1.60/1.60)	(0.74/0.75)	(0.54/0.54)	(0.07/0.07)	(-0.21/-0.21)
Ban list vs. Control Group	CAR	-2.78%	0.29%	-8.43%	-14.31%	4.44%	-1.04%
	t-stat	(-0.82/-1.00)	(0.06/0.07)	(-1.02/-1.23)	(-1.22/-1.48)	(0.31/0.37)	(-0.06/-0.07)
Ban list—Matching Firms	CAR	-3.10%	-0.67%	-10.27%	-13.38%	7.80%	-2.87%
	t-stat	(-0.90/-0.90)	(-0.14/-0.14)	(-1.22/-1.22)	(-1.13/-1.12)	(0.54/0.53)	(-0.16/-0.16)
Panel D: GRANK-*t* test statistic							
Raw CDS Returns	Entire Sample	-0.66	0.98	1.36	0.29	1.08	-0.66
	Ban list	-0.82	0.03	1.48	0.29	1.58	-0.78
	Control Group	-0.54	1.55	1.26	0.28	0.75	-0.57
	Ban list—Matching Firms	-0.59	-0.66	0.01	-0.06	1.14	0.31
Industry Excess CDS Returns	Entire Sample	-0.74	0.78	1.47	0.32	1.13	-0.64
	Ban list	-0.82	0.06	1.51	0.29	1.57	-0.74
	Control Group	-0.68	1.16	1.44	0.33	0.88	-0.58
Country Excess CDS Returns	Entire Sample	0.77	0.61	0.05	-0.13	0.23	-0.42
	Ban list	-0.04	0.02	-0.47	-0.95	0.26	-0.69
	Control Group	1.20	0.90	0.35	0.35	0.19	-0.23
	Ban list—Matching Firms	-0.60	-0.38	-0.48	-2.27(**)	0.55	-0.12

Notes: The table presents cumulative (excess) CDS returns for six alternative time horizons: [1;5], [1;10], [1;30], [1;60], [1;90] and [1;130]. The results are tabulated for the entire sample and by group. Statistical inference is conducted using parametric and non-parametric tests. With respect to the former, standard errors are computed by taking into account the cross-dependence of the observations, namely by employing the crude adjustment of Brown and Warner (1985) and the cross-dependency adjustment of standard errors developed by Kolari and Pynnönen (2010). *t*-statistics are displayed in parentheses. Panels A–C tabulate results utilizing raw returns, industry excess returns and country excess returns. Panel D reports the GRANK-*t* test statistic, a non-parametric test developed by Kolari and Pynnönen (2011) when event-induced volatility and cross-correlation are present. The results of the treatment group are compared to those of the control group and to those of matched firms. Significance at the 10%, 5% and 1% levels is indicated by *, ** and ***, respectively.

Table 6 Bid-Ask Spread, Volatility and Tail Risk after the Ban Implementation
Panel A

	Percentage bid-ask spread	Absolute bid-ask spread	Volatility	Vol+	Vol-	Kurt	Skew
Const	0.065*** (543.28)	12.988*** (161.97)	0.032*** (96.20)	0.021*** (68.57)	0.022*** (66.48)	1.403*** (20.67)	-0.036 (-1.40)
BAN	0.001** (2.33)	3.463*** (9.94)	-0.003* (-1.92)	-0.003*** (-2.69)	-0.002* (-1.70)	-0.450* (-1.76)	-0.077 (-0.74)
R-squared	49.4%	53.6%	62.7%	44.7%	56.7%	28.9%	19.7%
Cross-Section F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Calendar F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Frequency	Daily Obs.	Daily Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.
N	29957	29957	1445	1432	1431	1442	1444
S.E.	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)
Sample	1/04/2010– –2/09/2012	1/04/2010– –2/09/2012	2010M01- –012M01	2010M01– –2012M01	2010M01– –2012M01	2010M01– –2012M01	2010M01– –2012M01

