The Impact of Short Sale Bans on Option Trading Activity: Evidence from the Financial Crisis of 2007-2009 and its Aftermath

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Abstract

In the second half of 2008, the financial turmoil reached its peak. The situation was characterized by declining markets and intense speculations on the default of financial institutions. Financial regulators around the world responded by banning short sales of stocks. In most cases, however, the bans did not restrict trading in options. In this study, we use an international panel of options and stocks from 10 countries that adopted bans on short sales at some point in time from January 1, 2008 through February 28, 2011 to examine whether short sellers attempted to circumvent the bans by migrating to the option markets. More specifically, we do so by studying the impact of the bans on option trading activity. We find evidence that the short sale bans were associated with (i) an increase in option trading volume and open interest for both put options and call options, especially for options with high trading volume, and (ii) a decrease of the fraction of overall option trading volume that originates from put options, especially for options on U.S. stocks. Furthermore, the results indicate that bans on *naked* short sales were accompanied by an increase in option trading activity, whereas the estimated impact of bans on *all* short sales (both naked and covered short sales) is more ambiguous, especially for put options. Finally, our findings suggest that the bans increased the demand for call options more than the demand for put options. However, various subsample analyses and robustness tests indicate that the adoption of the bans is endogenous, and hence it is hard to assign the observed effects to an attempt by short sellers to circumvent the bans.

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1. Introduction

In the second half of 2008, the financial turmoil reached its peak. Stock prices of banks, insurance companies and other financial institutions experienced sharp declines and financial regulators around the world feared a loss of confidence for the financial markets and situations where short sellers pushed down prices well below the price level that would have resulted from a normal price discovery process. Many financial regulators responded by banning short sales of stocks. In most cases, however, the bans did not restrict trading in options. This means that it was still possible for speculators and market manipulators to profit from stock price declines by turning to the option markets.

In this study, we examine whether short sellers used the option markets as a substitute for direct short selling during the ban periods, and whether they did so primarily by trading put options or by trading call options. More specifically, we do so by studying the impact of the bans on option trading activity and its composition. As the bans had different introduction and lifting dates in different countries and featured different degrees of stringency, the effect of the bans are preferably examined through fixed effect panel data techniques. For this purpose, we put together a sample consisting of daily option and stock data from 10 countries that adopted bans on short sales at some point in time from January 1, 2008 through February 28, 2011, as well as data on various features of the bans adopted by these countries.

We address two principal questions. First, we examine whether the short sale bans affected option trading volume. To the extent short sellers tried to circumvent the short sale bans by taking corresponding positions on the option markets, we would expect option trading volume to increase during periods with restrictions on short selling. However, bans on short selling may also deteriorate the supply of liquidity on the option markets, making the net effect ambiguous. Market makers are major liquidity providers on the option markets, and even though they generally were exempted from the bans they were certainly not unaffected. The bans are likely to have reduced the informativeness of stock prices, slowed down the adjustment speed of prices to negative information, increased borrowing costs of banned stocks and imposed stricter delivery requirements also for market makers (Diamond and Verrecchia, 1987; Bai, Chang and Wang, 2006; and Kolasinksi, Reed and Thornock, 2009). As a result, it is likely that the bans reduced option market makers' ability to hedge, increased their hedging costs and made it more expensive for them to carry inventory and trade with informed investors. This, in turn, is likely to have suppressed option market makers' willingness to provide liquidity and hence increased the cost of trading in options on banned stocks. Increased cost of trading in options on stocks that are subject to short sale bans is indeed confirmed by empirical studies such as Battalio and Schultz (2009) and Grundy, Lim and Verwijmeren (2010). By also examining the impact on open interest of options, we are able to check whether a potential increase in option trading volume during periods with short sale bans can be assigned to an increase in the number of outstanding option contracts (potentially due to an increase in short exposure) or whether it is simply a reflection of existing contracts being traded more frequently.[1]

Secondly, we examine the impact of the bans on the composition of option trading volume. There are several ways by which investors can use the option markets to profit from declining stock prices, including buying a put option, writing a call option or taking a synthetic short position in the underlying stock consisting of a short call, a long put and a short bond. By examining the impact on the composition of option trading volume, we seek to identify the most commonly used substitute for direct short selling.

We find evidence that the short sale bans are associated with an increase in the daily number of traded option contracts for both put and call options. More specifically, the bans are associated with an increase in the number of traded puts by 1,117 contracts and an increase in the number of traded calls by 1,788 contracts. When measuring option trading activity as the number of option contracts outstanding (open interest), we find similar results: the bans are associated with an increase in open interest of puts by 39,827 contracts and an increase in open interest of calls by 64,151 contracts. Re-estimations for a subsample where the ten largest option classes in terms of average daily number of traded contracts are excluded indicate that these results can primarily be assigned to options with high trading volume.

In much of the analysis, we differentiate between two types of short sales; covered short sales and naked (uncovered) short sales. We define covered short selling as the practice of selling short a stock that has been borrowed (or arranged to be borrowed) from a third party in time to deliver it to the buyer within the settlement period, and naked short selling as the practice of selling short a stock that has *not* been borrowed (or arranged to be borrowed) in time to deliver it to the buyer within the settlement period.

When we allow for different impacts of naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), we find that naked bans generally are associated with a statistically and economically significant increase in option trading activity.[2] For covered bans, we obtain more ambiguous results: covered bans are generally associated with a statistically and economically significant increase in both trading volume and open interest of call options (even though the statistical significance generally is lower than for naked bans), but we find no evidence that covered bans had any impact on trading activity in put options. Theoretical models and results of previous empirical work suggest that stricter short sale restrictions have a more detrimental effect on the liquidity and market quality on the stock market (see, for example, Diamond and Verrecchia, 1987; and Beber and Pagano, 2010). Thus, a potential explanation to the observed difference between the

[1] Open interest is equivalent with the number of option contracts outstanding.

[2] Our definition of bans on naked short sales includes bans on selling short a stock that has not been borrowed (or arranged to be borrowed) in time to deliver it to the buyer within the settlement period *and* regulations imposing increased penalties on failure to deliver (that is, failure to deliver the stock to the buyer within the settlement period). The reason why we classify increased penalties on failure to deliver as a ban on naked short sales is that the adoption of stricter penalties was a commonly used way to limit abusive naked short selling.

impact of naked and covered bans is that covered bans may have resulted in higher inventory holding costs and adverse selection costs for option market makers than did naked bans. The results suggest that the short sale bans (both naked and covered bans) were accompanied by an increase in the demand for options on banned stocks, and that covered bans reduced the supply of liquidity in options on banned stocks more than did naked bans.

Furthermore, we find that the bans generally had a greater impact on the trading volume of calls than on the trading volume of puts. The bans are associated with a decrease of the fraction of overall option trading volume that originates from put options by 4.5 percentage points. This finding suggests that investors (i) speculated on declining stock prices by writing calls rather than buying puts, or (ii) viewed the low price levels prevailing during the ban periods as an opportunity to enter long positions at favorable prices (and hence demanded more long positions in call options). Re-estimations for a subsample where all option classes of U.S. stocks are excluded indicate that the estimated effect on the composition of option trading volume can primarily be assigned to options on U.S. stocks.

All results are robust to the introduction of variables reflecting time varying characteristics of the underlying stock. However, graphical results and various subsample re-estimations indicate that the adoption of the bans is endogenous. Hence, it is hard to assign the observed effects to an attempt by short sellers to circumvent the bans.

1.2 Contribution

There are at least two other recent studies that investigate the impact of the bans on short sales of stocks that were adopted in connection with the recent financial crisis on the option markets. Grundy et al. (2010) examine how the ban on short sales in specified U.S. financial stocks from September 19 through October 8, 2008 affected option trading volume, option bid-ask spreads and the relationship between option prices and stock prices. Studying the same event, Battalio and Schultz (2009) focus on a potential migration of short sellers to the option markets and examine how the confusion and regulatory uncertainty about option market makers' ability to hedge that arose in connection with the ban affected the cost of trading in options. Our study has many similarities with the papers by Grundy et al. and Battalio and Schultz. There are, however, several important differences, one being that it finds somewhat conflicting results.

Grundy et al. (2010) and Battalio and Schultz (2009) focus on the U.S. event exclusively. This study, on the other hand, uses an international panel of option and stock data from 10 countries, as well as data on various features of the short sale bans adopted by these countries. As pointed out by both Boehmer, Jones and Zhang (2009) and Beber and Pagano (2010), the 2008 U.S. ban on short sales in specified financial stocks coincided with the announcement of the Troubled Asset Relief Program (or simply

"TARP"), which makes it difficult to assign observed effects to the ban.[3] Our dataset allows us to run regressions for subsamples of countries and it reduces the potential endogeneity problem caused by the concomitant announcement of the program providing support to U.S. financial institutions. In a way then, by using an international panel of options and stocks, this study tests the robustness of the results reported in previous empirical work.

Moreover, we exploit cross-country differences in the inception and lifting of the bans by using fixed effect panel data regressions where we allow for stock-level fixed effects as well as time fixed effects. In contrast, Grundy et al. (2010) use an OLS difference-in-difference specification where they do not allow for stock-level fixed effects, while Battalio and Schultz (2009) run cross-sectional OLS regressions for each day from August 1, 2008 through October 21, 2008 for a number of different measures of liquidity and short exposure as left hand side variables.

