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Keywords: Short sale ban; jump risk; risk-neutral density; implied volatility skew; contagion risk

JEL Classification: G01, G14, G28

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The 2011 European Short Sale Ban: An Option Market Perspective*

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Abstract

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1. Introduction

On August 11, 2011, Belgium, France, Italy, and Spain imposed short sale bans. The European Securities and Markets Authority (ESMA) stated that the reason for the short sale bans was to curb market abuse and the spread of false rumors[1]. The spread of false rumors is dangerous because it may increase the risk of contagion, thereby endangering financial stability, which ultimately is the main condition that regulators intend to protect[2].

Recent academic studies argue that short sale bans, at best, do not impact stock price levels and, at worst, contribute to their decline and negatively impact market quality. For instance, Boehmer et al. (2013) conclude that it is unclear whether the 2008 SEC's imposition of short sale bans achieved the goal of providing a floor for US equity markets. Beber and Pagano (2013) investigate the impact of the 2008 bans on stock markets in 30 different countries and find that stocks that were banned underperform stocks not included in the bans.

Our paper focuses on the recent European short sale bans, whereas most existing work analyzes prior US bans. Thus, by comparing our European results to the US ones, we further increase the understanding of short sale bans' effects. Further, we explicitly take an option markets' perspective, as opposed to employing only the stock market itself. The argument is that the option market contains much information that is difficult, if not impossible, to extract from the stock market. Our paper focuses on changes in beliefs and expectations, like in Yan (2011), and in Chang et al. (2013). Forward-looking probabilities implied by options prices, i.e., risk neutral densities (RND), and the implied volatility (IV)

[1] ESMA stated on August 11, 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. [...] Today some authorities have decided to impose or extend existing short-selling bans in their respective countries. They have done so either to restrict the benefits that can be achieved from spreading false rumors or to achieve a regulatory level playing field, given the close inter-linkage between some EU markets."

[2] The ESMA's mission is to enhance the protection of investors and reinforce stable and well functioning financial markets in the European Union (available at http://www.esma.europa.eu/). Article 27 of Regulation No. 1095/2010 of the European Parliament and of the Council of 24 November 2010, establishes that the supervisory authority "shall contribute to developing methods for the resolution of failing key financial market participants in ways which avoid contagion, allow them to be wound down in an orderly and timely manner" (available at http://www.esma.europa.eu/system/files/Reg_716_2010_ESMA. pdf). ESMA produces a quarterly Risk Dashboard in which one of the four risks monitored is contagion risk.

skew are used to assess how the ban affects expected jump risk on banned and unbanned stocks. We focus not only on the outmost tails of RNDs but also on the ones of realized returns. We argue that it is the more extreme parts of the distributions that reflect best jump risk. To this purpose, we use Extreme Value Theory (EVT), which is especially suited for the analysis of tail probabilities. This approach allows us to assess how investors, through their perception of jump risk, differentiated between banned and unbanned stocks upon the introduction of the 2011 European short-selling ban.

We employ a dataset of daily IV across a range of different moneyness levels for all optionable European stocks listed in Belgium, France, Italy, and Spain. We use EVT for such assessment as this is a non-parametric method and, hence, independent of a specific distribution, as RNDs are. Our work is related to Melick and Thomas (1997) and Birru and Figlewski (2011), since it examines the behavior of RNDs over specific events. The rationale of using RND and IV skews to assess how the ban affected jump risk is supported by Bates (2000) and Rubinstein (1994). They show that before the 1987-crash, the probability of large negative stock returns was small and fairly close to that suggested by the normal distribution. Just prior to the crash however, the option-implied probability of jumps rose considerably at the same time that the IV skew became steeper. The left tail of the RND of returns became considerably fatter and thus negatively skewed with increased kurtosis, a phenomenon attributed to crash fear by Rubinstein (1994). We refer to, e.g., Pan (2002) and Eraker et al. (2003) for a discussion on how derivative pricing models are enhanced by incorporating jumps. As a result, out-of-the money (OTM) puts are systematically priced at a higher level relative to at-the-money (ATM) ones. Accordingly, a fat left tail in the RND of returns is a corollary of the fact that IV skew is steep, see also Bakshi et al. (2003).

The link between IV skews and jump risk is further strengthened by the fact that portfolio insurance often uses OTM puts to hedge equity portfolios, see Bates (2000) and Bakshi et al. (2003). OTM puts, by embedding more leverage, have greater downside hedging capacity than ATM puts and, hence, are considered the most explicit protection against jumps, see Bates (2000). Steep IV skews thus indicate the willingness of investors to pay for additional leverage to hedge such risk. The increase in the price of OTM puts since the 1987 crash could thus be attributed to a higher demand for portfolio insurance, which confirms a shift in beliefs of market participants who buy put options to hedge against negative jump risk.

The main contributions of our paper are threefold. First, using option market data, we provide evidence that the ban increased jump risk levels, thereby impacting especially the banned financial stocks. Jump risk may be proxied by measures based on IV skew, see, e.g., Pan (2002) and Yan (2011), who conclude that jump risk constitutes a priced risk premium in the cross-section of stocks. We compute the IV skew as the difference between the IV of OTM and ATM puts on individual shares. Unreported calculations show that our results are however robust to the use of IV from close to ATM options. Further, we also estimate extreme downside risk priced in RND tails to evaluate jump risk. While IV skews provide high frequency insight on jump risk changes, RND tails quantify the magnitude of expected jumps and their probabilities.

We show that it is the imposition of the ban itself that led to the increase in jump risk, rather than other causes such as information flow, options trading volumes, or stock specific factors. This finding is important because increased jump risk may provoke financial contagion and thus increase systemic risk. Financial contagion occurs when a relatively contained shock, which initially affects only one or few institutions, sectors or countries, propagates via larger shocks to the rest of the financial sector, economy or other countries. We refer to Bae et al. (2003), who use the tails of return distributions to model contagion. They argue that contagion occurs because of a sharp increase in the responses of one market, i.e., a jump, to a shock elsewhere. In an excellent paper, Ait-Sahalia et al. (2013) propose a model based on Hawkes jump-diffusion processes that mimic the two basic features of financial contagion: jump clustering in time, i.e., shock propagation in time, and across assets, i.e., shock propagation in space. In Hawkes processes, the probability of a next jump increases after a first jump is observed, thus naturally linking jump risk to contagion risk. Because of the connection between jump risk and contagion, shifts in jump risk are closely monitored by regulators[3].

[3] Poon and Granger (2003) discuss that the Bank of England uses implied volatilities to assess market sentiment. Further, Gai and Vause (2006) propose a method for measuring investor risk appetite from implied volatilities.

Second, using historical stock market returns, we find that after announcement of the ban contagion risk actually drops for banned stocks. This finding seems to run contrary to what one might expect, given the documented increases in jump risk levels for banned stocks. Interestingly, for the unbanned stocks we document that contagion risk levels do indeed increase after the ban, thus behaving in line with the rise in jump risk levels. Apparently, the ban affects banned stocks differently from the other stocks. We argue that this difference may be caused by (formal and informal) market makers' reluctance or prohibition to further increase their options' inventory risk, leading to relatively steep IV skews, reduced volumes, and widened bid-ask spreads for banned stocks. In support of this argument, Boehmer et al. (2013) find for large caps during the 2008 ban much poorer conditions for price impact, realized bid-ask spreads, and intra-day volatility. They argue that liquidity dried up because high-frequency traders, acting as informal market makers, were not exempt from the ban, as formal market makers were. Also, Beber and Pagano (2013) note that the ban led to an increase in bid-ask spreads due to market participants' inability to sell short, making inventory management more difficult for market makers. Battalio and Schultz (2011) show that during the 2008 US ban bid-ask spreads in options markets increased significantly more for options on banned stocks than for options on other stocks. Bid-ask spreads widen because market makers, formerly hedged, now try to protect themselves by keeping their inventories unchanged.

Third, we compare the effects of the 2011 European ban to its American 2008 counterpart. Investors may be able to obtain economic short exposure to banned stocks through a derivatives-based strategy that replicates the payoff of a stock's short sale. Such a "substitution effect" is characterized by a migration of trading volume from one instrument to another. We find that no substitution effect occurred between regular short selling and synthetic shorting through single stock puts during the 2011 European short selling ban. This result ties in with conclusions on the 2008 US ban reached by Battalio and Schultz (2011) and Grundy et al. (2012). They find that even if regulators do not explicitly prohibit net short positions in options, the ban by itself is effective in curbing these flows. They argue that substitution does not occur because bid-ask spreads increased significantly, whereas volume decreased, especially for derivatives on banned stocks. They suggest that the reason for the light volumes is the higher trading cost.

Instead of a substitution effect, our results suggest that a migration out of single stock puts into the EuroStoxx 50 index option market seems to have occurred. We hypothesize that this type of migration diversifies selling pressure initially concentrated in financial stocks across a larger share of the stock market, thus reducing systemic risks.

The remainder of this paper is organized as follows. Section 2 discusses the 2011 European short-selling ban on financial stocks. Section 3 describes the data and methodology. Section 4 presents our empirical findings and section 5 concludes.

2. The 2011 European Short sale Ban

The 2011 short sale ban on financial stocks in the Euro member countries Belgium, France, Italy, and Spain was established by a coordinated act of the European Securities and Market Authority (ESMA) and the national financial market regulators of these countries on August 11, 2011. The announcement was made via a public statement issued by the ESMA (reference ESMA/2011/266) and was followed by publications the same day by the Belgian Financial Services and Markets Authority (FSMA), the French Autorité Des Marchés Financiers (AMF), the Italian Commissione Nazionale per le Societá e la Borsa (Consob), and the Spanish Comision Nacional Del Mercado de Valores (CNMV). The ban entered into effect on August 12. Table 1 provides the list of banned financial stocks.

< *Please insert Table 1 about here* >

The ban on covered short selling not only prohibited the creation of new net short positions but also banned increases in existing ones, including intra-day operations. Naked short selling had already been prohibited in these European markets since 2008. Positions arising from formal market making activities were exempted from the ban. The ban targeted not only public markets but also OTC markets. In terms of scope, the announcements differed. The FSMA announced that the ban applied to net economic short positions of any kind, while the AMF communicated that derivatives could only be used to hedge, create or extend net long positions. For the Consob, the ban covered only shares and not ETFs or any derivatives, while the CNMV imposed the ban on all trades in equities or indices.

During the ban, holders of financial stocks could still use single stock derivatives or simply sell their holdings to hedge their portfolios. Investors exposed to stocks were allowed to hedge their overall equity market exposure by trading index or single stock derivatives. What was prohibited was the short selling of banned stocks, not hedging them or reducing equity market risk. The creation or extension of marginal net short positions in banned securities as a result of hedging equity market risk was still allowed.

The August 11, 2011 European short sale ban was initially intended to be in place for the next 15 days only, with the exception of Belgium, which announced that the ban would remain in effect indefinitely. Nevertheless, the ban was extended by the Spanish CNMV, the French AMF, and the Italian Consob several times. On February 13, 2012, both the FSMA and AMF announced the lifting of the ban with immediate effect in Belgium and with retroactive effect, on February 11, 2012, in France. On February 15, 2012, the CNMV announced the lifting of the Spanish ban from February 16 onwards. Finally, on February 24, 2012, the Italian ban expired. New short sale bans were re-imposed by Spain and Italy later in 2012.

3. Data and Methodology

Our analysis of jump risk for European stocks uses the RND of returns from single stock options. We implement Figleswki's (2009) method for obtaining RNDs. He builds on the Breeden and Litzenberger (1978) formulae and interpolates and smooths the IV structure, instead of interpolating option prices. A clear strength of the Figleswki method is its ability to fill in intermediate grid values of the IV curve between the available strikes (the body of the RND) with reduced noise and to extrapolate the RND beyond such observable strikes[4] with tails of flexible and reasonable shape.