Notes: Panel A presents results of regressing a market quality proxy variable (bid-ask spread, volatility, upward volatility, downward volatility, kurtosis and skewness) on a constant and on the binary variable BAN (which assumes the value of one in the period of the ban if the firm's stocks are subject to the short-selling restriction and zero otherwise). All panel data models include cross-section fixed effects and calendar fixed effects. The regression of the bid-ask spread on the aforementioned explanatory variables is run using daily data. This equation is estimated for the span ranging from 4 January 2010 to 9 February 2012. The regressions having volatility, upward volatility, downward volatility, kurtosis and skewness as the dependent variables are run utilizing monthly data. These equations are estimated for the span ranging from 4 January 2010 to 31 January 2012. White heteroskedasticity-consistent standard errors are used in the computation of *t*-statistics. (***), (**) and (*) indicate that the variable is statistically significant at the 1%, 5% and 10% level, respectively.

Table 6 Bid-Ask Spread, Volatility and Tail Risk after the Ban Implementation
Panel B

	Percentage bid-ask spread	Absolute bid-ask spread	Volatility	Vol+	Vol-	Kurt	Skew
Const	-0.007*** (-17.18)	1.312*** (12.97)	0.032*** (96.20)	0.021*** (68.57)	0.022*** (66.48)	1.403*** (20.67)	-0.036 (-1.40)
BAN	0.001 (0.87)	9.790*** (37.30)	-0.003* (-1.92)	-0.003*** (-2.69)	-0.002* (-1.70)	-0.450* (-1.76)	-0.077 (-0.74)
R-squared	44.0%	69.9%	62.7%	44.7%	56.7%	28.9%	19.7%
Cross-Section F.E.	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Calendar F.E.	No	No	No	No	No	No	No
Frequency	Daily Obs.	Daily Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.	Monthly Obs.
N	9529	9529	449	443	447	449	448
S.E.	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)	White diagonal standard errors & covariance (d.f. corrected)
Sample	1/04/2010– –2/09/2012	1/04/2010– –2/09/2012	2010M01– –2012M01	2010M01– –2012M01	2010M01– –2012M01	2010M01– –2012M01	2010M01– –2012M01

Notes: In Panel B, each banned firm is matched with a control firm with similar characteristics. Then, I take the difference of the value of each representative variable (bid-ask spread, volatility, upwards volatility, downwards volatility, kurtosis and skewness) for banned firms and matched firms. That difference is regressed on a constant and on the binary variable *BAN* (which assumes the value of one in the banning period and zero otherwise). All panel data models include cross-section fixed effects. White heteroskedasticity-consistent standard errors are used in the computation of *t*-statistics. (***), (**) and (*) indicate that the variable is statistically significant at the 1%, 5% and 10% level, respectively.

firms and matched firms. That difference is regressed on a constant and on the binary variable *BAN* (which assumes the value of one in the period of the ban and zero otherwise) and cross-section fixed effects. The use of this alternative approach leads to almost virtually identical conclusions (see *Table 6*, Panel B). The exception is the regression of percentage bid-ask spreads on *BAN* and cross-section fixed effects, where *BAN* loses its explanatory power. Nevertheless, the results hold for the absolute bid-ask spread. Overall, the ban seems to have produced a slight increase of CDS transaction costs along with a reduction in CDS volatility and tail risk.

4.4 The Effect of the Ban on Price Discovery

This subsection explores the cross-predictability between CDS and stock returns after the ban was imposed. If informed traders migrate to CDS markets after a ban and use their superior knowledge to purchase credit protection or revise quotes, CDS rates will convey more private information than before and their capacity to anticipate future stock returns should increase.