Another important difference is that our dataset allows us to exploit differences in the stringency of the bans across countries and across stocks. Some countries prohibited all short sales (e.g. Canada and the U.K.), while others only prohibited naked short sales (e.g. Germany and France). Thus, we have the opportunity to take into account that bans on all short sales and bans only on naked short sales may have had different impact on option trading activity. Last but not least, we rely on Thomson's Datastream as our primary data source, while Battalio and Schultz (2009) examine a proprietary database of intraday option prices and quotes and Grundy et al. (2010) use data from OptionMetrics.

The remainder of this paper is organized as follows. Section 2 provides some background to the shorting bans adopted around the world and describes the cross-country differences in the features of the bans. Section 3 reviews related literature and develops testable hypotheses. In section 4 we describe the dataset and methodology. In section 5 we examine the impact of the bans on option trading activity and investigate whether short sellers migrated to the option markets in an attempt to circumvent the bans. We also test the robustness of the results by controlling for time varying characteristics of the underlying stock, as well as by running regressions for different subsamples of options, countries and time periods. Section 7 concludes.

[3] The Troubled Asset Relief Program (TARP) is a program of the U.S. government to purchase assets and equity from financial institutions to stabilize and restore the confidence for the financial sector. The program went into effect on October 3, 2008. Beber and Pagano (2010) find that bans on short sales in financial stocks adopted around the world in connection with the 2007-2009 financial crisis were associated with better return performance in banned stocks only for the U.S., not for other countries. This finding supports the idea that it was the announcement of the TARP, rather than the ban, that caused the observed effects. However, when it comes to liquidity and market quality Beber and Pagano obtain significant results for most of the countries included in their sample.

2. Adoption of Short Sale Bans around the World

Stock prices of banks, insurance companies and other financial institutions experienced sharp declines during the early autumn of 2008. Financial regulators around the world feared a loss of confidence for the financial markets and situations where short sellers, by speculating and spreading unfounded rumors regarding the financial health of financial institutions, would push down prices well below the price level that would have resulted from a normal price discovery process, resulting in panic selling and withdrawals of funds from banks, which in turn would push down prices even further and attract even more speculators. By the issuance of an emergency order on July 15, 2008 prohibiting naked short sales in 19 financial stocks, the U.S. Securities and Exchange Commission (the SEC) took a first step to prevent short sellers from manipulating prices and restore the confidence for the financial markets.[4] The emergency order remained in effect until August 12, 2008. After continued price declines, the SEC issued a new order on September 17, 2008 banning naked short sales in all stocks listed on U.S. stock exchanges.[5] The ban came into effect on September 18. In the announcement, the SEC states that it has "become concerned about sudden and unexplained declines in the prices of securities" and that "such price declines can give rise to questions about the underlying financial condition of an issuer, which in turn can create a crisis of confidence without a fundamental underlying basis" (SEC, 2008b, p. 1-2).

The same day, on September 18, the U.K. financial regulator (the FSA) announced a temporary ban on all short sales (both naked and covered short sales) in specified financial stocks. The SEC followed suit the following morning by prohibiting all short sales in 797 financial stocks.[6] In the announcement, the SEC motivates the adoption of the stricter and more extensive ban in the following way: "Recent market conditions have made us concerned that short selling in the securities of a wider range of financial institutions may be causing sudden and excessive fluctuations of the prices of such securities in such a manner so as to threaten fair and orderly markets" (SEC, 2008c, p. 1). In addition, the SEC issued another order requiring institutional investors to file and report information concerning daily short sale activities.[7] The temporary ban prohibiting all short sales in 797 financials stock was lifted on the evening of October 8. However, the ban on naked short sales remained in place.

During the weeks that followed, in particular between September 19 and September 23, most major stock exchanges around the world adopted restrictions on short sales. However, the features of the bans

[4] See SEC Release no. 58166 (SEC, 2008a). In the release, the SEC states that it is forbidden to sell short a stock unless the investor (or its agent) has "borrowed or arranged to borrow the security or otherwise has the security available to borrow in its inventory prior to effecting such short sale and delivers the security on settlement date" (SEC, 2008a, p. 3-4). As a result of the rule, short sellers had to borrow the stock they wanted to sell short three days earlier than before. Prior to the adoption of the rule, short sellers did not have to borrow the stock until the end of the three day settlement period. If short sellers closed their positions within the settlement period they did not have to borrow the stock at all (Kolasinksi et al., 2009).

[5] See SEC Release no. 34-58572 (SEC, 2008b)

[6] See SEC Release no. 34-58592 (SEC, 2008c)

[7] See SEC Release no. 58591 (SEC, 2008d)

differed across countries. Some countries only prohibited naked short sales (e.g. Austria, Belgium, Denmark, France, Germany, Greece, Japan, Luxembourg, Portugal and Spain), other countries banned all short sales (e.g. Australia, Canada, Ireland, Norway, South Korea and the U.K.), and a third category of countries adopted some combination of bans on naked short sales and bans on all short sales (e.g. Italy, the Netherlands, Switzerland and the U.S).[8] Moreover, some countries adopted rules requiring short sellers to disclose their trades (e.g. Belgium, France, Ireland, the Netherlands, Spain, Sweden, the U.K and the U.S.), while others did not. Also, the scope of the ban regimes differed across countries. Most countries restricted the ban to financial stocks only (e.g. Canada, the U.S. and most of the European countries), while others included short sales of all stocks (e.g. Italy, Japan, South Korea and Spain). In addition, the introduction and removal dates, as well as the length of the ban periods, differed across countries (for example, Austria, Denmark, Italy and Spain adopted their bans later than the U.K. and the U.S., and the length of the ban periods in France, Italy, Spain and Switzerland by far exceeded that of the ban periods in Canada and the U.K.). In some countries (e.g. Austria, Denmark, Germany, Spain, Switzerland and the U.S.) the short sale bans are still in effect.[9] Introduction and removal dates, as well as cross-country differences in the stringency of the bans that were adopted by the countries included in our dataset are shown in Figure 2 and Table 1 (see section "4.1 Data" for further description of the dataset).

With few exceptions, the bans only prohibited short sales of stocks and not trading in options.[10] This means that in most countries it was still possible for speculators and market manipulators to profit from stock price declines by simply turning to the option markets. For example, investors could gain short exposure by buying a put, writing a call or creating a synthetic short position in the underlying stock consisting of a short call, a long put and a short bond.

Also potentially important for the impact of the bans on short sales of stocks on the option markets is whether and to what extent option market makers were exempted from the bans. Market makers are major liquidity providers on the option markets and the ability to sell stocks short is critical for their ability to hedge their positions (for example, if would-be short sellers attempted to circumvent the bans

[8] In the Netherlands, the financial regulator initially imposed a ban on naked short sales in financial stocks, which was later replaced by a ban on all short sales (both naked and covered short sales). In Italy, the financial regulator initially banned naked short sales. This ban was later replaced by a ban on all short sales (both naked and covered short sales). When the ban on all short sales expired, it was still prohibited to perform naked short sales for some time. In Switzerland, the regulator adopted a ban on all short sales (both naked and covered short sales) in specified financial stocks and a ban on naked short sales in all other stocks. The course of events in the U.S. is described in the text. (See Table 1 for a more detailed account of the cross-country differences in the scope and stringency of the bans).

[9] We know that the short sale ban adopted in Spain was still in effect in May 2010. We have found no information suggesting that the ban has expired since then, and hence we have assumed that it was in effect throughout the sample period. It is only Switzerland's ban on naked short sales that is still in effect. The ban on all short sales (both naked and covered short sales) in specified financial stocks expired on January 16, 2009. Germany initially banned naked short sales in specified financial stocks between September 20, 2008 and May 31, 2009. The German regulator issued a new ban on naked short sales in specified financial stocks on May 19, 2010 and it is this ban that is still in effect.

[10] Due to vague formulations in the announcements of the bans we find it difficult to be sure whether some of the bans also prohibited synthetic short selling of the underlying stock through the option and/or other derivatives markets. In Ireland, for example, the financial regulator prohibited investors from making profits on falling Irish bank stocks. In France, the ban applied to uncovered short sales of listed securities, including spot, forward, and option transactions involving the listed equities.

by migrating to the option markets we would expect them to demand long positions in puts and short positions in calls, and market makers usually hedge written puts and long positions in calls by shorting the underlying stock).[11] When the ability to hedge is reduced, the risk and cost of supplying liquidity increases. In as far the short sale bans reduced market makers' ability to sell short, the bans are likely to have had a negative impact on option market makers' willingness to supply liquidity, and hence a negative impact on the supply of liquidity in options on banned stocks.

Option market makers were generally exempted from the bans being examined in this study.[12] However, even in countries in which the restrictions did not apply to option market makers (or in which the rules provided at least some exemption for option market makers), there may still have been initial confusion whether and to what extent they were exempted. For example, Battalio and Schultz (2009) describe the initial confusion about option market makers' ability to hedge in connection with the 2008 U.S. ban. The emergency order issued on September 19, 2008 banning all short sales in 797 financial stocks provided a general exception for "registered market makers, block positioners, or other market makers obligated to quote in the over-the-counter market" conducting short sales as part of bona fide market making activity (SEC, 2008c, p. 3).[13] Moreover, the SEC provided a more specific exception for market makers "when selling short as part of bona fide market making and hedging activities related directly to bona fide market making in derivatives" until midnight on September 19, 2008 (SEC, 2008c, p. 4). The exemption was a way to facilitate expiration of options on September 20 (which was a monthly option expiration date). According to Battalio and Schultz, the latter exemption suggested that option market makers would be unable to sell stocks short throughout the remainder of the short sale ban. They also mention that by midday on September 19, several options market makers threatened to stop supplying liquidity if they were not allowed to hedge by shorting the underlying stock. This further supports that option market makers were confused whether and to what extent they were exempted from the ban.