To calculate RND for the stocks of our interest, we obtain daily IV data for seven moneyness levels, i.e., 80, 90, 95, 100, 105, 110, and 120, at the three-month maturity. Implied volatilities are extracted by reverse engineering the Black-Scholes model[5] from Bloomberg 16:00 hours closing mid-prices (Bloomberg, 2008). Our sample covers the period from February 15, 2008 to March 27, 2012 and includes 1,073 trading days. It consists of all stocks that had listed options as of February 2012 on the Belgian (Brussels Stock Exchange/Euronext Brussels), French (Paris Bourse or Euronext Paris), Italian (Milan Stock Exchange or Borsa Italiana), and Spanish (Bolsa de Madrid) stock exchanges. Overall, our sample comprises 185 stocks, of which 105 are included in these stock exchanges' main indices, i.e., the Belgian BEL20, the French CAC40, the Italian MID, and the Spanish IBEX35.

In line with Figleswki (2009), IVs for the 80 to 95 moneyness levels are obtained from puts, while IVs for the 105 to 120 moneyness levels are obtained from calls. For consistency with our IV skew measure, we use ATM IV from puts. Because we intend to compare RND from banned stocks to unbanned ones, we compute IV skews and extract RNDs for these two groups of stocks separately. The banned group constituents are the stocks that were prohibited from short sales. Our unbanned group constituents are the remaining stocks in our sample. We compute IVs for banned and unbanned stocks separately by equally averaging IV on each moneyness level available across all stock belonging to either the banned group or the unbanned one. This step produces one IV structure across our seven moneyness levels for both groups for every day in our sample. Then, we apply the Black-Scholes model to our IV data to obtain options prices (C) for the banned and unbanned groups of stocks. We set the instantaneous price level of both groups (S_0) equal to 100, and as a result the percentage moneyness level automatically reflects strike prices per group. When applying the Black-Scholes model, we calculate contemporaneous

[4] Interpolation using fourth order splines is argued by Figleswki (2009) to be superior to cubic splines, because it avoids kinks in the RND. The translation from interpolated IV curve into RND would require taking higher order derivatives than used by the construction of the spline.

[5] IV is calculated assuming constant interest rates and discrete dividends. Interpolation is used to calculate the IV at a fixed level of moneyness and at a fixed time to maturity.

dividend yield for banned and unbanned stocks by equally weighting dividend yield from the individual stocks. The risk-free rate applied is the Euribor three-months maturity.

Once options prices for the average banned and unbanned stocks are obtained, we can extract the RND of equity returns using the Breeden and Litzenberger (1978) formulae for the strikes along the body of our distribution, i.e., from 80 to 120 moneyness levels:

$$RND(S) = \exp(rT) \frac{\delta^2 C(T,K)}{\delta K^2}\bigg|_{K=S}, \tag{1}$$

which by computing the second derivative of the option price relative to strike prices via central differences leads to:

$$RND(S) \approx \exp(rT) \frac{C(T,S-\Delta K) - 2C(T,S) + C(T,S+\Delta K)}{(\Delta K)^2}, \tag{2}$$

where *RND(S)* is the risk-neutral probability of observing the terminal index level (*S*) at time *T*, *r* is the risk–free rate for the specific maturity, *K* is the strike price, and *C* is the index option price.

Following Figlewski (2009), extrapolation beyond the body of the RND[6] is done by fitting a Generalized Extreme Value (GEV) distribution using two extreme anchor points on each side of the body of the RND and extending a tail with the same shape. The GEV-based extrapolation is then used to model the tails of the RND towards the moneyness levels 0 and 200. We initially use the first and third percentiles of the RND's body as (outer and inner) anchor points for the left tail and the 99th and 97th percentiles as (outer and inner) anchor points for the right tail. Differently from Figlewski (2009), these anchor points are allowed to change if the fitted GEV curves produce implausible tails, e.g., zero probability under the tails. Equations (3) and (4) give, respectively, the GEV's cumulative distribution function and probability distribution function:

[6] The Figlewski (2009) method is close to the one by Bliss and Panigirtzoglou (2004), where body and tails are also extracted separately. These authors use a weighed natural spline algorithm for interpolation, which has the same decreasing noise effect in RNDs. Extrapolation is done by the introduction of pseudo-data points, which has the effect of pasting lognormal tails into the RND. One advantage of these two approaches is that extrapolation does not result in negative probabilities, which is possible when the spline interpolation is applied. We favor Figlewski's (2009) approach because the use of the lognormal tails by Bliss and Panigirtzoglou (2004) assumes that IV is constant beyond the observable strikes, resembling the Black-Scholes model.

$$F_{GEV}(S_T) = \exp\left[-\left(1+\omega\left(\frac{S_T-\mu}{\sigma}\right)\right)^{-1/\omega}\right], \tag{3}$$

$$f_{GEV}(S_T) = \frac{1}{\sigma}\left[1+\omega\left(\frac{S_T-\mu}{\sigma}\right)\right]^{\left(-\frac{1}{\omega}\right)-1} \exp\left[-\left(1+\omega\left(\frac{S_T-\mu}{\sigma}\right)\right)^{-1/\omega}\right], \tag{4}$$

where $\omega > 0$ sets a fat tail relative to the normal, $\omega = 0$ sets a tail such as the normal, and $\omega < 0$ sets a distribution tail that is thinner than the normal. The μ and σ are location and dispersion parameters. Because fitting GEV curves entails setting these three parameters, Figlewski (2009) also imposes three conditions on the tail: i) that the total probability in the tail of the body (up to the inner anchor point) is the same for the RND and the GEV approximation; ii) that the shape of the RND equals the shape of the GEV curve in the inner anchor point, and iii) in the outer anchor point[7].

Once the body and tails of the RND for terminal index levels are obtained for banned and unbanned stocks, we convert them into return RNDs by calculating log-returns relative to the starting index level S_0. Finally, we compute probabilities for every percentage return quantile of the PDF via linear interpolation, which are normalized to integrate to one.

We then compare tails of RNDs for banned and unbanned stocks by applying EVT. EVT allows us to compare returns implied by the extreme quantiles of our two RNDs, in other words, the extreme downside risk or Value-at-Risk (VaR) implied by RNDs for banned and unbanned stocks. EVT is appropriate to measure jump risk as it focuses on tail events, such as jumps in return distribution, see Hartmann et al. (2004). We first estimate the tail shape estimator (φ), using Hill (1975):

$$\hat{\varphi} = \frac{1}{k}\sum_{j=1}^{k} ln\left(\frac{x_j}{x_{k+1}}\right), \tag{5}$$

where x_j are ranked returns on ascending order j = 1, …, n, n is the sample size, k is the number of extreme returns used in the tail estimation, and $x_{k+1}$is the return 'tail cut-off point'. The tail shape estimator φ measures the curvature, i.e., the fatness of the tails of the return distribution: a high (low) φ

[7] See Appendix 1 for details on the fitting GEV algorithm as well as precise conditions imposed.

indicates that the tail is fat (thin). Once φ is obtained, we compute extreme downside risk, hereafter VaR, using a semi-parametric quantile estimator used in Hartmann et al. (2004):

$$\hat{q}_p = x_{k+1}\left(\frac{k}{pn}\right)^{\hat{\varphi}}, \quad (6)$$

where n is the sample size, p is a chosen exceedance probability, which means the likelihood that a return x_j exceeds the tail value q, and $x_{k+1}$is the 'tail cut-off point'. Note that $\hat{q}_p$ has as one of its inputs the estimated tail shape parameter $\hat{\varphi}$. The $\hat{q}_p$ statistic indicates the level of the worst return occurring with probability p.

Because the tail quantile statistic $\frac{\sqrt{k}}{ln\left(\frac{k}{pk}\right)}\left[ln\frac{\hat{q}\ (p)}{q\ (p)}\right]$ is asymptotically normally distributed, like in Hartmann et al. (2004) we use the following T-statistic for this estimator:

$$T_q = \frac{\hat{q}_1-\hat{q}_2}{\sigma[\hat{q}_1-\hat{q}_2]} \quad \sim N(0,1), \quad (7)$$

where the denominator is calculated as the bootstrapped difference between the estimated VaR for p = 1 and p = 2 using 1,000 bootstraps. The null hypothesis of this test is that $\hat{q}_1$ and $\hat{q}_2$ do not come from independent samples of normal distributions, therefore, VaR for p = 1 and p = 2 are equal across samples.

In the next step, we employ a bivariate EVT method to calculate commonality in jumps, hence, contagion risk from historical returns. EVT is well suited to measure contagion risk because it does not assume any specific return distribution. Our approach estimates how likely it is that one stock experiences a crash beyond a specific extreme negative return threshold conditional on another stock crash beyond a equally probable threshold. We refer to Hartmann et al. (2004) who use the so-called conditional co-crash (CCC) probability estimator, which is applied to each pair of stocks in our sample, as follows:

$$\widehat{CCC}_{ij} = 2 - \frac{1}{k}\sum_{t=1}^{N} I\{V_{it} > x_{i,N-k}\ or\ V_{jt} > x_{j,N-k}\ \}, \quad (8)$$

where the function I is the crash indicator function, in which I = 1 in case of a crash, and otherwise, I = 0; V_{it} and V_{jt} are returns for stocks i and j at time t; $x_{i,N-k}$and $x_{j,N-k}$ are extreme crash thresholds. Further, because the CCC-probability is asymptotic normal if $k/N \to 0\ as\ k, N \to \infty$ (see Hartmann, 2004), a t-test for such estimator is obtained by the same bootstrap-based approach that is used in Equation (7).

We use the daily IV skews of individual European equities as a second measure of jump risk. The IV skew is calculated as the difference between the IV of three-month OTM listed puts at the 80 percent moneyness level and ATM puts with the same maturity for every stock in our sample. As with RNDs, we construct two indices of IV skew, one for banned and the other for unbanned stocks, by equally averaging the stock specific IV skews of the constituents of each index. Next, we construct a daily excess IV skew measure by subtracting the unbanned IV skew index from the banned IV skew index. We call this new index the B-U index or simply the excess skew. We also calculate single country versions of the banned, unbanned, and B-U index for Belgium, France, Italy, and Spain, respectively.

Table 2 presents descriptive statistics for our IV skews and B-U index for the entire sample period. They are provided for the overall and single country levels and also for the daily differences of excess skew. Panel A shows that the average and median IV skews for banned stocks are higher than for unbanned stocks, an observation that pertains not only to the overall numbers but also to each country individually. The standard deviation of the IV skew is mostly higher for banned stocks. The distributions of the IV skew are as expected all positively skewed as skew tends to be bounded at zero. Most distributions of the IV skew are leptokurtic. As a consequence of these characteristics, all IV skew distributions reported here have fat right tails and are not normal, according to their Jarque-Bera statistics.

< Please insert Table 2 about here >

Panel B reports statistics, from the B-U index, for the excess skew. All average and median excess skews are positive, while all average and median daily differences in the excess skew, measured in volatility points, are close to zero. The highest median daily difference is 0.016 volatility points for Belgium, and the highest daily difference average is 0.004 volatility points for France. The distributions of the excess skew are leptokurtic, except for France, and most are positively skewed. All distributions of levels and differences of the excess skew are not normal, according to their Jarque-Bera statistics. This is the reason why we use non-parametric Mann-Whitney U-test to make statistical inferences[8].

[8] More specifically, we use the Mann-Whitney U-test to determine whether the median IV skews or the median excess skews from different periods differ statistically from one another. The null hypothesis is that there is no difference between measures

Finally, from Bloomberg we source daily trading volumes and numbers of shares outstanding per stock, trading volumes, and put-call volume ratios for listed options. Trading volumes for listed puts on the EuroStoxx50 index, the V2X index (the IV index from EuroStoxx 50 index), and generic series of five-year sovereign credit default swaps (CDS) for Belgium, France, Italy, and Spain are also downloaded from Bloomberg. Daily short stock positions (utilization rates) and costs of short selling (simple average fee, simple average rebate, and cost of borrow score) were kindly provided by Markit Securities Finance, former Data Explorers.

4. Discussion of Results

We start our discussion of results by first examining the short selling utilization rate[9] and the performance of banned and unbanned stocks around the ban announcement day. Subsequently, we inspect VaR implied by RND, as well as IV skews, information flow, contagion risk, trading activity, and finally we perform a regression analysis.