To assess the interaction between CDS and stock returns, the approach developed by Acharya and Johnson (2007) is followed. Their approach rests on two hypotheses: (i) the stock market is efficient with respect to the assimilation of publicly available information and (ii) the information flow from the CDS market to the stock market permanently impacts stock prices. Their method consists of estimating "CDS innovations" in the first phase. These innovations reflect new information or noise embedded in CDS returns that are explained neither by concurrent stock returns nor by past stock and CDS returns. CDS innovations are obtained by running time-series regressions for each firm:

$$\begin{aligned}\Delta(\ln CDSrate)_{i,t} &= \alpha_i + \sum_{k=0}^{5}\left(\beta_{i,k} + \gamma_{i,k} \times \frac{1}{CDSrate_{i,t}}\right) \times Stockret_{i,t-k} + \\ &+ \sum_{k=1}^{5}\left(\delta_{i,k} \times \Delta(\ln CDSrate)_{i,t-k}\right) + u_{i,t}\end{aligned} \tag{2}$$

The interaction between the stock returns and the inverse of the CDS rates is included to account for the nonlinear dependence between the CDS rate changes and the stock returns (Acharya and Johnson, 2007). The standardized residuals $\hat{u}_{i,t}$ capture "CDS innovations" for firm *i* on day *t* and are a measure of the arrival of new information in the CDS market, which, at that time, is neither known to stock market investors nor captured by past stock and CDS returns. Next, the impact of these innovations on stock returns is measured through the following regression:[18]

$$Stockret_{i,t} = a + b_1 \times StockRet_{i,t-1} + \sum_{k=1}^{2} c_k \times CDSinnovation_{i,t-k} + \varepsilon_{i,t} \tag{3}$$

The ability of innovations to predict stock returns implies that informed traders utilize the CDS market to exploit private information. Here, the interest is to verify whether the predictive ability of innovations increased after introduction

[18] The conclusions remain unchanged when more than two lags of CDS innovations are included in the estimation. Higher lags are not statistically significant.

Table 7 Cross-Predictability of Stock Returns and the Ban

Panel A

	(1)	(2)	(3)
Stock ret. (*t*-1)	0.020 (1.29)	0.020 (1.31)	0.020 (1.33)
CDS innov. (*t*-1)	-0.009 (-0.43)	-0.001 (-0.05)	0.000 (0.00)
CDS innov. (*t*-2)	-0.055** (-2.60)	(-2.60)	-0.058*** (-2.84)
CDS innov. (*t*-1)**BAN*(*t*-1)	0.030 (1.35)		
CDS innov. (*t*-2)**BAN*(*t*-2)	-0.004 (-0.19)		
CDS positive innov. (*t*-1)**BAN*(*t*-1)		0.544 (0.92)	
CDS positive innov. (*t*-2)**BAN*(*t*-2)		0.184 (0.75)	
CDS large positive innov. (*t*-1)**BAN*(*t*-1)			0.007 (0.36)
CDS large positive innov. (*t*-2)**BAN*(*t*-2)			0.015 (1.07)
C	0.000 (-0.56)	0.000 (-0.58)	(-0.60)
R-squared	18.3%	18.2%	17.8%
N	26050	26050	26050

Notes: Panel A reports the results of Fama-MacBeth regressions of stock returns on a constant, two lags of CDS innovations, two lags of the interaction of CDS innovations and the binary variable *BAN* (which assumes the value of one in the period of the ban if the firm's stocks are subject to the short-selling restriction and zero otherwise) and lagged stock returns. As an alternative to the introduction of the interaction of CDS innovations and the binary variable *BAN*, the interaction of positive CDS innovations and the binary variable *BAN* and the interaction of large positive CDS innovations and the binary variable *BAN* are also included in the regressions (columns 2 and 3). The equations are estimated for the span ranging from 4 January 2010 to 9 February 2012. *T*-statistics appear in parentheses. (***), (**) and (*) indicate that the variable is statistically significant at the 1%, 5% and 10% level, respectively.

of the ban for treatment-group firms. Hence, the interaction of *BAN* (which assumes the value of one in the period of the ban for treatment-group firms and zero otherwise) with CDS innovations is added to (3):

$$Stockret_{i,t} = a + b_1 \times StockRet_{i,t-1} + \\ + \sum_{k=1}^{2}(c_k + d_k \times BAN_{t-k}) \times CDSinnovation_{i,t-k} + \varepsilon_{i,t} \tag{4}$$

The above equation is estimated by means of Fama-MacBeth regressions for the period from 3 January 2010 to 10 February 2012. The results presented in *Table 7,* Panel A, show that past CDS innovations help predict stock returns. In effect, the second lag of $CDSinnovation$ presents a negative coefficient and is statistically significant. Nevertheless, lags of $CDSinnovation \times BAN$ are not statistically significant in line with the notion that cross-predictability did not increase after introduction of the ban for treatment group firms.