On the morning of September 22, the SEC issued an amendment release confirming the exception for option market makers and that the exception would continue for the duration of the order.[14] The SEC

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[11] The role of market makers is to take on the role of counterparty when an investor wants to buy or sell a financial instrument, and hence to supply liquidity in situations where no other counterparty wishes to sell or buy the same amount. The market maker quote a bid price at which it is obligated to buy and an ask price at which it is obligated to sell, and its main source of income is the spread between the higher ask price and the lower bid price.

[12] In some countries, however, the exemption for market makers was not as distinct as in others, and in a limited number of countries, market makers do not seem to have been exempted at all: In, Switzerland the prohibitions applied to all market participants. In Australia, transactions by all market participants were covered by the prohibition, with a "limited exception for covered short selling by market makers". In Canada, the order did not distinguish between different types of market participants. In Greece, market makers were exempted from the ban for "market making transactions under specific circumstances". In the U.S., the emergency order of September 17, 2008 imposed stricter delivery requirements on sales of all U.S. stocks, and no exceptions were provided for market makers.

[13] According to the U.S. SEC, factors that indicate that a market maker is engaged in bona-fide market making activities may include whether the market maker is putting their own capital at risk to provide continuous two-sided quotes in markets, taking the other side of trades when there are short-term buy-side and sell-side imbalances, or attempting to prevent excess volatility (SEC, 2008f).

[14] See SEC Release no. 58611 (SEC, 2008e)

states that the purpose of the accommodation was "to permit market makers to continue to provide liquidity to the markets" (SEC, 2008e, p. 3). However, in an attempt to prevent investors from circumventing the ban by trading options, the SEC added a regulation that market makers were not allowed to sell short if the market maker knew that the customer or counterparty's transaction would result in the customer or counterparty establishing or increasing an "economic net short position (i.e., through actual positions, derivatives or otherwise)" in stocks subject to the ban (SEC, 2008e, p. 4). This means that, even when the exception for option market makers had been clarified, there were still special circumstances preventing them from engaging in normal hedging activities.

In some cases, the bans imposed stricter requirements on option market makers also in other dimensions. For example, the emergency order issued by the SEC on September 17, 2008 included a temporary rule (Rule 204T), effective immediately, imposing stricter delivery requirements on sales of all U.S. stocks.[15] The temporary rule imposed "a penalty on any participant of a registered clearing agency, and any broker-dealer from which it receives trades for clearance and settlement, for having a fail to deliver position at a registered clearing agency in any equity security" (SEC, 2008b, p. 2). No exceptions were provided for market makers. Historically, the U.S. regulator had not been as strict in penalizing failures to deliver. Confusion about option market makers ability to hedge and stricter requirements imposed on option market makers are probably not unique for the U.S. event, but rather the situation is likely to have looked about the same in most of the countries that adopted bans on short sales.[16]

To summarize, even though trading in options and earning profits from declining prices in banned stocks through the option markets generally were allowed, there is reason to believe that confusion about option market makers' ability to hedge and stricter requirements imposed on option market makers may have suppressed a potential migration of would-be short sellers to the option markets. In the next section we will look at factors that may have affected option market makers and the supply of liquidity in options on banned stocks in more detail. Which of the two effects – the wish of investors to sell banned stocks short (and hence their attempt to find alternative ways to profit from declining prices) and the reduced willingness of option market makers to supply liquidity – that were the stronger remains to be seen.

3. Previous Literature – Theory and Evidence

At present, there are many studies on the impact of short sale restrictions on liquidity, overpricing and price discovery on the stock market. The effect of short sale restrictions on the equity option markets has, however, not obtained the same level of attention. The limited work that has been done concentrates mainly on the effect on option trading volume and migration of short sellers from the stock market to the

[15] See SEC Release no. 34-58572 (SEC, 2008b). Rule 204T was later adopted permanently by the SEC in July 2009 and is known as SEC Rule 204.

[16] Ideally, we would have conducted similar research about how the bans affected option market makers' hedging activities for each country included in our dataset. However, this would be far too time consuming to fit in the scope of this thesis.

option markets, liquidity on the option markets, as well as the relationship between stock prices and option prices. In this thesis we address the two former effects: the effect on option trading volume and the migration to the option markets. In the remainder of this section we give a brief description of the effects as predicted by theory and the evidence that has been found so far. However, we think it is well motivated to start off by giving an account of theory and evidence related to the effects of short sale bans on the underlying market, the stock market.

3.1 Impact on the Stock Market

3.1.1 Stock Prices

Theoretical models that deal with the impact of short sale bans on stock prices can be divided into two main categories. In models such as Miller (1977), investors have differences in beliefs. In this world, short sale restrictions prevent pessimists' views to be impounded in stock prices, and optimists do not take any notice of the absence of the views of pessimists. Hence, under a short sale ban, stocks subject to the ban end up overvalued.

Diamond and Verrecchia (1987) show that if all investors have rational expectations, short sale restrictions do not cause stock prices to be overpriced on average. As short sellers are not likely to perform short sales for liquidity reasons, they are more likely to be informed than the average investor. Under a short sale ban, investors take into account that well informed would-be short sellers are shut out of the market and adjust the price they are willing to pay accordingly. However, even if stock prices are unbiased, short sale restrictions reduce the adjustment speed of stock prices to negative news and result in less accurate pricing. Empirical evidence such as Desai, Krishnamurthy and Venkataraman (2006) and Boehmer, Jones and Zhang (2008) support that short sellers are well informed and manage to earn abnormal returns.

Bai et al. (2006) add another assumption to the rational investor setting and show that when rational investors are also risk-averse, they require higher expected return which put downward pressure on stock prices. The idea is that the slower price adjustments amplifies the risk perceived by uninformed investors, which in turn results in higher required return and lower prices. Thus, in the setting proposed by Bai et al. bans on short sales may amplify price declines rather than support stock prices (the latter being one of the most frequently used arguments used by politicians to motivate short sale bans). Hence, the impact of short sale restrictions on stock prices, as predicted by theory, is ambiguous.

In general, there is more evidence in support for the assumption that investors have differences in beliefs than the assumption that all investors have rational expectations. Jones and Lamont (2002) use data from the NYSE from the 1920's and 1930's and find evidence that stocks which are expensive to short show relatively high valuations and low future returns compared to stocks which can be shorted more cheaply. Using data from the Hong Kong stock market, Chang, Cheng, and Yu (2007) find that short sale constraints tend to cause overvaluation by analyzing price impacts when stocks are added to a list with stocks designated as eligible for shorting.

Recent studies use panel data techniques to examine price impacts of the short sale bans that were adopted around the world in connection with the 2007-2009 financial crisis. Boehmer et al. (2009) focus on the U.S. stock market and find that the announcement of a temporary ban on all short sales in specified financial stocks on September 19, 2008 was followed by a sharp increase in the prices of stocks subject to the ban (however, they admit that it is hard to assign this effect to the ban because of the concomitant announcement of the Troubled Asset Relief Program). Beber and Pagano (2010) use an international panel of stocks from 30 countries, of which 20 adopted bans on short sales at some point in time between January 1, 2008 and June 23, 2009. They find that neither bans on naked short sales nor disclosure requirements were associated with excess returns relative to a group of unbanned control stocks. Bans on all short sales (both naked and covered short sales), on the other hand, were associated with significant return underperformance. When narrowing the sample to U.S. stocks only, Beber and Pagano find that the U.S. ban was associated with a cumulative excess return for banned stocks exceeding that of a control group of non-banned stocks. In terms of impact on stock prices then, the U.S. stock market behaved quite differently compared to the rest of the world.[17] Another recent study by Harris, Namvar and Philips (2009) examines the price impacts of the U.S. ban using a factor approach. Consistent with the findings of Boehmer et al. (2009), they find that the ban led to a substantial increase in the prices of banned stocks. Interestingly, they also find results suggesting that the TARP legislation is *not* a significant factor in explaining the inflation in stock prices.

3.1.2 Market Liquidity

Diamond and Verrecchia (1987) suggest that by preventing investors from trading on bad news, short sale bans result in slower price adjustments and lower informational efficiency, which in turn widens the bid-ask spreads. In addition, by replacing informative short sale transaction with less informative no-trade outcomes, the information revelation becomes slower, which also tend to widen the spreads.

Overall, the empirical evidence supports the idea that short sale restrictions worsen market liquidity. Boehmer et al. (2009) provide evidence that liquidity, as measured in terms of bid-ask spreads and price impacts, for stocks subject to the 2008 U.S. short sale ban was significantly damaged compared to non-banned control stocks. These results are supported by Beber and Pagano (2010) who find that the short sale bans adopted around the world in connection with the 2007-2009 financial crisis were detrimental for liquidity, as measured by bid-ask spreads and the Amihud (2002) illiquidity indicator.

Other studies, such as Jones (2008) and Charoenrook and Daouk (2005) do, however, find more ambiguous results. When investigating the altered strictness of short sale restrictions in the U.S. in the

[17] This finding supports Boehmer et al.'s (2009) worries that the positive effect found for the U.S. may result from the concomitant announcements of the TARP.

1920's and 1930's, Jones finds that the introduction of stricter requirements for brokers when lending shares of customers worsened liquidity, while a rule only allowing short sale transactions to be performed on an uptick had the opposite effect.[18] Using stock data and information on the history of short sale regulation and put option trading from 111 countries, Charoenrook and Daouk find that short sale restrictions are associated with higher dollar trading volume.