Figure 1 shows that the imposition of the 2011 European ban strongly affects the short selling of stocks in Belgium, France, Italy, and Spain. Short selling levels, measured by utilization rates, fall for banned stocks from 32 percent to 27 percent in the months of August and September 2011, especially after the ban announcement on August 11, 2011. This drop in short selling utilization is widespread across the four countries. For Belgium and Italy, short selling utilization drops from 29 percent to 23 percent and 24 percent, respectively. For France, it remains unchanged at approximately ten percent during this period. For Spain it drops by 8 percent from 53 percent on ban announcement day to 45 percent per September 30, 2011. We find that such drops in the utilization rate come from the decrease in value of short selling,

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from two distinct periods. We use the normal approximation of the U statistic, given the large number of observations available. We transform the U-values into Z-values and compare these with Z-critical values at the one, five, and ten percent levels of significance.

[9] The short utilization rate is calculated as *Utilization*=100*(*ValueOnLoan*/*InventoryValue*), where *ValueOnLoan* is the beneficial owner value of the loan and *InventoryValue* is the beneficial owner inventory value. Utilization measures the value of a stock utilized for securities lending against the total value of inventory available for lending so that it indicates the short selling demand, which ranges from zero to 100 percent. Data Explorers Limited (2011) mention that their database captures "stock loan trading information from over 100 participants and approximately 85 percent of the OTC securities lending market".

the numerator of the utilization rate, as inventories of stocks available for lending in the four countries remain relatively unchanged. The decreasing utilization rates indicate that the ban was effective in reducing short selling, despite the fact that market makers were still allowed to short banned stocks.

< *Please insert Figure 1 about here* >

The reduction in short selling of financial stocks is especially noteworthy when utilization rates for banned and unbanned stocks are compared. Figure 1 shows that the utilization rate for unbanned stocks increases on average over all four countries from 16 percent to 18 percent during August and September 2011, an increase observed across all four Euro countries.

Despite such changes in utilization rates, short selling in banned financial stocks far exceeds the level measured in unbanned stocks during August and September 2011. In August, average short selling utilization for financial stocks is twice the level reported for stocks of other sectors (32 percent versus 16 percent, see Figure 1). This difference, which is also present in the single country data, indicates that by August 2011, short-sellers would benefit much more from further deterioration of financial stocks rather than from a potential weakness in the average stock. Despite such dichotomy, utilization rates for the overall market around the ban announcement day were at their highest levels since 2010 for the four countries. For Italy and Spain, the short selling activity was concentrated in mid-caps (see Data Explorers Limited, 2011), which matches a large short selling interest in Italian and Spanish banks, as most of these institutions are mid-caps.

From the end of June until the ban announcement, shares by European Banks (Euro Stoxx Banks index) dropped by around 32 percent, whereas the Euro Stoxx 50 index fell by 22 percent. In the first ten days of August 2011, before the ban, shares by European banks fell by 23 percent, whereas the European index corrected by 17 percent. In the subsequent month after the ban was announced, European banks' stock lost an additional 18 percent while non-financial equity dropped by only six percent. Our data on short selling positions and returns suggest that financial stocks were indeed under strong pressure.

4.1. Risk-Neutral Densities

After extracting RNDs for both banned and unbanned stocks, we next determine the optimal number of observations k used for estimation of parameter φ in Equation (5). For this purpose, we produce Hill-plots for the left tail of our two RNDs. Hill-plots depict the relationship between k and φ as a curve. The optimal value of k is selected as the minimum level for which value of φ stabilizes, thus where a stable trade-off between the approximation of the tail shape by the Pareto distribution and the uncertainty of such approximation occurs (because of the use of fewer observations, i.e., a lower k). We set k equal to four percent or 43 observations, which matches the level used previously in the literature (see, e.g., Hartmann et al. (2004)).

In the following analysis, where we analyze VaR-levels implied by RNDs, we distinguish five sub-periods: (1) the US recession period (February 15, 2008 to June 30, 2009); (2) the period of the 2009/2010 stock market rally (July 1, 2009 to April 26, 2010); (3) the European crisis period (April 27, 2010 to August 10, 2011), initiated by Standard and Poor's downgrade of Greece's sovereign bonds to "junk" status; (4) the ban period (August 11, 2011 to February 16, 2012); and (5) the post-ban period, from February 17, 2012, one day after the short-selling ban was lifted in Belgium, to March 27, 2012, when our data sample ends.

The VaRs implied by RNDs provide interesting insights into perceived jump risk. Panel A of Table 3 shows that during the ban-period the RND-implied VaR-levels for banned stocks are significantly higher than for unbanned stocks. The same conclusion holds for the post-ban period.

< Please insert Table 3 about here >

Next, we compare the VaR-levels across the different sub-samples. VaR-levels from RNDs during the ban period are significantly higher than during the preceding period, i.e., the pre-ban European crisis. The ten percent VaR for banned stocks it increases from 37 to 62 percent for banned stocks, whereas for unbanned stocks increases to a much lesser extent, from 35 to 46 percent. Similar differences in extreme downside risk for these two sub-samples are observed for the five and one percent VaR-levels. Interestingly, the VaR-levels for the post-ban periods are not statistically different from the ban period for both banned and unbanned stocks. The other sub-sample that had very distinct downside risk in comparison to the preceding one was the 2009 stock market rally. The latter period had statistically significant less VaR priced in RND-returns than the US recession period, especially for banned stocks but also for unbanned ones. The 10 percent VaR banned stocks was 50 percent during the recession and 40 percent during the rally, whereas for unbanned stocks they were 47 and 41 percent, respectively. Overall we conclude that the VaR-levels for banned stocks were generally higher than for unbanned stocks and downside-risk priced in RND reached its peak during the ban.

4.2 Volatility Skews

Figure 2 depicts the historical behavior of our proxy for jump risk, the average IV skew, for banned and unbanned stocks. The ban period is highlighted, with the beginning of the shadowed part representing the ban announcement day. We observe that jump risk rises strongly just prior to announcement of the ban for both banned and unbanned stocks. Such spikes in jump risk occur during the day of August 11, 2011, whereas the ban was officially announced only after the market close. The increase in the average IV skew on August 11, 2011, for banned stocks is equivalent to 2.16 volatility points, while for unbanned stocks it is equivalent to 1.05 volatility points. Both differences exceed the 99th percentile of all daily IV skew changes in our sample. On August 12, 2011, the average IV skew continued to rise sharply, by 0.78 volatility points for banned stocks, a movement exceeding the 94th percentile of all daily IV skew changes in our sample. On that same day, the IV skew for unbanned stocks rose by 0.55 volatility points, exceeding the 96th percentile. We observe that jumps in the IV skews around ban announcement day are clearly outliers in our sample. More importantly, the rise in the IV skew for banned stocks is much more pronounced than for unbanned stocks. The daily difference in the B-U index on August 11, 2011, is 1.11, exceeding the 97th percentile, and 0.23 on August 12, 2011, exceeding the 73rd percentile[10].

< Please insert Figure 2 about here >

[10] It is unclear whether any information on the upcoming short selling ban leaked before the market closed on August 11. However, given that a ban on covered short selling on all stocks was already introduced in Greece on August 8, 2011, the extension of the ban to other European countries might have been expected by some market participants.

We observe from Figure 2 that between 2008 and 2012, spikes in average IV skews well above their mean coincide with periods of market turmoil. Figure 2 depicts that the IV skews strongly rise in 2008 around the Lehman collapse and wane after the market trough in March 2009. In 2010, the IV skews jump on April 27, the day that Greek government bonds were downgraded by Standard and Poor's to "junk" status. The IV skew then strongly reverses on October 18, 2010, when a task force of European leaders agreed on a package to improve the European Union's economic governance in order to tackle the financial crisis. The 2011 jump in IV skews coincides with the ban announcement day on August 11, 2011. The announcement was not accompanied by any major event related to the European financial crisis, to equity markets in general or to the financial sector. We observe that on all these three occasions, the IV skew of banned stocks exceeded the one for unbanned stocks. This comes as no surprise, as these IV skew increases corresponded with financial sector crises.

Figure 2 also suggests that after the announcement of the short sale ban, the IV skew levels for both banned and unbanned stocks tended to hover at an elevated level for several weeks. During the entire ban, the IV skew of banned stocks remained relatively high, whereas the IV skew for unbanned stocks slowly declined to pre-ban levels. This persistence in the high level of jump risk indicates that the ban did not diminish market participants' concerns regarding European financial stocks. Table 4 presents the corresponding medians for the whole period and for the five sub-periods separately.

< *Please insert Table 4 about here* >

Table 4 presents the IV skew medians for the full sample and for the same five sub-periods previously defined. We observe that the sub-periods 1, 3, and 4 have the highest IV skews. They also roughly match the periods of market turmoil and volatility humps highlighted in Figure 2: the global financial crisis, the European sovereign debt crisis, and the 2011 European ban period. The median IV skew for banned stocks is 7.34, significantly higher during the ban period than before it, when it was 6.05. For unbanned stocks, the median IV skew during the ban period is 6.05 only slightly higher than during the European crisis, when it was 5.78. Moreover, the median IV skew during the European crisis period is also higher than during the preceding period, the period of the stock market rally. Medians for these IV skews are statistically different from each other at the one percent level, based on the Mann-Whitney U-test.

The median IV skews during the ban period, for both banned and unbanned stocks, are the highest among all other sub-periods in our sample. Furthermore, the median IV skew for banned stocks during the ban period rises much more than that for unbanned stocks. We observe similar patterns in the country-specific data. Such empirical findings suggest that the ban contributes to an increase in jump risk, especially with respect to banned stocks. Conversely, once the ban is lifted, IV skews drop. For banned and unbanned stocks, the post-ban period records IV skews that are statistically smaller than during the ban period.

Figure 2 also indicates that the IV skew for banned stocks exceeds that for unbanned stocks in most periods. Panel B shows that the B-U index has a positive median value of 0.71 volatility points across the entire sample. We observe similar results for the country-specific excess skew indices. More importantly, overall median excess skew is higher during the period of the short sale ban than during any other period, reaching 1.18 in comparison to 0.68 during the pre-ban European crisis and 1.15 during the post-ban period.

Our overall impression from the data examined thus far is that the short sale ban does not reduce jump risk during the European financial crisis. Otherwise, VaR levels and IV skews would have receded. However, we find significant evidence that VaR levels increased strongly and that IV skews jumped instantly when the short sale ban was introduced and remained high during the period of the ban, especially for banned stocks.

4.3 Information Flow

A potential flaw in the empirical analysis so far is that large movements in the IV skew, observed during the ban or at the time of its announcement, may have been contemporaneous to the dissemination of other relevant information. If so, one cannot draw a clear connection between the ban announcement and IV skew behavior. We refer to Beber and Pagano (2013) and Boehmer et al. (2013), who both analyze the 2008 US short sale ban and the concurrent US 2008 TARP announcement.

< *Please insert Figure 3 about here* >

Figure 3 indicates that the IV skews rise on August 11, 2011, the ban announcement day, even though no negative shocks occurred in our chosen information flow proxies, i.e., we do not observe negative shocks within country-specific CDS spreads and the V2X index[11]. Information flow for the four countries was relatively benign around the ban announcement date, as CDS spread levels remain unchanged, whereas the V2X even decreases after showing a large spike in the days preceding the ban. Equity market movements around that period further support the presence of such positive information flow[12]. The EuroStoxx 50 index rose by 2.86 percent on August 11 and by 4.15 percent on August 12, whereas the EuroStoxx Banks index rose by 2.96 percent and 5.26 percent, respectively. Moreover, no other major announcement was made during these days. The absence of negative information strongly suggests that the ban announcement itself catalyzed the rise in jump risk.

4.4. Contagion Risk

In this section, we assess the development of contagion risk, using conditional co-crash probabilities. Estimation of these probabilities requires the setting of k as the number of observations used in Equation (8). For consistency with our Hill-estimator, we again use k=43. We estimate the bivariate CCC-probabilities or contagion risk for all pairs of banned and all pairs of unbanned stocks using empirical returns. We note that these calculations are not based on options market data, nor are they forward-looking in nature. Contrary to previous sections, the calculations of the CCC-probabilities are based on realized, historical stock price returns. We compute the average CCC-probabilities for banned and unbanned stocks

[11] We use spreads on sovereign credit default swaps (CDS) to proxy the country-specific information flow. We adopt the V2X, the European counterpart of the VIX (the IV index for S&P500 index options), as a proxy for the European equity market information flow.