Table 7 Cross-Predictability of Stock Returns and the Ban

Panel B

	Before the ban			After the ban		
	ALL	**BAN**	**NOT BAN**	**ALL**	**BAN**	**NOT BAN**
Theta	-0.054** (-2.62)	-0.074** (-2.72)	-0.044 (-1.57)	-0.012 (-0.19)	-0.004 (-0.05)	-0.016 (-0.19)
Theta*	-2.63**	-2.16**	-1.58	-1.20	-0.87	-0.80
Theta (Matched Samples)	-0.054** (-2.27)	-0.074** (-2.72)	-0.035 (-0.97)	-0.010 (-0.14)	-0.004 (-0.05)	-0.015 (-0.15)
Theta* (Matched Samples)	-2.19**	-2.16**	-0.67	-0.99	-0.87	0.65

Notes: Panel B tabulates results of time series regressions of stock returns on a constant, two lags of CDS innovations and lagged stock returns. Here, the sample period is broken into two parts: the period that precedes the ban (3 January 2010 to 11 August 2011) and the period that follows the ban (12 August 2011 to 10 February 2012). To save space, only the cumulative effect of the two lags of CDS innovations is reported. Theta estimates for each time series regression $\left(\hat{\theta} = \sum_{k=1}^{2} \hat{c}_k\right)$ are averaged for the full sample of firms and by group. *t*-statistics appear in parentheses. Theta* denotes the average value of the standardized coefficient and is also computed for the entire sample of firms and by group. (***), (**) and (*) indicate that the variable is statistically significant at the 1%, 5% and 10% level, respectively.

The introduction of the ban is expected to induce informed trading based on negative information in CDS markets. As such, the predictive capacity of positive innovations should increase *vis-à-vis* negative innovations. Bearing that in mind, lags of the interaction of BAN with positive CDS innovations are also added to (3). The second column of *Table 7*, Panel A, details the results of the regression and shows that neither of the lags of the interaction of *BAN* with positive CDS innovations is statistically significant. A similar conclusion is achieved with the introduction of the lags of the interaction of *BAN* with large positive CDS innovations (higher than one standard deviation).

As an alternative to Fama-MacBeth regressions, time series regressions of (3) are run for each firm using two alternative subsamples: the period that precedes the ban (3 January 2010 to 11 August 2011) and the period that follows the ban (12 August 2011 to 10 February 2012). After that, the estimated coefficients are averaged for the entire sample of firms and for the two groups separately. *Table 7*, Panel B, reports the average $\hat{\theta} = \sum_{k=1}^{2} \hat{c}_k$, i.e. the cumulative effect of the two lags of CDS innovations, together with the corresponding *t*-stat for different group/period combinations. It is curious that past CDS innovations predict stock returns in the period prior to the ban when the full sample is considered ($\hat{\theta}$ negative and statistically significant), but that predictive capacity disappears after introduction of the ban. Indeed, the reduction in the predictive capacity is more striking for firms subject to the ban. The latter procedure is replicated using a matched-sample approach. $\hat{\theta}$ is not statistically significant in either of the groups in the period that follows the ban, which is consistent with previous results. On balance, the results do not support the hypothesis that the information flow between CDS and stock markets increased with the introduction of the ban. Consequently, H4 is rejected.

5. Conclusions

As long as the cost of CDS trading is not prohibitively expensive, the possibility of buying CDS protection opens an indirect channel through which short-selling restrictions might be circumvented. As CDSs are insurance against the default of a firm, they are naturally sensitive to negative information. Additionally, they allow investors to leverage their positions when trading. This paper examines the existence of a migration of informed investors to the CDS markets with the goal of circumventing short-selling restrictions imposed by financial regulators after a specific event—the August 2011 ban on covered short selling of European financial stocks. In doing so, this paper investigates whether that decision produced effects on the trading behavior of CDS contracts, namely on open interest, prices, volatility, liquidity and price discovery.