3.1.3 Price Discovery

As pointed out in the previous section, the model presented by Diamond and Verrecchia (1987) suggests that by preventing agents from trading on bad news, short sale restrictions lead to slower price discovery. Most empirical evidence supports this prediction. With data from 46 stock markets, Bris, Goetzmann and Zhu (2007) find that new information is impounded in stock prices more quickly in countries where there are no restrictions on short sales. Saffi and Sigurdsson (2008) use lending transaction data for 17,015 stocks from 26 stock markets and find that stocks in which short selling is less constrained, as measured by lending supply and borrowing fees, have shorter price delays. Using an international panel of stocks and fixed effect panel data techniques, Beber and Pagano (2010) find that bans on short sales adopted in connection with the financial crisis of 2007-2009 reduced the adjustment speed of prices, especially in declining markets.

3.2 Impact on the Option Markets

The availability of theoretical models predicting the impact of restrictions on short sales of stocks on the options markets is rather limited. In some cases, we are therefore left with intuition. Fortunately, the empirical research is somewhat more extensive. Much of the evidence comes in indirect form by showing that the short sale bans adopted around the world in connection with the 2007-2009 financial crisis had different impact on stocks with listed options and stocks without listed options. Interestingly, the limited number of studies that focus directly on the impact on the option markets report contrasting results.

3.2.1 Testable Hypotheses

Intuitively, if the short sale restrictions induced short sellers to switch to the option markets, we would expect to see an increase in the demand for options (for long put positions and short call positions) and hence an increase in both option trading volume and open interest of options.[19] Boehmer et al. (2009), who study the 2008 U.S. ban, report that for stocks on the SEC's original ban list (published on September 19, 2008), short sales account for a cross-sectional average of 21.78 percent of total trading volume before the ban and that this figure drops to 7.71 percent during the ban period.[20] A proportionate

[18] Under the uptick rule, short selling was only allowed if the order was placed at a price above the last traded price of the security, or at the last traded price if that price was higher than the price in the previous trade. The rule went into effect in 1938 and was removed in 2007.

[19] Open interest is equivalent with the number of option contracts outstanding.

[20] According to Comerton and Putnins (2009), the fraction of total trading volume that involved a short sale is even higher. They report that 26 percent of the total trading volume on the NYSE in 2006 involved a short sale. For 2008, the corresponding share

move of short sale activity in banned stocks to the option markets would hence have resulted in a quite substantial increase in both option trading volume and open interest. For stocks that were never subject to the ban, the proportion of short sales of total trading volume only drops from 19.48 percent to 18.23 percent.

As pointed out earlier, the short sale bans are also likely to have affected the supply side of the option markets, mainly by increasing the inventory holding costs and adverse selection costs of option market makers.[21] Market makers are major liquidity providers on the option markets and the ability to sell stocks short is critical for their ability to hedge their positions. When the ability to hedge is reduced, the risk and cost of supplying liquidity increases. More specifically, the inventory holding costs increase as the value of the market makers' inventory becomes more exposed to unfavorable price movements of the underlying stock. Furthermore, when naked short selling is prohibited and covered short selling is allowed, the demand for equity loans of banned stocks should increase. This in turn, should drive up the price of borrowing banned stocks and hence increase the inventory holding costs of market makers by increasing the cost of hedging (at least as long as market makers are not fully exempted from the ban). Kolasinksi et al. (2009) do indeed report that the emergency order of July 15, which imposed stricter delivery requirements on the sales of all U.S. stocks, led to a substantial increase in the fees charged by lenders.

Short sale bans are also likely to increase the adverse selection costs of option market makers. Models such as Diamond and Verrecchia (1987) and Bai et al. (2006), which assume that investors have rational expectations, suggest that by preventing investors from trading on bad news, short sale restrictions lead to slower price discovery and less informative prices. These predictions are empirically confirmed by Beber and Pagano (2010), Bris et al. (2007) and Saffi and Sigurdsson (2008). By reducing the adjustment speed of stock prices to negative news and reducing the informative content of prices, short sale bans create a greater information asymmetry between informed and uninformed market participants. Hence, in as far market makers are uninformed, short sale bans are likely to increase the information asymmetry between informed investors and market makers, and hence the adverse selection costs of market makers.

Bai et al. (2006) show that when rational investors are also risk-averse, the slower adjustments of stock prices to negative news and reduced informative content of prices caused by short sale bans will increase the risk perceived by uninformed market participants. Beber and Pagano (2010) argue that to the extent market makers are uninformed, this will increase market makers' inventory holding costs.

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is 38 percent. Boehmer et al. (2009) suggest that the short sales in banned stocks observed during the ban period probably were performed by market makers, who were still able to sell short as part of their market making and hedging activities.

[21] Inventory holding costs are costs associated with the carrying of positions acquired by market makers when supplying liquidity. The inventory holding cost has two main components: the opportunity cost of capital tied up in the market makers' inventory and the risk that the value of the inventory declines as a result of unfavorable price movements. Adverse selection costs arise because market makers, in guaranteeing to buy and sell securities at the quoted bid and ask prices, may trade with investors that are better informed (Bollen, Smith and Whaley, 2004).

According to models such as Stoll (1978) and Bollen, Smith and Whaley (2004), the inventory holding costs of market makers increases with the volatility of stock prices. This is very intuitive: increased volatility increases the uncertainty about the future value of the positions held in inventory. The impact of short sale bans on stock price volatility is, however, ambiguous. Bai et al. (2006) show that if investors have rational expectations and are risk-averse, and in the absence of information asymmetry, short sale constraints reduce stock price volatility by limiting the fluctuation in demand of stocks due to negative news. Under information asymmetry, however, short sale constraints can cause the volatility to increase, the intuition once again being that short sale restrictions reduce the informativeness of stock prices and increase the uncertainty of uninformed investors. Reduced informational efficiency of market prices, they argue, causes trades to have greater impact on prices and increases the variance. Boemer et al. (2009) report that the 2008 U.S. ban was associated with a large increase in price volatility and that the volatility increased more for banned stocks than for matched control stocks that were never subject to the ban.[22] Even though these results may suffer from reverse causality (that is, short sale bans were adopted due to increasing volatility and they were particularly imposed on stocks for which the variance increased the most), inventory holding costs of option market makers may have increased during the ban periods as a result of increased volatility of stock prices of banned stocks.

According to Stoll (1978) and Bollen et al. (2004), market makers require compensation for bearing inventory risk and adverse selection costs. Hence, increased inventory holding costs and adverse selection costs of market makers should result in market makers increasing their bid-ask spreads for options on banned stocks. This corresponds to a decrease in the supply of liquidity in options on banned stocks. Furthermore, bid-ask spreads of options are also likely to be affected by the competition among suppliers of liquidity. If market makers are the only market participants being exempted from a short sale ban, the ban should reduce the competition from other providers of liquidity and hence allow market makers to lower their bid prices and increase their ask prices (Bollen et al., 2004).

The two effects of short sale bans on option trading activity – the increased demand for options on banned stocks and the reduced willingness of option market makers to supply liquidity in options on banned stocks – are summarized in Figure 1. For simplicity, let's use option trading volume as a measure for option trading activity. The quantity of options is on the x-axis and the price of option liquidity is on the y-axis. The increase in demand is illustrated by a shift of the blue solid demand curve denoted "Pre ban demand" to the right. The ban is also likely to increase the adverse selection costs and inventory holding costs of option market makers, which reduces their willingness to provide liquidity. This is illustrated by a shift of the red solid supply curve denoted "Pre ban supply" to the left. As we can see, the net effect of the ban on option trading volume is ambiguous. If the demand increases more than the corresponding decrease of the supply, we would expect option trading volume to increase from Q1 to Q2. On the other hand, if the supply decreases more than the corresponding increase in the demand, we

[22] Boehmer et al. (2009) measure volatility as the difference between the highest and the lowest transaction price recorded for a given stock on a given day, divided by the volume-weighted average trade price for that day.

would expect option trading volume to decrease from Q1 to Q2*. In both scenarios, the equilibrium price paid by the party demanding liquidity to the party supplying liquidity (mainly option market makers) increases (shown by the increase of the price from P1 to P2=P2*). On the option markets, this increase would be reflected by an increase in the bid-ask spreads of options on banned stocks.

3.2.1 Direct Evidence

In a memorandum produced in January 2009, the SEC's Office of Economic Analysis investigates the impact of the SEC's emergency order of July 15, 2008, which restricted naked short selling in stocks of 19 financial entities, on option market activity (OEA 2009). By analyzing open interest and option trading volume, the OEA finds no evidence that the emergency order had a significant impact on option activity, and hence it finds no evidence that short sellers switched from the stock market to the option markets.

Grundy et al. (2010) examine how the ban on all short sales in 797 U.S. financial stocks from September 19 through October 8, 2008 affected option trading volume, option bid-ask spreads and the relationship between prices of options and stocks. They find that the ban was associated with a decline in the trading volume of put options and that the decline was significantly larger for banned stocks than for non-banned stocks. When controlling for changes in stock trading volume, stock return and the level of the VIX, they find that the average daily trading volume of put options on banned stocks declines by 3,194 contracts during the ban, while the corresponding decrease for unbanned stocks amounts to 490 contracts. As a result, they conclude that the ban was not circumvented by a migration of short sellers to the option markets. Grundy at al. also report that bid-ask spreads increased for options on both banned and non-banned stocks during the ban. In line with theory, the increase in bid-ask spreads is found to be significantly larger for options on banned stocks. In a two-stage simultaneous equation estimation, they find that the increase in bid-ask spreads explains some, but not all, of the decline in option trading volume observed during the ban. Hence, they argue that one reason that they do not observe the expected increase in option trading volume is that the cost associated with trading in options was higher during the ban period.