[12] Figure 3 shows that country-specific IV skew moves match very well the sovereign CDS spread behavior in this period. CDS spreads moved sideways for Belgium, France, and Italy, while those for Spain rose. This divergence can be explained by the fact that on February 13, 2012, Spain's sovereign debt rating was downgraded by Moody's by two notches, from A3 to A1, much more severely than the rating changes for the other three countries.

separately. The average CCC-probability measures the likelihood that a banned (unbanned) stock crashes given that another banned (unbanned) stock crashes. Results from the estimation of Equation (8) for the full sample and individual sub-samples are presented in Table 5.

< Please insert Table 5 about here >

The CCC-probability in the full sample for banned stocks is 32 percent, while for unbanned ones it is 29 percent. This difference in CCC-probability between the two groups is statistically not significant. In the first sub-sample periods, the contagion risk of the banned stocks reaches a similar level as it does for the other stocks. However, during the pre-ban European crisis period, we find that contagion risk for banned stocks, at 42 percent, becomes statistically different from unbanned stocks, at 33 percent, at the five percent level. Surprisingly, this statistical difference is no longer observed during the ban period, when the CCC-probability for banned stocks decreases to 32 percent, while it increases to 41 percent for unbanned stocks. This decrease in contagion risk for banned stocks is one of the major findings of our paper. Apparently, imposition of the ban decreased systemic risk, based on historical returns. This effect occurred despite the increase in forward-looking jump risk across the same sample and period.

In a next step we analyze whether CCC-probabilities for banned and unbanned stocks are different across samples. We observe that CCC-probabilities for banned stocks during the US recession (27 percent) and the 2009 equity market rally period (28 percent) are not statistically different. The same condition applies to unbanned stocks, for which CCC-probabilities are respectively 26 and 23 percent for these two periods. The pre-ban European crisis period, however, witnesses an abrupt and statistically significant increase in CCC-probability for banned and unbanned stocks. CCC-probabilities for banned stocks rises from 28 to 42 percent, whereas for unbanned stocks such increase is from 23 to 32 percent. Clearly, contagion risk is higher across the board once the European crisis is triggered but especially so for financial stocks. After the ban is announced, banned stocks' contagion risk falls from 42 to 32 percent, while for unbanned stocks, contagion risk rises from 32 to 41 percent. Obviously, the short-selling restrictions impacted banned stocks differently from unbanned stocks.

4.5. Trading Activity

In this section, we study an additional variable that may shed light on our analysis of jump and contagion risk: trading activity. Hypothesizing the existence of supply-demand imbalances in the options market, Bollen and Whaley (2004) suggest that the IV skew might be closely linked to trading activity in the options market. They argue that changes in the shape of the IV function are directly related to net buying pressure on options from end-users' public order flow. End-users trade options for portfolio insurance, agency, and speculative reasons, rather than for market making reasons. This conclusion is confirmed by Gârleanu et al. (2009), who find that the size of the IV skew is positively and significantly related to demand pressures. They show that IV skews are most impacted by institutional investors seeking portfolio insurance in the index option market but their results also apply to the bottom-up IV skew.

We inspect in the following daily put and call trading volumes as well as the put-call volume ratio as proxies for trading pressure, as suggested by Dennis and Mayhew (2002). We measure volume as the median number of contracts traded on a specific day for all stocks in the sample[13]. We obtain an overall put-call volume ratio by averaging the stock-specific contracts. Again, we evaluate these measures over the five periods previously identified in our dataset.

The results in Table 6 document that the median number of single stock puts for each banned stock traded per day decreases significantly from 1,905 during the pre-ban European crisis period to 1,727 during the ban period. For unbanned stocks, the median volume of puts also drops, from 1,157 during the pre-ban period to 943 during the ban. Similar results are obtained for the median number of options (puts and calls) traded during the ban period relative to the pre-ban period. The median put-call volume ratio for unbanned stocks significantly increases during the ban, from 6.3 to 8.7, whereas the median put-call volume ratio for banned stocks remains nearly the same.

< Please insert Table 6 about here >

[13] Evidence that OTM put options have the heaviest trading volume among puts (see Bollen and Whaley, 2004) supports our approach.

Our findings from Table 6 provide no evidence that individual stock options, in particular puts, affected by the ban experienced a large rise in trading activity. Thus, we find no evidence of a substitution effect of short selling of common stock into single stock put options. Nor, however, does trading activity completely dry up during the ban period. This result is in line with Grundy et al. (2012), who find that the overall volume of options trading dropped during the 2008 US short-selling ban. This behavior of trading volumes suggests that during the ban, the IV skew does not increase as a result of increased selling pressure, as originally suggested by Bollen and Whaley (2004) and Gârleanu et al. (2009). We next investigate the supply and demand dynamics of the single stock put options market.

We assume that once short selling activity in banned stocks diminishes, the demand for synthetic shorts via put options should increase. When analyzing the supply-side of the market, we first take the perspective of informal market makers in options. Such agents are typically high frequency traders and hedge funds that make money by writing puts. We suggest that, as these market participants can no longer delta-hedge by short selling stocks, they become less willing to sell protection. Boehmer et al. (2013) note that around 50 percent of all options trading is currently supplied by such informal market makers. Thus, we assume that the ban impaired the offering of puts normally provided by these market participants. For formal market makers, given that they are still allowed to short sell stocks to hedge their short put inventories, the ban does not place a hard constraint on their put-writing capabilities. Nevertheless, because stocks' bid-ask spread and price impact increase, as strongly suggested by the literature, hedging by shorting stocks became too expensive[14]. Market makers however, may also hedge short options by trading other options. However, as noted by both Battalio and Schultz (2011) and Grundy et al. (2012), bid-ask spreads on options on banned stocks also rose significantly during the 2008 US short sale ban. Thus, hedging with options may also have become prohibitively expensive during the 2011 European short sale ban.

[14] For example, Beber and Pagano (2013) illustrate that the 2008 US ban is associated with an increase in bid-ask spreads ranging between 1.64 and 1.98 percentage points in a sample of international stocks where the average bid-ask spread is 3.93 percentage points. As securities lending programs were in less demand by short-sellers during the ban, it became cheaper to borrow stocks as evidenced by three common measures of borrowing costs: the simple average fee, the simple average rebate, and the daily cost of borrow score, all provided by Markit Securities Finance. However, costs incurred by bid-ask spreads and price impact may easily outpace borrowing costs in times of market distress and thin trading activity.

One additional explanation for a smaller supply of puts during the ban is that options-sellers become more risk-sensitive following equity market declines, as noted by Gârleanu et al. (2009). The authors find that end-users have a net long-position in equity index options with a corresponding large net position in OTM puts. Conversely, market makers are short in OTM puts. Therefore, market makers face significant unhedged risk and, following a market decline, they experience losses and become more risk averse. In becoming more risk averse, they become more reluctant to expand trading inventories further and write additional puts. At the end of July 2011, days before the introduction of the European short sale ban, equity markets strongly corrected on the back of an intensifying European financial crisis. Therefore, it is not difficult to envision a high level of risk aversion among market makers during the ban. Hence, diminished willingness of market makers to sell puts could potentially have caused the supply of puts to strongly decrease. Holders of financial stocks interested in hedging their positions suddenly must pay much higher prices to buy protection[15].

On the ban announcement day, trading volume for puts on the EuroStoxx50 index reached 2,573,868, which is the second highest daily trading volume for this instrument in our sample[16]. A potential explanation for such an increase in trading volume is that, after the imposition of the ban, the skew from stock options relative to index options became too costly. The spread between the IV skew from the Eurostoxx 50 index put options and single stock puts, which is normally highly positive, was just marginally positive during the ban, reaching zero on December 20, 2011. Because index puts are far more liquid than single stock puts, a liquidity premium no longer existed and a migration from single stock puts to index puts may have taken place. Such explanation is consistent with "flight-to-liquidity" models suggested by Pástor and Stambaugh (2003) and Acharya and Pedersen (2005).

4.6. Panel Regression Analysis

[15] To illustrate, three-months 80 percent moneyness OTM puts on financial stock on average became 16 percent more expensive on August 11, 2011, compared to the previous 21 trading days.

[16] The heaviest trading in EuroStoxx50 puts took place on October 10, 2008, when the Belgian bank Dexia was bailed out, and 2,604,185 contracts were traded.

To further assess the effect created by the ban and trading activity into IV skews, we run a panel regression analysis with the IV skew (*VolSkew*) as the dependent variable. This regression allows us to isolate the relationship between the IV skew, banned stocks, and trading activity by controlling for other determinants of the IV skew such as information flow and idiosyncratic factors. We use the following firm specific control variables: daily turnover (*Turnover*), systematic risk (*Beta*), and firm size (*Size*). We use turnover as a proxy for stock liquidity, following Dennis and Mayhew (2002).

We calculate an individual stock's daily turnover by dividing its daily trading volume by its number of shares outstanding. The stock's beta is our control variable for systematic risk. The market return is assumed to be the equal-weighted average daily return for all stocks in our sample. The daily estimation of beta uses a rolling window of one year's worth of data, where the data begin one year before the first sample-date. Firm size is calculated as the number of shares outstanding on a specific day multiplied by the stock price. Single stock put option trading volume is computed as the average daily trading volume of puts divided by 1,000. Put trading volume is not used as an additional cross-sectional factor, as data are only available for a limited set of stocks (122 out of 186). Finally, we use the daily trading volume of puts on EuroStoxx50 divided by 1,000,000 (*E50PutVolume*) to capture the potential indirect substitution effect of trading pressure on single stocks' puts by index puts.

Control variables are uncorrelated with each other in a cross-sectional and time-series dimension (unreported here). We employ de-trended levels of sovereign CDS spreads for Belgium, France, Italy, and Spain (*CountryCDS*) and the V2X volatility index (*V2X*) as a control variable for country-specific and equity market information flows[17]. Additionally, we proxy firm-specific information flows with daily stock returns (*StockRt*) and trading pressure via single stock put option trading volume (*PutVolume*) and trading volume of puts on EuroStoxx50 (*E50PutVolume*). Our resulting model 1 is specified as follows:

[17] Based on the Johansen cointegration test, we find no cointegration between the four countries' de-trended CDS spreads and the V2X index at the five percent significance level.

$$
\begin{aligned}
VolSkew_{i,t} &= c + V2X_t + CountryCDS_{i,t} + StockRt_{i,t} + Turnover_{i,t} + Size_{i,t} + Beta_{i,t} + \dots \\
&+ D_t^{BanPeriod} + D_t^{BannedStock} + D_t^{BanPeriod} * D_i^{BannedStock} + D_t^{PostbanPeriod} + D_t^{PostBan} * D_i^{BannedStock} + \dots \\
&+ PutVolume_t + E50PutVolume + \varepsilon_{i,t}
\end{aligned}
\quad , \qquad (9)
$$

where $D_t^{BanPeriod}$ is a dummy variable equal to one if the date is within the ban period (August 11, 2011 until February 16, 2012), and zero otherwise. $D_t^{BannedStock}$ is a dummy variable equal to one if the underlying stock is a banned stock, and zero otherwise. $D_t^{PostBan}$ is a dummy variable equal to one if the date is after the lifting of the ban (from February 17, 2012 onwards), and zero otherwise. An additional dummy variable is created as an interaction term for these two dummies, $D_t^{BanPeriod} * D_t^{BannedStock}$. This variable captures the effect on the IV skew when two conditions — that the stock is banned and that the ban is in place — hold. The joint dummy $D_t^{PostBan} * D_t^{BannedStock}$ captures the relation between the IV skew and banned stocks in the post-ban period. *PutVolume* is the average daily trading volume of single stock puts, and *E50PutVolume* is the daily trading volume of puts on the EuroStoxx50 index.

Further, we use as our estimation method the Generalized Least Squares (GLS) approach to account for potential serial correlation in the residuals. We estimate our panel regression over three different periods: (a) the full period, ranging from February 15, 2008 to March 27, 2012; (b) the period that starts on April 27, 2010, when the European sovereign crisis is deemed to have begun, to March 27, 2012, and (c) the ban period, ranging from August 11, 2011 to February 16, 2012. Table 7 reports the results.