The results obtained do not support the view that investors saw credit protection buying as a viable substitute for short interest in stocks. First, it follows that the CDS open interest of firms included in the ban lists actually declined after introduction of the ban. In parallel, there is no evidence of price pressure on CDSs written on firms on the ban lists resulting from potential order imbalance, and CDS volatility actually declined. Finally, the ability of CDS spreads to predict stock prices fell during the ban, which is inconsistent with a surge in the flow of information from CDS to stock markets after that event in the wake of a migration of informed traders to the CDS markets.

Overall, the results of this paper contrast with those of Ni and Pan (2011) and Courtney (2010). Ni and Pan's (2011) findings suggest that following the 2008 SEC ban, it took more time for the negative information contained in either the option market or the CDS market to be incorporated into stock prices. The analysis carried out by Courtney (2010) indicates that the CDS prices of entities covered by the SEC ban rose at a higher rate during the period of the ban than did the prices of entities that were not covered. I did not reach a conclusion similar to those of the aforementioned studies for European entities following the 2011 ban.

There are several explanations for the inexistence of a migration of informed investors to the CDS market. The first is that transaction costs are too high in the CDS market. On average, the daily volatility of stock returns represents 314% of the stock percentage bid-ask spreads, whereas the daily volatility of CDS rate changes represents 58% of CDS percentage bid-ask spreads. Therefore, large movements of CDS spreads (almost two daily standard deviations) will be required in order to profit from a speculative investment in the CDS market. This justification is consistent with the finding of Hilscher *et al.* (2015) that informed traders are active primarily in the equity market rather than the CDS market given the existence of higher transaction costs in the CDS market. Duarte *et al.* (2007) also claim that only a small fraction of the mispricing between CDS and stock prices can be arbitraged due to the large transaction costs in the CDS markets. Even if short selling is restricted, informed investors may opt to stay out of the CDS market, as the value of their private information may not compensate for the incurrence of large CDS bid-ask spreads. Interestingly, the results of this study tell us that bid-ask spreads increased after introduction of the ban, primarily for firms affected by the ban. Certainly, these results are in accordance with the notion that any attempt on the part of informed traders to move to the CDS market may have prompted an adjustment of bid-ask spreads by dealers to compensate for the increment of information asymmetry in that market.

The second hypothesis rests on the concept of market segmentation. Goldstein *et al.* (2014) argue that financial markets are populated by different types of traders who have different trading opportunities. While some markets are used for speculation, others are used for hedging purposes. Therefore, it may be that some informed investors are limited to trading in the spot market in order to speculate. In reality, some investors may be prohibited from trading in the CDS market (e.g. mutual funds facing regulatory restrictions) or they are simply not specialized in or do not have access to the CDS market, where only major players are allowed to trade. In that regard, Duarte *et al.* (2007) argue that arbitrage trading between stock and CDS markets requires a high level of "intellectual capital" to identify the arbitrage opportunities and to hedge out the risks using complex models.

The third hypothesis relates to the existence of limits to arbitrage and synchronization risk that may affect CDS and equity market integration. This hypothesis is in line with the finding of Kapadia and Pu (2012) that arbitrage capital is slow moving and that CDS mispricing may persist in the short run but tends to decrease in the long run. In this view, mispricing of CDSs may exist in the short-run, which elevates the risk for investors betting on a certain view on the fundamentals of a firm. This additional risk rests on the timing of CDS price correction. Counterparty risk may also fuel this mispricing by reducing liquidity in the market.

Finally, the last explanation concerns sell-side elasticity. Foley-Fisher (2010) claims that only a fixed proportion of all investors is allowed to sell CDS protection. This friction is amplified in periods of crisis, as the distribution of beliefs is such that there is excess demand for CDS that cannot be satisfied without that friction playing a role. Additionally, in the presence of only a few sellers, informed traders may find it difficult to invest without revealing their strategies.

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