In another recent paper studying the 2008 U.S. ban, Battalio and Schultz (2009) examine how the confusion and regulatory uncertainty about option market makers ability to hedge that arose in connection with the ban affected the option markets. More specifically, they examine whether short sellers used the option markets to circumvent the ban, whether the cost of trading in options increased during the ban period and whether the ban resulted in divergences in the relative prices of options and stocks. Battalio and Schultz examine the potential migration of short sellers to the option markets by studying option-to-stock trading volume and changes in short exposure taken through the option markets. They find that the ratio of option-to-stock trading volume for banned stocks is similar to that of a control group of non-banned stocks during the entire sample period.[23] Furthermore, using data on

[23] Battalio and Schultz (2009) multiply the trading volume of put and call contracts by 100 (since each contract covers 100 shares) and divide the product by the number of shares traded of the underlying stock on that day.

individual option transactions from the Chicago Board Options Exchange (CBOE) and International Securities Exchange (ISE) they compute the change in short exposure in the underlying stock taken through the option markets.[24] They find that aggregated short exposure is comparable for options on banned and non-banned stocks throughout the sample period. Hence, they find little evidence that short sellers migrated to the option market in an attempt to circumvent the ban.

When it comes to trading costs, Battalio and Schultz (2009) find that the ban had a substantial impact on option bid-ask spreads. From September 22 through the lifting of the ban on October 8, they find that the quoted bid-ask spreads are on average 10 percentage points higher for options on stocks that were subject to the ban than for options on non-banned control stocks. The effect on the bid-ask spreads is comparable for put and call options. In line with the thoughts of Grundy et al. (2010), Battalio and Schultz mean that the inflated cost of gaining short exposure in banned stocks through the option markets is one of the main reasons why short sellers do not seem to have switched to the option markets.

3.2.2 Indirect Evidence

Harris et al. (2009) examine the price impacts of the 2008 U.S. short sale ban using a factor approach and find evidence that stock prices increased the most for stocks without listed options. They argue that this finding points towards that the option markets enabled investors to form synthetic short positions and hence served as an effective substitute for regular short selling on the stock market. Supporting results are presented by Beber and Pagano (2010), who report that the short sale bans adopted around the world in connection with the 2007-2009 financial crisis deteriorated liquidity, as measured by bid-ask spreads and the Amihud (2002) illiquidity indicator, particularly for stocks with no listed options, small market capitalization and high volatility.

Kolasinksi et al. (2009) suggest that informed traders are more likely to use options as a substitute for direct short selling (as the substitute provided by the option markets is only available to investors sophisticated enough to perform synthetic short selling). Hence, for stocks with traded options, bans are likely to increase the fraction of informed short sellers, and thus increase the informative content of short sales. When testing this implication, they find that the 2008 U.S. short sale ban increased the informativeness of short sales, as measured by the relation between short selling volume and returns, in banned stocks and that the increase was particularly large for stocks with listed options. Moreover, in contrast to the findings of Beber and Pagano (2010), Kolasinksi et al. find that the ban's deteriorating effect on liquidity, as measured by Amihud's (2002) illiquidity indicator and share turnover, was especially strong for stocks with listed options.

[24] Battalio and Schultz (2009) classify option trades as being "open-buys" (trades that are initiated by buyers to open a position), "open-sells" (trades that are initiated by sellers to open a position), "close-buys" (trades that are initiated by buyers to close a position) or "close-sells" (trades that are initiated by sellers to close a position). Then they compute the change in short exposure as follows: $change\ in\ short\ exposure = (Put\ Open\ Buy + Call\ Open\ Sell) - (Put\ Close\ Buy + Call\ Close\ Sell)$.

4. Data and Methodology

4.1 Data

Most of the analysis is based on an international panel of options and stocks from 10 countries that adopted restrictions on short sales at some point in time during our sample period – January 1, 2008 through February 28, 2011. The data consists of option trading volume and the number of option contracts outstanding (open interest) for put and call options respectively, bid and ask prices of the underlying stocks, stock trading volume, and country of listing.[25] All option and stock data is from Thomson's Datastream and measured with daily frequency. In order to arrive at a dataset of manageable size, we only include financial stocks and options on financial stocks in our sample.[26]

Information regarding which stocks that were subject to bans on short sales at some point in time during the sample period, as well as information on which time periods the bans were in force, is obtained from the webpage of the government agency responsible for financial regulation in each country respectively, as well as from Beber and Pagano (2010) and Strömqvist (2009). From above mentioned sources, we also obtain information on various features of the bans, including whether the bans prohibited naked short sales or all short sales (both naked and covered short sales), whether and to what extent the bans also prohibited investors from taking corresponding short positions on the option markets, and whether and to what extent option market makers were exempted from the bans. Inception and removal dates, as well as cross-country differences in the stringency of the bans are illustrated in Figure 2. A somewhat more detailed description of the features and scopes of the bans, and the overall structure of our dataset is presented in Table 1.

Figure 3 illustrates the diffusion of short sale bans by showing the fraction of stocks in our sample that is subject to a ban over time. The fraction of stocks subject to a ban increases sharply from 0 percent in June 2008 to about 35 percent in September. In October it jumps further to 95 percent. Subsequently, the fraction of banned stocks gradually decreases back to about 80 percent of the sample throughout the remaining part of the sample period. Naked bans represent a majority of the bans in all months but September 2008, the reason for the exception being the announcement by the U.S. SEC on September 19 of a ban prohibiting all short sales (both naked and covered short sales) in 797 financial stocks. The fraction of stocks subject to a covered ban reaches its peak of about 35 percent in October 2008. As from June 2009, all bans in our sample are represented by bans on naked short sales.

[25] Option trading volume is measured in number of traded contracts and stock trading volume is measured in thousands of shares. All prices are measured at closing. Each option contract covers 100 shares.

[26] Even with a sample of financial stocks only, the generation of some variables is very time consuming. For example, one of our control variables, a rolling standard deviation of the return of the underlying stock based on the previous 10 observations, took some 20 hours to generate.

Beber and Pagano (2010) identify 20 countries around the world that adopted restrictions on short sales in connection with the financial crisis of 2007-2009 (Australia, Austria, Belgium, Canada, Denmark, France, Germany, Greece, Ireland, Italy, Japan, Luxembourg, the Netherlands, Norway, Portugal, South Korea, Spain, Switzerland, the U.K. and the U.S.). Datastream only has option data for 13 of these countries (for all countries except Denmark, Greece, Ireland, Japan, Luxembourg, Portugal and South Korea). In addition, we exclude countries for which Datastream only has option data for one of the following time periods: (i) before the ban, (ii) during the ban or (iii) after the ban. This means that options on stocks that are listed in Australia, Belgium and Norway, for which Datastream only provides option data for the post-ban period (Australia and Norway) and the ban period (Belgium), are left out of the dataset.

For countries where short sale restrictions were imposed on financial stocks only, but where the financial regulator did not issue a list on which financial stocks that were to be subject to the ban, we determine which stocks to categorize as banned ourselves. For this purpose, we use a function in Datastream that allows the user to search for equities by industry and country of listing. We identify 5 industry categories that we consider belong to the financial sector.[27] Thereafter, we make the assumption that all companies in these industry categories were subject to the ban in question. For countries where short sale restrictions were imposed on all stocks, we use the same function and the same 5 industry categories to determine which stocks are to be considered as financial stocks. After these adjustments, we arrive at a sample of 1,278 financial stocks from 10 countries.

Subsequently, we manually search for listed options for each of the 1,278 financial stocks in Datastream. As we do not aim to identify different impacts of short sale bans on options with certain strike prices or expiry dates, all option data is based on a data category in Datastream called "Option Class". An option class includes all put *or* call options on an underlying stock listed on a marketplace, irrespective of the exercise prices and expiry dates of the individual options.[28] Hence, in an ideal world each stock has two option classes: one for puts and one for calls. Occasionally, stocks have options listed on more than one marketplace (and hence the stock has more than one pair of option classes). In such cases, we only include the pair of option classes listed on the "domestic" option marketplace, which we define as the option marketplace of the country in which the underlying stock is primarily listed. In the rare case where a stock only has options listed in another country than the country in which the underlying stock is primarily listed, the stock (and hence the foreign-listed pair of option classes) is left out of the sample. Finally, we only include classes of American options, and there are two main reasons for this. First, the liquidity in American options is generally substantially higher than the liquidity in European options on

[27] The industry categories are Banks, Financial Services (Sector), Life Insurance, Nonlife Insurance and Real Estate Investment & Services.

[28] For example, the daily number of traded put (call) contracts of an option class is measured as the sum of all put (call) contracts of all put (call) options on the underlying stock listed on a certain derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options.

the same stock. Secondly, all stocks that have listed European options also have listed American options, while opposite does not hold true. Because the dataset only contains one option class for puts and one option class for calls per stock, the dataset only contains one observation per underlying stock per day.

To be included in the sample, an observation must fulfill the following condition: data on trading volume of the call and the put option classes has to be contemporaneously available. This requirement leaves out 18 percent of the number of observations, whereof the majority is excluded due to missing option data from January 1, 2008 until May 20, 2008 (the dropped observations primarily belong to option classes of stocks listed in the U.S.). The final dataset contains 171 pairs of option classes (and hence 171 financial stocks) and 116,528 observations.