< Please insert Table 7 about here >

The results in Table 7 indicate that over the full sample period (column a) all coefficients are statistically significant at the one percent level, except for the joint dummy variable $D_t^{PostBan} * D_t^{BannedStock}$. The results for *V2X*, country CDS spreads, *Beta*, and *Turnover* are in line with results reported in the literature and with our expectations. We expected V2X and CDS spreads to be positively related to the IV skew, as jump risk priced for individual stocks is likely to increase with equity market volatility (see Dennis and Mayhew, 2002; Engle and Mistry, 2008) and with country credit risk. Contrary to our expectations, stock returns and size are positively related to the IV skew. Nevertheless, our size-skew estimates are in line with the results reported in Engle and Mistry (2008). They suggest that size proxies for beta, warranting a positive relationship between size and skew.

The results obtained from our dummy variables over the full sample period (column a) confirm that the short sale ban positively affected the IV skew for banned stocks: $D_t^{BanPeriod} * D_t^{BannedStock}$ has a positive sign and is statistically significant. This is a strong result, given the large set of control variables used. This finding suggests that the IV skew for banned stocks during the European short sale ban was abnormally high compared to that for unbanned stocks and that for banned stocks in other periods. The fact that $D_t^{BannedStock}$ has a positive sign and is significant implies that financial stocks have, on average, higher excess skew than other stocks, a finding that is consistent with insights provided by our descriptive statistics (Table 2). The three dummy estimates confirm that the IV skew for all stocks was higher during the short sale ban and that the effect was more pronounced for banned stocks.

Column b of Table 7 shows that during the Euro crisis pre-ban period, all parameter estimates for control variables have identical signs and comparable statistical significance levels compared to the results obtained in estimating Model 1 over the full period (column a). However, the empirical results change more strongly when we estimate Model 1 for the 2011 European short sale ban period, from August 11, 2011 to February 16, 2012. Column c of Table 7 shows that all control variables still have the same signs and the results are strongly statistically significant. Because we use such a short period, the dummies $D_t^{BanPeriod}$, $D_t^{PostBan}$, $D_t^{Ban} * D_t^{BannedStock}$, $D_t^{PostBan} * D_t^{BannedStock}$ are no longer applicable. The estimate of $D_t^{BannedStock}$ is no longer statistically significant. This outcome suggests that, within the ban period, banned financial stocks are no longer associated with higher IV skews relative to the average stock. The lack of significance of $D_t^{BannedStock}$ is, however, connected to its cross-correlation with beta for financial stocks during the ban (i.e., 1.30 relative to 0.90 for unbanned stocks). Additionally, *PutVolume* becomes negative and significantly related to *VolSkew*. Thus, a rise in the skew during the ban period was associated with a lower volume of single stock puts. This result is consistent with our hypothesis that a supply shift likely drove the IV skew during the ban, rather than a change in demand. In such a setting, large upward movements in the skew could have been caused by low trading volumes in OTM puts. At the same time, the link between *E50PutVolume* and *VolSkew* turns positive and significant. This relation is explained by the above-noted increase in the volume of index puts traded, in parallel with the supply-led rise in the IV skew during the ban.

A ban may be considered to be ineffective when selling pressure migrates from banned securities to alternative instruments. However, in the case of the 2011 European short sale ban, we argue that the migration of selling pressure from financial stocks to put options on European indices may not to have jeopardized the efficacy of the short sale ban. As a result of the migration, the ban appears to have diverted selling pressure initially concentrated in financial stocks to a larger share of the market. This hypothesis is consistent with the fact that contagion risk decreased for banned stocks during the ban but increased for unbanned stocks.

When the short sale ban was introduced on August 11, 2011, any further selling pressure on financial stocks could have led to destabilizing shocks and financial contagion. The price of OTM puts on banned stock likely rose as a result of lower trading volume, rather than through a substitution effect. The richness of OTM puts made it substantially more expensive for market participants to take a synthetic short position. Hence, imposition of the ban likely helped curbing downward price pressure, thereby benefiting financial sector stability.

4.7. *Robustness Tests*

In the above regression Model 1, we observed shifts in the signs of *PutVolume* and *E50PutVolume* across different sample periods. Hence, we now run an additional GLS panel regression as a robustness check to control for any influence of the short sale ban. We estimate a reduced form of Model 1 that excludes all dummies related to the ban, while using pre-ban data only. Thus, our Model 2 is specified as follows:

$$\begin{aligned} VolSkew_{i,t} &= c + V2X_t + CountryCDS_{i,t} + StockRt_{i,t} + Turnover_{i,t} + Size_{i,t} + Beta_{i,t} + \ldots \\ &+ PutVolume_t + Euro50PutVolume_t + \varepsilon_{i,t} \end{aligned} \quad , (10)$$

where the variables are defined as in Model 1. We estimate the panel regression Model 2 for the entire pre-ban period and for three different sample periods: (a) the pre-ban period (February 15, 2008 to August

10, 2011); (b) the US recession period (February 15, 2008 to June 30, 2009); (c) the ensuing stock market rally (July 1, 2009 to April 26, 2010); and (d) a sample that starts on April 27, 2010, when the European sovereign crisis is deemed to have begun, and runs through August 10, 2011, the last trading day before the short sale ban was implemented.

Panel A of Table 8 presents the Model 2 results. Column a shows that the Model 2 estimates for the pre-ban period are consistent with the Model 1 estimates (Table 7, column a) for the full sample period. Hence, we find that increased trading activity in single stock puts is linked to a high IV skew of single stock options. This relation confirms the findings of Bollen and Whaley (2004) and Gârleanu et al. (2009).

< Please insert Table 8 about here >

Columns b and c of Panel A show that the estimates of the Model 2 parameters during all three sub-periods have the same signs and statistical significance levels as those attained over the full pre-ban period (column a). Hence, the findings remain stable across various time periods and within two different model specifications.

As an additional robustness check, we analyze next whether the IV skews for stocks in other European countries increased around the date of the short sale ban announcement. Such an increase could be evidence of financial contagion effects in option markets. If so, the steep rise in jump risk would also be observed in other European countries that did not adopt the ban and that were vulnerable to or already hit by the financial crisis. European countries that fit such criteria are Greece, Ireland, and Portugal, see Grammatikos and Vermeulen (2011). We compile the IV skew data for only Ireland and Greece because Portugal does not have a public equity options market. In unreported results, we find no indication that jump risk for these stocks materially changes when the ban is introduced. These observations strengthen our earlier conclusion that the rise in the level of jump risk on the day of the ban announcement is connected to the short sale ban itself, as opposed to other reasons, like financial contagion.

Furthermore, results of our regression analysis are not materially changed when Model 1 is estimated without Belgian shares; a potential justification for excluding the Belgian data from the analysis is the fact that they experienced heavy governmental intervention in the period. Finally, using the IV slope measure of Yan (2011), which is the IV of close-to-ATM puts minus the one of calls, as the dependent variable instead of our IV skew measure, also does not alter our findings.

5. Conclusion

Recent research suggests that the short sale bans introduced during the 2008 crisis may have reduced market quality around the world, perhaps even to the extent that the ban's benefits were outpaced (see Battalio and Schultz, 2011; Grundy et al., 2012; Boehmer et al., 2013 and Beber and Pagano, 2013). Nevertheless, European market regulators in Belgium, France, Italy, and Spain re-introduced a short sale ban on financial stocks in August 2011 to combat the European financial crisis.

In order to analyze the effects of the European 2011 short sale ban on jump risk and contagion risk, we extract RNDs and IV skews from single stock options. RNDs indicate that jump risk of banned stocks is higher during the ban-period than in any other period analyzed. We find that on the day of the ban announcement, jump risk levels for both banned and unbanned stocks show a significant rise. Jump risk tends to increase for banned stocks even more than for unbanned ones. Furthermore, during the imposition of the ban, the banned stocks' average IV skews remain at an elevated level, whereas this metric drops for the unbanned stocks. During the ban, the median IV skews for both the banned and unbanned stocks, as well as the excess skews, reach their highest levels when compared to any other period in the sample. Thus, the short sale bans themselves seem to increase jump risk, especially for the banned stocks, even after controlling for information flow and stock specific factors.

Along the same vein, we document that contagion risk for both banned and unbanned stocks increases significantly already during the pre-ban period. For unbanned stocks, contagion risk rises even more upon imposition of the ban. However, per contrast, we find that contagion risk for banned stocks decreases during the ban relative to the pre-ban period. Obviously, the ban impacted banned stocks quite differently from unbanned stocks.

Furthermore, for banned stocks, trading volumes for both single stock puts and put-call ratios decline during the ban period. Such lower trading activity likely is caused by the reluctance of risk-averse market makers to sell options, increasing hedging costs, and because informal market makers are virtually prohibited to supply puts during the ban. Our findings suggest that supply-side frictions may also have pushed IV skew higher during the ban. OTM single stock puts become relatively expensive, thus the ban appears to curb synthetic shorting activity in financial stocks.

While the short sale ban is effective in restricting both outright and synthetic shorts on banned stocks, we do find evidence of trading migration to the index option market. Investors seem to switch from single stock puts to index puts because of “flight-to-liquidity” incentives. The selling pressure potentially diverted from the financial stocks to a larger share of the stock market, thereby reducing destabilizing effects in the financial sector. Such hypothesis is consistent with the decrease of contagion risk for banned stocks during the ban and its increase for unbanned stocks. Overall, our findings suggest that the 2011 European short sale bans reduced contagion risk in the financial sector, though at the cost of higher jump risk.

Appendix 1 – The Figlewski (2008) approach for extracting RND from implied volatilities

The Figleswki (2009) method extrapolates the RND beyond its body by fitting Generalized Extreme Value (GEV) distributions using two extreme anchor points on each side of the body of the RND and extending a tail with the same shape as the GEV curve. Equations (3) and (4) give, respectively, the GEV's cumulative distribution function and probability distribution function. Figleswki (2009) imposes three conditions to fit such GEV curves: i) that the total probability in the tail is the same for the RND's body and the GEV approximation; ii) that the shape of the RND's body equals the shape of the GEV curve in the inner anchor point; and iii) in the outer anchor point. For the right tail such conditions are:

Condition 1)	$F_{GEV}(X(\alpha_{innerR})) = \alpha_{innerR}$,
Condition 2)	$f_{GEV}(X(\alpha_{innerR})) = f_{Body}(X(\alpha_{innerR}))$,
Condition 3)	$f_{GEV}(X(\alpha_{outerR})) = f_{Body}(X(\alpha_{outerR}))$,

where $X(\alpha_{innerR})$ represents the exercise price corresponding to the α-quantile of the RND used as inner anchor point in the right tail, whereas $X(\alpha_{outerR})$ denotes the same but for the outer anchor point in the right tail. For the left tail such conditions have to be modified, becoming:

Condition 1)	$F_{GEV}(-X(\alpha_{innerL})) = 1 - \alpha_{innerL}$,
Condition 2)	$f_{GEV}(-X(\alpha_{innerL})) = f_{Body}(X(\alpha_{innerL}))$,
Condition 3)	$f_{GEV}(-X(\alpha_{outerL})) = f_{Body}(X(\alpha_{outerL}))$,

Therefore, we fit such GEV curves by implementing the following optimization:

$$GEV(\omega, \mu , \sigma) = Min \arg(y),$$

where, the objective function y_R for the right tail, following the three conditions above is:

$$y_R = [F_{GEV}\big(X(\alpha_{innerR})\big) - \alpha_{innerR}]^2 + [f_{GEV}\big(X(\alpha_{innerR})\big) - f_{Body}(X(\alpha_{innerR}))]^2 + \cdots$$
$$[f_{GEV}\big(X(\alpha_{outerR})\big) - f_{Body}(X(\alpha_{outerR}))]^2$$

whereas, for the left tail, such objective function y_L is:

$$y_L = [F_{GEV}\big(-X(\alpha_{innerL})\big) - 1 + \alpha_{innerL}]^2 + [f_{GEV}\big(-X(\alpha_{innerL})\big) - f_{Body}(X(\alpha_{innerL}))]^2 + \cdots$$
$$[f_{GEV}\big(-X(\alpha_{outerL})\big) - f_{Body}(X(\alpha_{outerL}))]^2$$

Figleswki's (2009) approach performs nicely for many observations in our sample. However, for some observations, the fitted GEV curves are implausible. We illustrate the problem encountered in Figure 4, where the right tail of the RND is reasonably fitted by GEV but the left tail is not. To avoid that we end up with implausible tails, we allow the inner anchor points to change by a predefined amount (ΔIAnchor), following a loop-algorithm from iteration *m*=1,…, *M*. Within such algorithm, inner anchor points are mainly the ones to shift to accommodate a better-behaved GEV curve. Exceptionally, however, outer anchor points are also shifted. The algorithm is the following, for the left tail:

1. Let the α-quantile inner anchor point (α_{innerL}) increase by ΔIAnchor as $m \rightarrow M$ a loop until $y_L^m > 5^{-25}$ and median of $\frac{\delta^2 f_{GEV}}{\delta K^2}\Big|_{K=0}^{K=X(\alpha_{innerL})} < 0$; otherwise stop loop.
2. If $y_L^{m-1} < y_L^m$, then, evaluates if median of $\frac{\delta^2 f_{GEV}}{\delta K^2}\Big|_{K=0}^{K=X(\alpha_{innerL})} > 0$. If yes, stop loop and use α-quantile inner anchor point (α_{innerL}) of y_L^m for GEV estimation. If median of $\frac{\delta^2 f_{GEV}}{\delta K^2}\Big|_{K=0}^{K=X(\alpha_{innerL})} < 0$, continue loop by increasing α-quantile inner anchor point (α_{innerL}) by ΔIAnchor.
3. If $y_L^{m-1} < y_L^m$, then, evaluate if $0.05 > \int_0^{X(\alpha_{outerL})} F_{GEV}^{m-1} > 0.1$. If so, stop loop and use α-quantile inner anchor point (α_{innerL}) of y_L^{m-1} for GEV estimation; otherwise continue loop.
4. If the α-quantile inner anchor point (α_{innerL}) increases up to the mode (peak) of the RND, then it stops increasing and the α-quantile outer anchor point (α_{outerL}) starts increasing by a very small step (ΔOAnchor) of 0.01 percent. If the α-quantile outer anchor point (α_{outerL}) increases more than 10 times, then stop loop and use α-quantile outer and inner anchor points from the iteration with lowest y_L is used for GEV estimation.

Thus, our modification to the Figlewski (2009) is that this approach will always take the RND-body as extracted from the IV, using the Breeden and Litzenberger (1978) formulae. In contrast, Figlewski (2009) substitutes the original RND in the interval between the inner anchor point and the end of the original RND.

< Please insert Figure 4 about here >

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Table 1. Overview of Banned Financial Stocks

This table lists the financial stocks banned from short selling on August 11, 2011 in Belgium, France, Italy, and Spain by their respective national financial market regulators in a coordinated act with the European Securities and Market Authority (ESMA).

Belgium	France	Italy	Spain
Ageas	April Group	Azimut Holding	Banca Cívica, S.A.
Dexia	Axa	Banca Carige	Banco Bilbao Vizcaya Argentaria, S.A.
KBC Group	BNP Paribas	Banca Finnat	Banco de Sabadell, S.A.
KBC Ancora	CIC	Banca Generali	Banco de Valencia
	CNP Assurances	Banca Ifis	Banco Español de Crédito, S.A.
	Crédit Agricole	Banca Intermobiliare	Banco Pastor, S.A.
	Euler Hermés	Banca Monte Paschi di Siena	Banco Popular Español, S.A.
	Natixis	Banca Popolare Emilia Romagna	Banco Santander, S.A.
	Paris Ré	Banca Popolare Etruria e Lazio	Bankia, S.A.,
	Scor	Banca Popolare Milano	Bankinter, S.A.
	Société Générale	Banca Popolare Sondrio	Bolsas y Mercados Españoles, S.A.
		Banca Profilo	Caixabank, S.A.
		Banco di Desio e Brianza	Caja de Ahorros del Mediterráneo
		Banco di Sardegna Rsp	Grupo Catalana de Occidente, S.A.
		Banco Popolare	Mapfre, S.A.
		Cattolica Assicurazioni	Renta 4 Servicios de Inversion, S.A.
		Credito Artigiano	
		Credito Emiliano	
		Credito Valtellinese	
		Fondiaria – Sai	
		Generali	
		Intesa Sanpaolo	
		Mediobanca	
		Mediolanum	
		Milano Assicurazioni	
		Ubi Banca	
		Unicredit	
		Unipoland	
		Vittoria Assicurazioni.	

Table 2. Descriptive Statistics

Panel A of this table provides descriptive statistics for the IV skews of unbanned and banned stocks calculated over the full sample period (February 15, 2008 to March 27, 2012) for the overall group of stocks as well as individually for Belgium, France, Italy, and Spain. Jarque-Bera normality tests are performed for all groups of stocks at the one percent significance level. Panel B presents the descriptive statistics for the level and daily differences in the excess IV skew calculated over the entire sample for the overall group of stocks as well as individually for Belgium, France, Italy and Spain. The excess skew (B-U index) is calculated as the difference between the average IV skew of all banned stocks and the average IV skew of all unbanned ones. Jarque-Bera normality tests are performed for the level and difference of all groups of stocks at the one percent significance level.

Panel A: Implied volatility skew

	Overall		*Belgium*		*France*		*Italy*		*Spain*	
	Unbanned	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*
Average	5.73	6.49	5.33	7.11	6.23	8.57	6.49	7.49	3.85	4.96
Median	5.31	5.98	5.14	6.94	6.09	8.16	6.23	7.24	3.19	3.98
Standard deviation	1.22	1.63	1.40	2.25	1.28	2.04	1.26	1.70	2.78	3.84
Skew	1.15	1.91	0.67	0.48	0.56	0.54	1.02	0.89	4.03	4.35
ExcessKurtosis	0.88	5.24	0.29	0.64	-0.11	-0.51	0.89	1.16	18.87	22.44
Jarque-Bera	266.0	1835.4	81.0	58.5	54.9	61.9	214.5	195.8	18395.7	25297.4

Panel B: Excess skew

	Overall		*Belgium*		*France*		*Italy*		*Spain*	
	Level	*Difference*	*Level*	*Difference*	*Level*	*Difference*	*Level*	*Difference*	*Level*	*Difference*
Average	0.76	0.002	1.78	0.001	2.34	0.004	1.00	0.000	1.11	-0.001
Median	0.71	-0.006	1.89	0.016	2.15	-0.003	1.06	0.005	0.89	-0.002
Standard deviation	0.77	0.567	2.08	1.576	1.41	0.679	1.11	0.898	1.78	1.396
Skew	1.47	0.749	-0.02	-0.219	0.65	0.199	-0.80	0.409	2.26	0.061
Kurtosis	7.65	11.32	0.93	6.59	-0.13	9.34	4.33	13.52	14.69	20.76
Jarque-Bera	2933.1	5697.5	38.2	1903.0	75.2	3814.5	932.7	8011.9	10314.1	18821.4

Table 3. Extreme Downside Risk Value Analysis

This table shows the ten, five and one percent extreme downside risk estimates, or VaR, of the RNDs for all unbanned and all banned stocks during the full sample period (February 15, 2008 to March 27, 2012), as well as for five different sub-periods. Panel A compares VaR of unbanned stocks versus the ones of banned stocks using the *t*-tests specified in Equations (6) and (7). The null hypothesis (H_0) for these tests is that there is no difference between VaR from banned and unbanned stocks. Rejection of H_0 is denoted by ***, **, and *, for the one, five, and ten percent significance levels, respectively. Panel B compares VaR of unbanned stocks across different samples. The superscripts are placed in the box of the second sub-sample that is compared. Therefore, the result for the statistical test that compares VaR between the first and the second sub-sample is shown in the box belonging to the 2009 stock market rally sub-sample. Likewise, Panel C compares statistics of Banned stocks in different samples.

Panel A: Comparison of unbanned and banned stocks within whole sample and within sub-samples

	10% VaR			*5% VaR*			*1% VaR*		
	Unbanned	*Banned*	*T-stat*	*Unbanned*	*Banned*	*T-stat*	*Unbanned*	*Banned*	*T-stat*
Full: 02/15/2008 – 03/27/2012	-0.31	-0.34	-1.3	-0.36	-0.40	-1.5	-0.51	-0.57*	-1.9
US recession: 02/15/2008 – 06/30/2009	-0.47	-0.50	-0.9	-0.53	-0.58	-1.2	-0.73	-0.82*	-1.8
2009 stock market rally: 07/01/2009 – 04/26/2010	-0.41	-0.40	0.3	-0.47	-0.46	0.2	-0.63	-0.63	0.0
Pre-ban European crisis: 04/27/2010 – 08/10/2011	-0.35	-0.38	-1.1	-0.40	-0.44	-1.0	-0.58	-0.62	-1.0
Ban period:08/11/2011 – 02/16/2012	-0.46	-0.62***	-4.0	-0.52	-0.70***	-4.0	-0.69	-0.94***	-4.1
Post-ban period: 02/17/2012 – 03/27/2012	-0.45	-0.56**	-2.1	-0.50	-0.64**	-2.2	-0.66	-0.84**	-2.2

Panel B: Comparison unbanned across sub-sample

	10% VaR		*5% VaR*		*1% VaR*	
	Unbanned	*T-stat*	*Unbanned*	*T-stat*	*Unbanned*	*T-stat*
Full: 02/15/2008 – 03/27/2012	-0.31	NA	-0.36	NA	-0.51	NA
US recession: 02/15/2008 – 06/30/2009	-0.47	NA	-0.53	NA	-0.73	NA
2009 stock market rally: 07/01/2009 – 04/26/2010	-0.41**	-2.1	-0.47**	-2.1	-0.63**	-2.3
Pre-ban European crisis: 04/27/2010 – 08/10/2011	-0.35**	-2.3	-0.40*	-1.9	-0.58	-1.0
Ban period:08/11/2011 – 02/16/2012	-0.46***	3.9	-0.52***	3.5	-0.70**	2.5
Post-ban period: 02/17/2012 – 03/27/2012	-0.45	-0.3	-0.50	-0.4	-0.66	-0.6

Panel C: Comparison banned across sub-sample

	10% VaR		*5% VaR*		*1% VaR*	
	Banned	*T-stat*	*Banned*	*T-stat*	*Banned*	*T-stat*
Full: 02/15/2008 – 03/27/2012	-0.34	NA	-0.40	NA	-0.57	NA
US recession: 02/15/2008 – 06/30/2009	-0.50	NA	-0.58	NA	-0.82	NA
2009 stock market rally: 07/01/2009 – 04/26/2010	-0.40***	-2.9	-0.46***	-3.2	-0.63***	-3.7
Pre-ban European crisis: 04/27/2010 – 08/10/2011	-0.37	-0.9	-0.44	-0.7	-0.62	-0.1
Ban period:08/11/2011 – 02/16/2012	-0.62***	6.4	-0.70***	6.1	-0.94***	5.4
Post-ban period: 02/17/2012 – 03/27/2012	-0.56	-0.9	-0.63	-1.0	-0.84	-1.2

Table 4. Median Implied Volatility Skew and Excess Skew

Panel A of this table shows the median IV skews for the overall group of stocks during the full sample period (February 15, 2008 to March 27, 2012), as well as for five different sub-periods. Two different groups of stocks are distinguished: All unbanned and all banned stocks. Mann-Whitney (MW) U-tests are applied to the IV skew of paired sample splits to infer whether the medians are statistically different from each other. The null hypothesis (H_0) for the MW U-test is that there is no difference between the two unrelated samples. Rejection of H_0 and its implication that medians between sub-samples are different is denoted by ***, **, and *, for the one, five, and ten percent significance levels, respectively. The superscripts are placed in the box of the second sub-sample that is compared. Therefore, the result for the statistical test that compares IV skew between the first and the second sub-sample is shown in the box belonging to the 2009 stock market rally sub-sample. Panel B shows the median IV excess skews between all unbanned and banned stocks for the overall sample (all) and for five different sub periods. The excess skew (B-U index) is calculated as the difference between the average IV skew of all banned stocks and of all unbanned ones. Similarly to Panel A, MW) U-tests are applied to the IV excess skews of paired sample splits to infer whether medians are statistically different from each other.