As seen in Table 1, 91 of the 171 pairs of option classes are listed on an U.S. exchange, and these pairs represent more than half of the total sample in terms of number of observations (52 percent of the total number of observations). Moreover, as can be seen in Figure 4, a handful of stocks (firms) represent a major fraction of the aggregated average daily trading volume: half of the average daily number of put contracts traded belongs to options on seven stocks, half of the average daily number of call contracts traded belongs to options on five stocks, and half of the average daily number of traded shares in the underlying stock is represented by seven stocks.

Figure 5 illustrates the distribution of the monthly cumulative number of traded put and call contracts, as well as the monthly cumulative trading volume in the underlying stock across countries over time. Notice that stocks and pairs of option classes of some countries lack observations for some time periods, the most obvious example being that we have no data for U.S. options and stocks until May 2008. Figure 5 also illustrates the development of average daily trading volume for an average option class (stock) over time. Table 4 reports average daily trading volume, open interest and put-to-option ratio for an average option class (stock) for each month throughout the sample period.[29] Further summary statistics are reported in Table 2.

4.2 Methodology

We examine the impact of the bans on the level and composition of option trading activity graphically and in fixed effect panel regressions. The graphic illustrations show different measures of option trading activity over time. Most of the figures compare the average level of the relevant measure for each day during a 41 day window ranging from 20 days before the *introduction date* up to 20 days after the *introduction date* with the average level of the measure for all non-ban observations except those belonging to the 20 days preceding the introduction of the bans. We complement these figures with similar graphic illustrations focusing on the average level of the relevant measure for each day during a 41 day window

[29] The put-to-option ratio is calculated as the daily number of traded put contracts over the total daily number of traded option contracts (puts+calls).

ranging from 20 days before the *removal date* up to 20 days after the *removal date*.[30] The graphic illustrations provide indications on how the short sale bans impacted option market activity, even though formal testing is left to the fixed effect panel regressions.

Using our panel of option and stock data from 10 countries adopting bans on short sales at some point in time during the period from January 1, 2008 through February 28, 2011 and various subsamples, we estimate the following fixed effect model:

$$Y_{it} = \beta_0 + \alpha_i + \beta_1 D_{it}^{ban} + \gamma_t + (\theta X_{it}) + \varepsilon_{it} ,$$

where Y_{it} is some measure of option trading activity for option class i at time t. On the right-hand side of the equation, we include

(i) a constant β_0;

(ii) a constant α_i representing time-invariant characteristics of the underlying stock of option class i;

(iii) a dummy variable that is equal to one if short sales (either naked short sales or all short sales – both naked and covered) are forbidden in the underlying stock of option class i at time t and zero otherwise;

(iv) weekly time dummies γ_t;

(v) a vector $\mathrm{X_{it}}$ for a number of time varying control variables, including return of the underlying stock, bid-ask spreads of the underlying stock and the standard deviation of the return of the underlying stock based on the previous 10 observations.[31] These control variables are only used in selected regressions and are not part of our standard specification.

All three control variables are chosen to represent time varying characteristics of the underlying stocks, and they have all been pointed out as being correlated with the adoption of short sale bans by previous empirical work of, among others, Beber and Pagano (2010) and Boehmer et al. (2009). By controlling for these variables in selected specifications, we aim to capture the direct effect of the short sale bans on

[30] More specifically, we examine option trading activity around the introduction of the bans by estimating the following fixed effect model: $Y_{it} = \beta_0 + \alpha_i + d'20 + \dots + d'1 + d0 + d1 + \dots + d20 + \gamma_t + \varepsilon_{it}$, where Y_{it} is some measure of option trading activity for option class i at time t, α_i represents time-invariant characteristics of the underlying stock of option class i (stock-level fixed effects), γ_t represents weekly time dummies (time fixed effects), $d'20$ is a day dummy that equals one if the relevant ban is introduced in 20 days and zero otherwise, $d'19$ is a day dummy that equals one if the ban is introduced in 19 days and zero otherwise, and so on… This continues up to $d19$, which is a day dummy that equals one if the ban was introduced 19 days ago and zero otherwise. $d20$ stands out by equaling one throughout the remainder of the ban period (that is, from ban day 20 up to the removal of the ban) and zero for all other days. We then plot the estimated coefficients of the dummies $d'20$ through $d20$. In this setup, all non-ban observations, except those belonging to the 20 days *preceding* the ban introduction date, perform the role as control. In the graphs focusing on the removal of the bans, a corresponding procedure is used. The main difference is that all non-ban observations, except those belonging to the 20 days *following* the removal of the bans perform the role as control. In the graphic illustrations, we make no difference between naked and covered bans.

[31] Stock return is based on closing prices and expressed in decimals. Bid-ask spread is the percentage bid-ask spread expressed in decimals and calculated as the ask price minus the bid price (at closing) divided by the bid-ask midpoint. Standard deviation of the return of the underlying stock is based on the previous 10 observations and expressed in decimals.

option trading activity, that is, changes in option trading activity due to investors trying to find alternative ways to profit from declining prices, rather than impacts resulting from (i) the bans having a detrimental effect on the liquidity and market quality on the underlying market and (ii) increased inventory holding costs and adverse selection costs of option market makers. Hence, the extended specification allows us to check whether the results of our standard specification are not simply a reflection of the short sale bans having an impact on the underlying market (other than not allowing short sales).[32]

We use the following measures as left-hand side variables:

(i) number of option contracts traded;
(ii) number of outstanding option contracts (open interest);
(iii) a ratio of the number of put option contracts traded over the total number of option contracts traded (puts+calls).

For the measures not being ratios, we also test for nonlinear relationships by including regressions with log-level functional form.

Our specification of the regression equation exploits the following features of the panel dataset:

(i) different countries adopted and removed their short sale bans on different dates;
(ii) the stringency of the bans varied across countries: in some countries the ban only prohibited naked short sales, while in others, the ban ruled out all short sales (both naked and covered short sales).[33]

In this setup, option classes of stocks that are not yet subject to a ban, and option classes of stocks that have been subject to a been but for which the ban has been removed, as well as each option class before and after the ban periods of their underlying stocks, perform the role as control. All regressions allow for stock-level fixed effects as well as time fixed effects (weekly). This implies that we allow time-invariant characteristics of the underlying stock of option class i to be correlated with other explanatory variables in each time period (for example whether the underlying stock of option class i, at some point in time, will be subject to a ban or not), and for unobserved variables that are constant across stocks but change over time (e.g. broad market moves). All regressions are estimated with heteroskedasticity robust standard

[32] Investors buy puts when they believe that the price of the underlying stock will decline going forward, and equivalently they buy calls when they believe that the stock price will increase. Hence, if investors base their beliefs about tomorrow's stock return on today's return, the return of the underlying stock is likely to affect option trading volume. In addition, we find it reasonable to suspect that stock return is negatively correlated with the adoption of bans on short selling. Bid-ask spreads is a widely used measure for liquidity on the stock market, and Beber and Pagano (2010) and Boehmer et al. (2009) find evidence that the size of the bid-ask spread is positively correlated with the adoption of short sale bans. Moreover, bid-ask spreads may also be correlated with the inventory holding costs and adverse selection costs of option markets makers, which in turn is likely to affect the supply of liquidity in options (Grundy et al., 2010). We also choose to control for the standard deviation of the return of the underlying stock, since its changes may increase the inventory risk of option market makers, and hence reduce their willingness to provide liquidity (Bollen et al., 2004).

[33] See section "2. Adoption of Short Sale bans around the world" and "4.1 Data" for a more detailed account of how the features of the short sale bans differed across countries.

errors which are clustered at the level of the underlying stock, the latter allowing for dependence in the residuals across time for a given option class i (and hence for correlation in ε_{it} within i).

All regressions are run with two different sets of ban dummies. In the first set up, we include one ban dummy for all types of short sale bans, and hence we make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered). In the second set up, we allow for different effects of naked and covered bans by including one ban dummy for naked bans and one ban dummy for covered bans.[34] In addition, all regressions are run in two further versions: one version where the control variables presented above are excluded (the "standard specification"), and one version where the control variables are included (the "extended specification"). Thus, we use four specifications for each dependent variable.

A major issue with our specification is that there is reason to believe that the adoption of short sale bans is endogenous. Most financial regulators imposed the bans as a reaction to sharp price declines, high return volatility, intense speculations on the default of financial institutions and a loss of confidence for the financial markets. Probably, the problems were more severe and emerged earlier in some countries in our sample than in others, which may have resulted in some countries adopting short sale bans earlier and/or bans with stricter requirements than others. In as far option trading activity is correlated with these time varying variables related to the crisis, and to the extent the variables are country specific (that is, not market-wide), we run the risk of obtaining biased results.