Panel A: Implied volatility skew

	Overall		*Belgium*		*France*		*Italy*		*Spain*	
	Unbanned	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*
Full: 02/15/2008 – 03/27/2012	5.31	5.98	5.14	6.94	6.09	8.16	6.23	7.24	3.19	3.98
US recession: 02/15/2008 – 06/30/2009	5.02	5.99	4.70	6.97	5.46	7.83	6.31	8.02	3.47	4.87
2009 stock market rally: 07/01/2009 – 04/26/2010	5.05*	5.41***	4.46***	6.24***	5.69**	7.46***	5.99***	6.00***	2.42***	3.26***
Pre-ban European crisis: 04/27/2010 – 08/10/2011	5.78***	6.05***	5.63***	7.81***	6.42***	8.14***	5.82	7.11***	3.60***	3.98***
Ban period:08/11/2011 – 02/16/2012	6.05**	7.34***	5.90***	7.28	6.99***	11.97***	7.14***	7.98***	3.06***	3.98
Post-ban period: 02/17/2012 – 03/27/2012	5.17***	6.37***	5.32***	5.25***	5.59***	10.65***	6.73***	5.49***	2.73***	4.98***

Panel B: B-U Index

	Overall	*Belgium*	*France*	*Italy*	*Spain*
	Banned-Unbanned	*Banned-Unbanned*	*Banned-Unbanned*	*Banned-Unbanned*	*Banned-Unbanned*
Full: 02/15/2008 – 03/27/2012	0.71	1.89	2.15	1.06	0.89
US recession: 02/15/2008 – 06/30/2009	0.74	2.04	2.29	1.57	1.34
2009 stock market rally: 07/01/2009 – 04/26/2010	0.30***	1.79	1.75***	0.05***	0.83***
Pre-ban European crisis: 04/27/2010 – 08/10/2011	0.68***	2.10	1.50***	1.20***	0.60***
Ban period:08/11/2011 – 02/16/2012	1.18***	1.64**	4.87***	0.90***	0.88***
Post-ban period: 02/17/2012 – 03/27/2012	1.15	0.01***	5.09	-1.05***	2.28***

Table 5. Conditional-Co-Crash Probabilities

This table shows the average conditional-co-crash (CCC) probabilities calculated by Equation (8) among all unbanned and all banned stocks during the full sample period (February 15, 2008 to March 27, 2012), as well as for five different sub-periods. Panel A compares CCC-probabilities of unbanned stocks versus the ones of banned stocks using a *t*-test similar to Equations (7) for each sample split. The null hypothesis (H_0) for these tests is that there is no difference between the average CCC-probability from banned and unbanned stocks. Rejection of H_0 is denoted by ***, **, and *, for the one, five, and ten percent significance levels, respectively. Panel B compares CCC-probabilities of unbanned stocks across different samples. The superscripts are placed in the box of the second sub-sample that is compared. Therefore, the result for the statistical test that compares CCC-probabilities between the first and the second sub-sample is shown in the box belonging to the 2009 stock market rally sub-sample. Likewise, Panel C compares statistics of banned stocks in different samples.

Panel A: Comparison unbanned vs banned stocks within whole sample and within sub-samples

	Conditional-Co-Crash-Probabilities		
	Unbanned	*Banned*	*T-stat*
Full: 02/15/2008 – 03/27/2012	0.29	0.32*	1.7
US recession: 02/15/2008 – 06/30/2009	0.26	0.27	0.4
2009 stock market rally: 07/01/2009 – 04/26/2010	0.23	0.28	1.6
Pre-ban European crisis: 04/27/2010 – 08/10/2011	0.32	0.42**	2.2
Ban period:08/11/2011 – 02/16/2012	0.41	0.32	-1.3
Post-ban period: 02/17/2012 – 03/27/2012	1.00	1.00	0.0

Panel B: Comparison un-banned and banned stocks across sub-sample

	Conditional-Co-Crash-Probabilities			
	Unbanned	*T-stat*	*Banned*	*T-stat*
Full: 02/15/2008 – 03/27/2012	0.29	NA	0.32	NA
US recession: 02/15/2008 – 06/30/2009	0.26	NA	0.27	NA
2009 stock market rally: 07/01/2009 – 04/26/2010	0.23	0.6	0.28	-0.2
Pre-ban European crisis: 04/27/2010 – 08/10/2011	0.32*	-2.0	0.42**	-2.2
Ban period:08/11/2011 – 02/16/2012	0.41	-1.2	0.32	1.2
Post-ban period: 02/17/2012 – 03/27/2012	1.00*	-2.0	1.00**	-2.2

Table 6. Median Trading Volume for Options on Unbanned and Banned Stocks

This table shows the median daily trading volume, measured by the number of contracts traded for puts and for all options (calls and puts together) as well as the median daily put-call volume ratio for all unbanned and banned stocks for the overall sample period and for five different sub periods. We apply Mann-Whitney U-tests to the median daily volume for puts, for all options and for the put-call ratio of unbanned and banned stocks of paired sample splits to test whether the medians are statistically different from each other. The null hypothesis is that there is no difference between the populations of the two unrelated samples. Rejection of the null and its implication that the medians between our five sub-periods are different is denoted by the asterisks ***, **, and *, indicating significance at the one, five, and ten percent level, respectively.

Trading volumes

	Put volume		*Options volume*		*Put-call volume ratio*	
	Unbanned	*Banned*	*Unbanned*	*Banned*	*Unbanned*	*Banned*
Full: 02/15/2008 – 03/27/2012	1,064	1,690	2,166	3,566	7.0	3.8
US recession: 02/15/2008 – 06/30/2009	877	1,377	1,849	3,046	7.1	4.1
2009 stock market rally: 07/01/2009 – 04/26/2010	1,200***	1,747***	2,488***	3,619***	7.0	3.4***
Pre-ban European crisis: 04/27/2010 – 08/10/2011	1,157	1,905**	2,325**	3,765	6.3	3.7**
Ban period:08/11/2011 – 02/16/2012	943***	1,727***	1,950***	3,526*	8.7***	3.8
Post-ban period: 02/17/2012 – 03/27/2012	1,245***	2,758***	2,554***	4,713***	7.9	5.7***

Table 7. Panel Regression Results

This table reports the panel regression results for Model 1. We distinguish three different periods: (a) Full, February 15, 2008 to March 27, 2012; (b) Euro Crisis, April 27, 2010 to March 27, 2012; and (c) Ban, August 11, 2011 to February 16, 2012. The single stock IV skew is the dependent variable and information flow (*CountryCDS* and *V2X*), firm specific control variables (*StockRt, Turnover, Size, Beta*), trading volume on single put options (*PutVolume*) and on index options (*Euros50PutVolume*), and dummies are the explanatory variables. The intercept is estimated as common to all cross-sections and no weighting is used in the cross-sections for estimation. Residuals are not normal for most cross-sections. We apply White-Heteroskedaticity consistent standard error and covariance estimates. The asterisks ***, **, and *, indicate significance at the one, five, and ten percent level, respectively.

	(a) Full Vol skew	(b) Euro crisis Vol skew	(c) Ban Vol skew
Intercept	2.823***	2.198***	-0.283
	(0.155)	(0.329)	(0.344)
V2X	0.053***	0.094***	0.035***
	(0.004)	(0.011)	(0.009)
Country CDS	0.006***	0.003***	0.008***
	(0.001)	(0.001)	(0.001)
Stock returns	5.509***	9.663***	7.055***
	(0.921)	(1.710)	(1.385)
Stock turnover	-19.317***	-22.535***	-25.602***
	(1.834)	(2.468)	(3.313)
Stock size	0.048***	0.066***	0.109***
	(0.001)	(0.001)	(0.001)
Stock *Beta*	1.015***	1.710***	3.702***
	(0.036)	(0.060)	(0.097)
Dummy Ban Period	-0.180	-1.104***	
	(0.118)	(0.195)	
Dummy Stock Banned	0.719***	0.357***	0.037
	(0.035)	(0.064)	(0.049)
Dummy Ban Period*Stock	0.306***	0.451***	
	(0.098)	(0.119)	
Dummy Post Ban	-0.201	-0.695***	
	(0.200)	(0.224)	
Dummy Post Ban*Stock	0.280	0.389*	
	(0.207)	(0.230)	
Overall put volume	0.305***	0.185	-0.297**
	(0.074)	(0.147)	(0.144)
EuroStoxx50 put volume	-0.723***	-1.358***	0.397**
	(0.119)	(0.220)	(0.189)
R^2	0.105	0.129	0.276
Observations	146,201	73,327	21,298

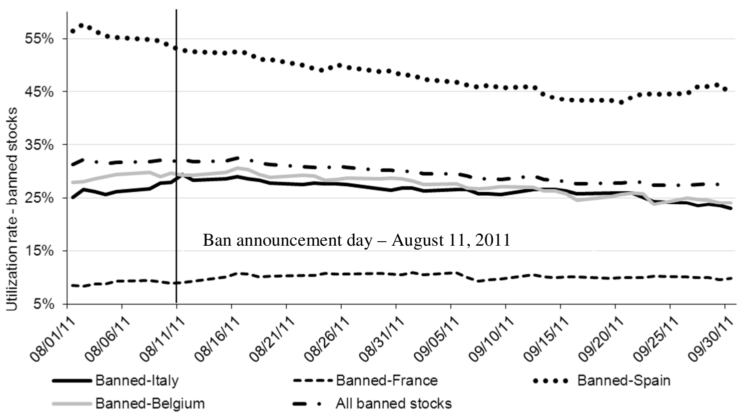
Table 8. Robustness Checks

Panel A reports the panel regression results for Model 2, in which we do not specify the dummies, using the pre-ban period, ranging from April 27, 2010 to March 27, 2012.We distinguish four different periods: (a) Full (pre-ban), February 15, 2008 to- August 10, 2011; (b) US recession, February 15, 2008 to June 30, 2009; (c) Market rally, July 1, 2009 to April 26, 2010, and (d) Euro Crisis, April 27, 2010 to August 10, 2011. Panel B reports the panel regression estimates excluding Belgium. Here we distinguish three different periods: (a) Full, 15 February 2008 to 27 March 2012; (b) Euro Crisis, 27 April 2010 to 27 March 2012; and (c) Ban, August 11, 2011 to February 16, 2012. The single stock IV skew is the dependent variable and information flow (*CountryCDS* and *V2X*), firm specific control variables (*StockRt, Turnover, Size, Beta*), trading volume on single put options (*PutVolume*) and on index options (*Euros50PutVolume*), and dummies are the explanatory variables. The intercept is estimated as common to all cross-sections and no weighting is used in the cross-sections for estimation. Residuals are not normal for most cross-sections. We applied White-Heteroskedaticity consistent standard error and covariance estimates. The asterisks ***, **, and *, indicate significance at the one, five, and ten percent level, respectively.