We deal with the potential endogeneity problem related to the adoption of the bans by the introduction of our stock-level control variables. After having read the ban announcements of several financial regulators, we have noticed that return and volatility on the stock market were two of the most important measures taken into account by financial regulators when deciding whether and when they should impose bans on short sales. By adding stock-level return and stock-level volatility (as well as stock-level bid-ask spreads) to the list of explanatory variables, we aim to control for cross-country differences in time varying variables affecting the likelihood of a country imposing a ban. In addition, we rerun all regressions for a sample period starting in December 2008 (rather than in January 2008). In December, all countries in our dataset had adopted short sale bans. This means that we rely solely on the removal of the bans in our difference-in-difference re-estimation. This procedure eliminates the potential endogeneity problem

[34] More specifically, we allow for different impacts of naked and covered bans by estimating the following fixed effect model: $Y_{it} = \beta_0 + \alpha_i + \beta_1 D_{it}^{nakedban} + \beta_2 D_{it}^{coveredban} + \gamma_t + (\theta X_{it}) + \varepsilon_{it}$, where $D_{it}^{nakedban}$ is a dummy variable that is equal to one if naked short sales are forbidden and covered short sales are allowed in the underlying stock of option class i at time t and zero otherwise; and $D_{it}^{coveredban}$ is a dummy variable that is equal to one if even covered short sales are forbidden in the underlying stock of option class i at time t and zero otherwise (that is, $D_{it}^{coveredban}$ is equal to one when all short sales, both naked and covered short sales, are forbidden and zero otherwise). Coefficients of interest are β_1 and β_2, which represent the respective estimated effects of naked bans and covered bans on Y_{it}. Estimation of this specification yields the exact same results as when estimating the following specification: $Y_{it} = \beta_0 + \alpha_i + \beta_1 D_{it}^{ban} + \beta_2 D_{it}^{coveredban} + \gamma_t + (\theta X_{it}) + \varepsilon_{it}$, where D_{it}^{ban} is defined as in the text.

by excluding the endogenous event (however, the results may still suffer from the removal of bans being endogenous).

Another potential endogeneity problem is that most of the bans included in our dataset were adopted at a time when the situation on the financial markets all around the world was very turbulent. For example, most of the bans were adopted within a week after the bankruptcy of Lehman Brothers. In as far the financial turmoil per se resulted in increased option trading activity (e.g. by attracting more speculators) we run the risk of obtaining biased results by assigning this effect to the concomitant short sale bans. We deal with the potential problem caused by market-wide developments on the financial markets by allowing for weekly time fixed effects, which will take out broad market moves. In addition, we deal with both potential endogeneity problems by presenting graphic illustrations of the changes in option trading activity around the removal of the bans, the intuition being that the financial turbulence was not as strong at the time of the lifting of the bans and that the lifting of the bans were more random than were the adoption of the bans.

5. Results

5.1 Descriptive Results

Table 3 reports average daily trading volume and open interest for an average option class separately for the ban periods and non-ban periods. As seen in the table, the average daily trading volume of put options is higher during the ban periods (4,447 contracts) compared to the non-ban periods (3,920 contracts). The bans also seem to be associated with an increase in the trading volume of call options (6,320 contracts for the ban periods vs. 4,064 contracts for the non-ban periods). Similar indications are given for open interest of both puts and calls. Note that we make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered).

Next, we examine the impact of short sale bans on option trading activity graphically. We compare the average level of the relevant measure for each day during a 41 day window ranging from 20 days before the introduction date up to 20 days after the introduction date with the average level of the measure for all non-ban observations except those belonging to the 20 days preceding the introduction of the bans (henceforth called *the control group*).[35]

Figure 6 illustrates the level of trading volume of put options around the introduction of the bans. There is a slight increase in the trading volume of puts (relative to the control group) from an average of 743 contracts for the 20 days preceding the introduction of the bans (green horizontal line) to 1,815 contracts for the 20 days following the introduction date (red horizontal line). However, from the look of the graph, it is hard to assign any increase in option trading volume to the introduction of the bans, as the

[35] A more detailed description of how the graphs are created is given in the methodology section ("4.2. Methodology").

trading volume starts to rise about 10 days prior to the introduction date. Moreover, the trading volume of puts reaches its highest point a couple of days *before* the introduction date and then drops substantially during the days that follow.

The noticeable peak a couple of days before the introduction of the bans is a bit puzzling. One potential explanation to why we observe the peak is that most of the bans in our sample were adopted within a week after the bankruptcy of Lehman Brothers. However, to the extent the bankruptcy of Lehman had a market-wide impact on option trading activity (e.g. by attracting more speculators), this impact should be captured by the time fixed effects (week dummies). Moreover, September 19 was the last trading day in options expiring on September 20 (which was a monthly expiration date), and many countries included in our sample adopted their bans shortly after this date (see Figure 2 and Table 1). In untabulated results, we find that the trading volume of puts is higher during days preceding option expiration dates. More specifically, the trading volume is on average 767 contracts higher during the three days preceding an expiration date compared to all other days, and the result is significant at the one percent level. However, the results illustrated in Figure 6 remain unchanged when we control for expiration of options, and hence the peak is likely to have another cause.[36] Furthermore, the fact that the trading volume of puts starts to rise about 10 days before the introduction of the bans may indicate that the adoption of the bans is endogenous. Financial regulators are likely to have introduced the bans as a reaction to country specific time varying variables related to the crisis, and the endogeneity problem emerges when these variables are correlated with option trading activity.

In Figure 7 we do a similar graphical analysis for open interest of put options. In line with the results presented in Figure 6, open interest of puts (relative to the control group) is on average higher during the 20 days following the introduction date (60,631 contracts) than for the 20 days preceding the introduction of the bans (30,797 contracts). However, as in the case with trading volume we see signs of pre-trends, and it is hence problematic to assign any effect to the introduction of the short sale bans. The sharp decline around the introduction date coincides with the expiration of options for several of the countries included in the sample.

Furthermore, in Figure 8 and 9 we examine the level of trading volume and open interest of put options around the *removal* of the bans. As mentioned in the methodology section, this is done in order to mitigate potential endogeneity problems related to the adoption of the bans. As in the case with ban inception dates, we see no clear changes in option trading activity that can be assigned to the removal of the bans. The noticeable drop in open interest 13-14 days after the lifting date coincides with the expiration of options for several of the countries included in the sample. The results illustrated in Figure 8 do not change when we control for option expiration dates.

[36] We control for the expiration of options by including a dummy variable that is equal to one for the three days preceding an option expiration date and zero otherwise to the list of explanatory variables. When including the expiration dummy in our standard specification, the estimated coefficient for the dummy equals 767 and it is significantly different from zero at the one percent level.

The graphical results provide no strong evidence that short sellers attempted to circumvent the bans by migrating to the option markets. Even though trading volume and open interest of put options on average is higher during the period after (before) the introduction (removal) of the bans, it is hard to assign the results to the bans.

5.2 Regression Analysis

In order to examine the entire sample period more thoroughly, formalize the results obtained in the graphical illustrations and allow for different effects of naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), we now turn to the regression analysis. We start out by focusing on the impact of the short sale bans on option trading volume. We then turn to open interest and the composition of option trading activity. Subsequently, we test the robustness of the results by introducing control variables representing time-variant characteristics of the underlying stock. Finally, as a further robustness test, we rerun the regressions for various subsamples of countries, option classes and time periods. All results from the regression analysis presented here are significant at the five percent level (at least) if not stated otherwise.

5.2.1 Option Trading Volume

We start out by focusing on the impact on option trading volume, as measured by the daily number of traded put and call contracts. When analyzing the entire sample, the coefficient of the *Ban* dummy in column 2 of Table 5, panel C, indicates that the bans on short sales (irrespective of type) resulted in increased overall trading volume of options. The coefficient suggests that the bans are associated with an increase in the daily number of traded options by 2,905 contracts. In column 1, panel C, where we allow for different impacts of naked and covered bans, the results suggest that bans on naked short sales are associated with an increase in the daily number of traded options by 3,400 contracts. Interestingly, covered bans have no significant impact (even though the coefficient is positive).

In panel A and B of Table 5 we investigate the impact on the daily number of traded contracts for put options and call options separately. The coefficient of the *Ban* dummy in column 2, panel A and B, suggests that the bans (irrespective of type) have a statistically significant positive impact on both the number of traded puts and the number of traded calls. The bans (irrespective of type) are associated with an increase in the number of traded puts by 1,117 contracts and an increase in the number of traded calls by 1,788 contracts. Again, the results indicate that naked and covered bans have different impacts, the difference being larger for puts than for calls. The coefficient of the *Naked ban* dummy shown in column 1, panel A and B, indicates that naked bans are associated with an increase in the number of traded puts by 1,580 contracts, whereas covered bans do not seem to have any impact at all. Naked bans are associated with an increase in the number of traded calls by 1,820 contracts, while covered bans are associated with an increase in the number of traded calls by 1,713 contracts (significant at the 10 percent level). To get a feeling of the economic significance of these results, the average daily number of traded put and call contracts per option class during the pre-ban and post-ban periods is 3,920 and 4,064 contracts respectively (see Table 3).

When rerunning the regressions with a log-level functional form, we obtain significant coefficients for the ban dummies only for call options (see column 5 and 6 of Table 5, panel B). The results suggest that the short sale bans (irrespective of type) are associated with an increase in the daily number of traded call contracts by 19.8 percent. Bans on naked short sales are associated with an increase in the daily number of traded call contracts by 23.1 percent, while there is no evidence that covered bans had any impact at all.