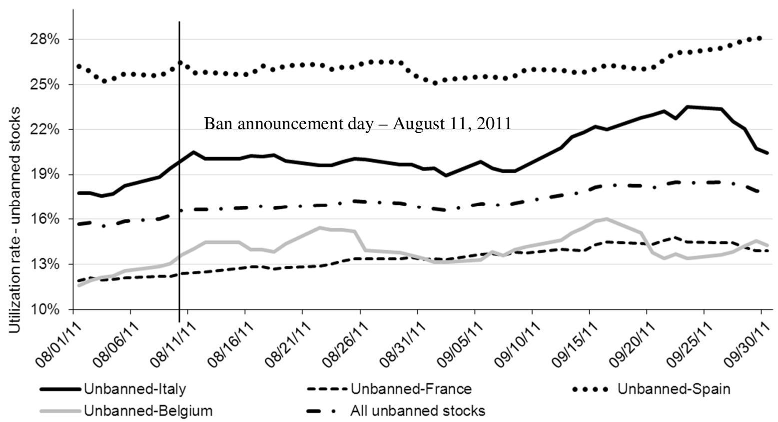
	Panel A				Panel B		
	(a) Full (pre-ban) Vol skew	(b) US recession	(c) Market rally	(d) Euro Crisis	(a) Full Vol skew	(b) Euro Crisis Vol skew	(c) Ban Vol skew
Intercept	3.213***	2.196***	6.035***	2.429***	2.893***	1.981***	-0.728**
	(0.170)	(0.102)	(0.281)	(0.473)	(0.164)	(0.353)	(0.335)
V2X	0.052***	0.052***	-0.022**	0.125***	0.050***	0.103***	0.037***
	(0.005)	(0.002)	(0.010)	(0.017)	(0.004)	(0.012)	(0.009)
Country CDS	0.005***	0.025***	0.024***	0.000	0.006***	0.003***	0.009***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
Stock returns	4.861***	4.085***	3.569***	11.958***	5.383***	9.735***	7.090***
	(1.060)	(0.667)	(1.234)	(2.927)	(0.995)	(1.832)	(1.374)
Stock turnover	-22.602***	-18.110***	-5.120	-30.131***	-19.694***	-23.984***	-27.241***
	(2.200)	(3.245)	(4.332)	(3.727)	(1.890)	(2.572)	(3.330)
Stock size	0.042***	0.034***	0.038***	0.054***	0.049***	0.069***	0.114***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Stock Beta	0.939***	0.944***	0.383***	1.262***	1.002***	1.739***	3.893***
	(0.031)	(0.046)	(0.037)	(0.052)	(0.040)	(0.068)	(0.099)
Dummy Ban Period					-0.184	-1.217***	
					(0.124)	(0.207)	
Dummy Stock Banned					0.712***	0.339***	0.087
					(0.036)	(0.068)	(0.056)
Dummy Ban Period*Stock					0.318***	0.497***	
					(0.106)	(0.128)	
Dummy Post Ban					-0.277	-0.743***	
					(0.210)	(0.237)	
Dummy Post Ban*Stock					0.438**	0.578**	
					(0.223)	(0.244)	
Overall put volume	0.364***	0.052	-0.016	0.370*	0.341***	0.236	-0.274*
	(0.081)	(0.054)	(0.056)	(0.201)	(0.077)	(0.156)	(0.140)
EuroStoxx50 put volume	-0.901***	-0.252***	-0.165	-2.188***	-0.724***	-1.452***	0.423**
	(0.132)	(0.082)	(0.164)	(0.308)	(0.124)	(0.233)	(0.184)
R^2	0.083	0.141	0.276	0.099	0.099	0.127	0.285
Observations	119759	44644	28230	46735	128757	64470	18702

Figure 1. Short Positions around Ban Date

A. Banned stocks

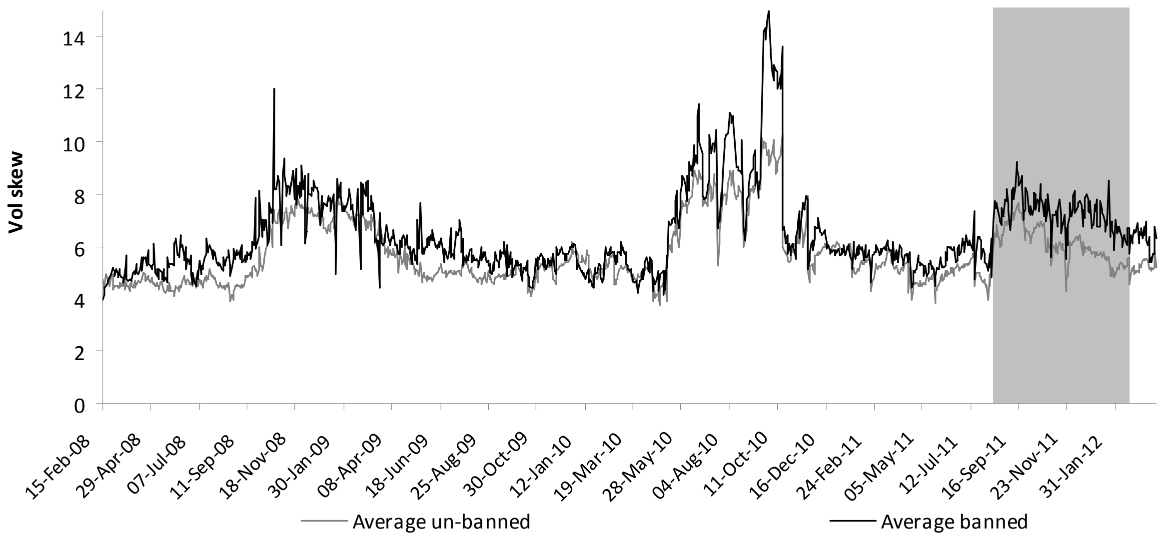


B. Unbanned stocks



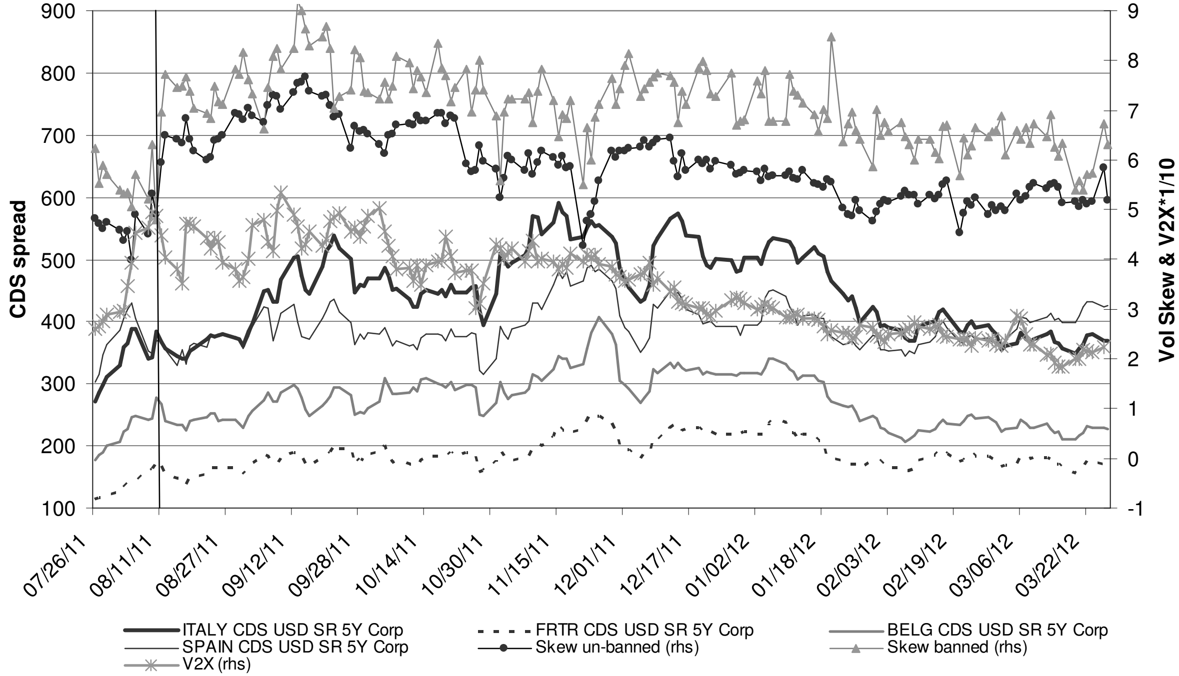
These figures present average short utilization rates calculated for banned (Section A) and unbanned stocks (Section B) in our sample. Utilization rates have been calculated for all banned and all unbanned stocks as well as for stocks in Belgium, France, Italy and Spain only. The utilization rate is calculated by $Utilization = 100 * \left(\frac{ValueOnLoan}{InventoryValue}\right)$, where *ValueOnLoan* is the beneficial owner value on the loan and *InventoryValue* is the beneficial owner inventory value. The utilization rate indicates the value of a stock utilized for securities lending against the total value of inventory available for lending. It is the demand to short as measured by the portion of shares in lending programs currently out on loan and ranges from zero to 100 percent.

Figure 2. Averaged Implied Volatility Skews



This figure depicts the average IV skew for banned and for unbanned stocks over the entire sample period. Averages are calculated over all stocks in Belgium, France, Italy, and Spain that have listed options. The IV skew per stock is calculated as the difference between the IV of the 80 percent moneyness OTM put option and the ATM put option. The European short-selling ban period (August 12, 2011 to February 16, 2012) is shadowed in the figure.

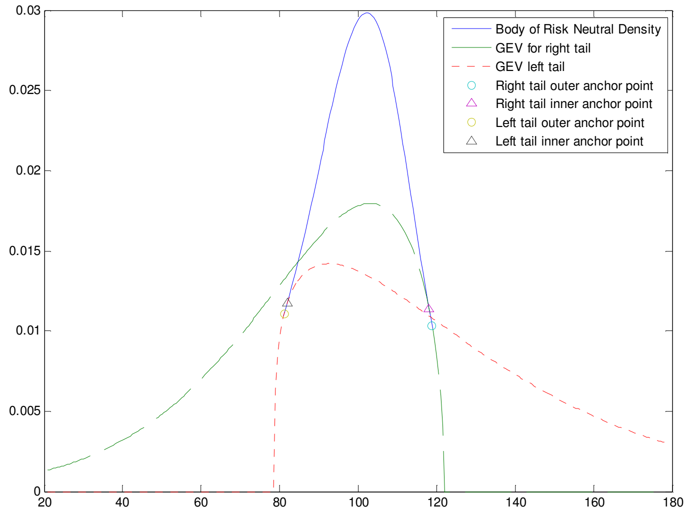
Figure 3. Sovereign CDS Spreads, V2X and Implied Volatility Skews



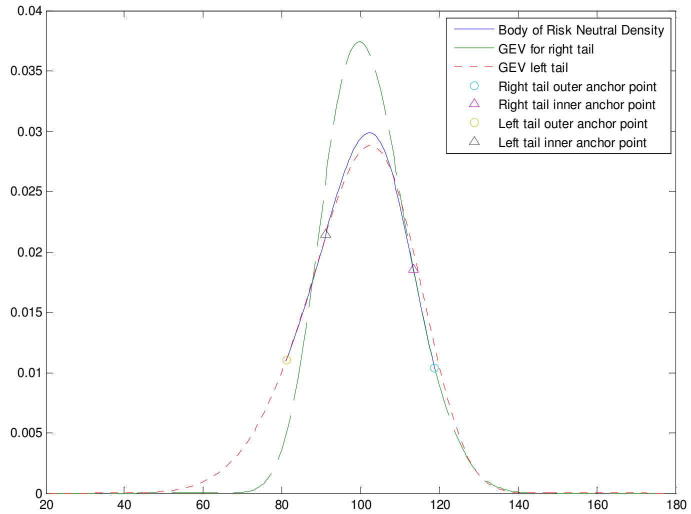
This figure depicts the five-year sovereign CDS spreads for Belgium, France, Italy, and Spain, the V2X, and the IV skews for unbanned and banned stocks. Sovereign CDS spreads are used as proxy for country-specific information flow in our study. V2X is the implied volatility index from the EuroStoxx 50 index and proxies for market-wide information flow. The V2X plot is multiplied by a factor of 1/10 to fit the same scale of the IV skew. The IV skew per stock is calculated as the difference between the IV of the 80 percent moneyness OTM put option and the ATM put option. The ban announcement day of August 11, 2011 is made distinct by a vertical line in the figure.

Figure 4. RND Extraction Using Figlewski (2009) Method to European Banned Stocks

A. Figlewski method



B. Modified Figlewski method



Plot A depicts the RND extract for banned stocks for March, 24, 2011 employing the Figlewski (2009) method, whereas Plot B depicts the RND extract for banned stocks for same date using the modified Figlewski method described in section 3. We note that in Plot A both the left and the right tails of the RND, fitted by GEV curves, are implausible because they contain abruptly declining tails under which the probability is close to zero. It is not the methodology that causes such distortion but the limited range of moneyness in our data set.