5.2.2 Open Interest

We investigate the impact of the short sale bans on option trading activity further by examining the effect on open interest, which is the number of option contracts outstanding on a given day. In Table 6, panel A, B and C, the dependent variable is open interest of puts, calls and all options (put+calls) respectively. The results are in line with those obtained when using trading volume as dependent variable. This is expected since the correlation between open interest and the daily number of traded contracts is high (0.72 for puts and 0.60 for calls). For puts, the bans on short sales (irrespective of type) are associated with an increase in the number of outstanding puts by 39,827 contracts (see column 2, panel A), whereas the corresponding figure for calls is 64,151 contracts (see column 2, panel B). In column 1, where we allow for different impacts of naked and covered bans, bans on naked short sales are seen to increase the number of outstanding puts by 48,901 contracts and the number of outstanding calls by 55,489 contracts. Consistent with the results obtained when using trading volume as dependent variable, covered bans do not have a statistically significant impact on the open interest of put options. For calls, however, the coefficient for the *Covered ban* dummy, displayed in column 1, panel B, is significantly positive at the 10 percent level and corresponds to an increase in the number of outstanding calls by 85,848 contracts during the ban period. When rerunning the regression with a log-level functional form we fail to obtain significant coefficients for all ban dummies, both for put and call options (see column 5 and 6, panel A, B and C). The average number of outstanding contracts per option class during the pre-ban and post-ban periods is 208,481 contracts for puts and 211,060 contracts for calls.

The results so far suggest that the bans are associated with an increase in option trading volume and that this increase can partially be assigned to an increase in the number of outstanding option contracts. Furthermore, the results indicate that the impact on option trading activity is greater for naked bans than for covered bans, especially for put options. This finding is a bit surprising. By prohibiting all short sales, both naked and covered short sales, covered bans are stricter than are naked bans. Intuitively, covered bans should therefore have an impact on the demand for options at least as large as the impact of naked bans. There are, however, several potential explanations to why we obtain these results. First of all, some countries that adopted covered bans imposed increased delivery requirements for market makers, and the exemptions for market makers in some covered regimes were not as extensive as in countries only banning naked short sales.[37] Secondly, theoretical models and results of previous empirical work suggest that stricter short sale restrictions have a more detrimental effect on the liquidity and market quality on the stock market. Thus, a potential explanation to the observed difference between the impact of naked and covered bans on option trading activity is that covered bans may have resulted in higher inventory holding costs (including higher hedging costs) and adverse selection costs for option market makers than did naked bans, and hence covered bans may have had a more detrimental effect on the supply of liquidity in options on banned stocks.

5.2.3 Composition of Option Trading Volume

The results presented above already indicates that the short sale bans had a larger economic impact on the daily number of traded call contracts than on the daily number of traded put contracts. This observation is confirmed by the results presented in column 1 and 2 of Table 7, where the dependent variable is a ratio of the daily number of traded put contracts over the total daily number of traded option contracts (puts+calls). The coefficient of the *Ban* dummy in column 2 indicates that the short sale bans (irrespective of type) are associated with a decrease in the fraction of put contracts by 4.5 percentage points. In column 1, where we allow for different impacts of naked and covered bans, the coefficients indicates that both types of bans are associated with a statistically significant decrease in the fraction of put contracts (even though the estimated effect is greater for covered bans). To get a feeling of the economic significance of these results, the average fraction of overall option trading volume that originates from put options is 46.9 percent during the pre-ban and post-ban periods.

5.1.3 Introduction of Control Variables

In column 3 and 4 of Table 5, panel A, B and C, we investigate whether the results so far are robust to the introduction of stock-level time varying control variables. We control for the following characteristics of the underlying stock: return, bid-ask spreads and the standard deviation of return based on the previous 10 observations.[38] All three variables have been pointed out as being correlated with the adoption of short sale bans by previous empirical work. The reason why we choose to control for these variables is that we want to (i) capture the direct effect of the short sale bans on option trading activity, that is, changes in option trading activity due to investors trying to find alternative ways to profit from declining prices, rather than impacts resulting from the bans having an effect on the underlying market

[37] In, Switzerland the prohibitions applied to all market participants. In the U.S., the temporary rule (Rule 204T), which was announced on September 17, 2008 (two days before the announcement of the ban prohibiting all short sales in specified financial stocks), imposed enhanced delivery requirements on sales of all equity securities and limited market makers ability to hedge by penalizing failure to deliver. In Canada, the order did not distinguish between different types of market participants.

[38] All control variables are expressed in decimals.

(other than prohibiting short sales), and (ii) mitigate the potential endogeneity problem related to the adoption of the bans (further described in the methodology section).[39]

When the extended specification is estimated, the signs, significance levels and magnitude of the ban dummies remain the same as in the standard specification estimated in column 1 and 2. In line with theoretical predictions, the estimated coefficient of the bid-ask spread of the underlying stock is estimated to be negative and it is significantly different from zero at the one percent level for puts as well as for calls (see column 3 and 4 of Table 5, panel A and B). Moreover, in line with the intuition that investors buy puts when prices fall and calls when prices rise, the coefficient of the return of the underlying stock shown in column 3 and 4, panel A and B respectively, is estimated to be negative for puts and positive for calls (however, it is only significantly different from zero in the former case). Finally, in contrast to the theory that higher volatility raises the inventory holding costs of market makers, the coefficient of the rolling standard deviation of return based on the previous 10 observations is positive and significant for puts and negative but insignificant for calls.[40] However, the economic significance is questionable for all control variables. For example, an increase in the bid-ask spread of the underlying stock by one percentage point is associated with a decline in the daily number of traded puts by 33 contracts and a decline in the daily number of traded calls by 76 contracts (when we allow for different impacts of naked and covered bans), which are rather small declines in comparison to the average daily trading volume during the pre-ban and post-ban periods of 3,920 contracts for puts and 4,064 contracts for calls. The average bid-ask spread in our sample is 0.34 percent, which means that an increase of the spread by one percentage point is quite substantial.

Neither does the introduction of the three control variables alter the signs, significance levels nor the magnitude of the estimated coefficients of the ban dummies in column 3 and 4 of Table 6, panel A, B and C, where we measure option trading activity as open interest. Of all three control variables, the bid-ask spread of the underlying stock is the only one that is estimated to have a significant impact on open interest. A one percentage point increase in the bid-ask spread is associated with a decrease in the open interest of puts by 1,343 contracts and a decrease in the open interest of calls by 1,749 contracts (when we allow for different impacts of naked and covered bans). Also the results regarding the impact of the bans on the composition of option trading volume are unaffected by the introduction of the control variables, which can be seen in column 3 and 4 of Table 7.

The results in this subsection confirm that option trading activity, as measured in terms of option trading volume and open interest, was higher during ban periods than during non-ban periods. As the results are robust to the introduction of stock-level time varying variables that are likely to reflect variables that

[39] The introduction of variables reflecting time varying characteristics of the underlying stock that are correlated with the adoption of short sale bans is an indirect way to control for country specific time varying variables that could potentially be correlated with both the adoption of short sale bans *and* option trading activity.

[40] However, the fact that we do not obtain a negative and significant coefficient for the volatility of the return of the underlying stock is not surprising, as return volatility is also likely to be positively correlated with the demand for options.

affect the supply of liquidity in options, there is reason to believe that much of the observed increase in option trading activity can be explained by an increased demand for options on banned stocks. However, even after having controlled for time varying characteristics of the underlying stock that are likely to affect option market makers willingness to supply liquidity, naked bans are still estimated to have a greater impact on the trading volume and open interest of put options than are covered bans. This finding is a bit puzzling and indicates that (i) naked bans are associated with a larger increase in the *demand* for put options than are covered bans, or (ii) our control variables do not effectively control for variables that affect the supply of liquidity in options. Furthermore, to the extent the control variables pick up country specific time varying variables that are correlated with both the adoptions of short sale bans *and* option trading activity, the robustness of the results reduces the likelihood that the results so far are biased due to the adoption of bans being endogenous.

5.1.4 Subsample Analyses

As a further robustness test we rerun the regressions for various subsamples of countries, option classes and time periods. More specifically, we estimate the standard specification for a subsample where all option classes of U.S. stocks are excluded, as well as for a subsample where the ten largest option classes in terms of average daily number of traded contracts are excluded. Furthermore, we re-estimate the standard specification for a sample period starting in December 2008 (rather than in January 2008), when all countries in our data set had adopted bans on short sales. Finally, we rerun the regressions for a subsample only containing option classes of stocks listed in countries that only adopted bans on naked short sales, and another subsample of option classes of stocks listed in countries that only adopted bans on all short sales (both naked and covered short sales).

5.1.4.1 Exclusion of Options on Non-U.S. Stocks

As pointed out in the data section, 91 of the 171 pairs of option classes in the main dataset are listed on an U.S. exchange, and these pairs represent more than half of the total sample in terms of the number of observations. The announcement of the U.S. ban prohibiting all short sales (both naked and covered short sales) in 797 financial stocks on September 19, 2008 coincided with the announcement of the TARP. If the announcement of the TARP had an impact on option trading activity by itself, this impact will automatically be assigned to the ban due to their concomitant announcements. As options on U.S. stocks represent a majority of the number of observations this could in turn result in biased results. A subsample where all option classes of U.S. stocks are excluded allows us to check whether the results presented in previous sections are not simply a reflection of a differential between the impact of the short sale bans on trading activity for options on U.S. stocks and options on stocks listed in the rest of the world.

As seen in column 1, 2, 3 and 4 of Table 8, panel A and B, most of the results obtained when using the entire sample remains unchanged when options on U.S. stocks are excluded. The estimates of the coefficients of the ban dummies generally remain sizeable and significantly different from zero in the specifications examining the impact of the bans on the number of traded puts, number of traded calls and open interest of puts. Some of the coefficients are smaller in absolute value, but this probably reflects the fact that the average daily number of traded contracts of U.S. options is higher than for options on stocks listed in other countries. There are, however, two main differences in the results. First, the estimated coefficients of the ban dummies in column 3 and 4, panel B, where the dependent variable is open interest of calls, all lose their significance. Secondly, there is no longer evidence that the bans had an impact on the fraction of overall option trading volume that originates from put options (see panel C). These results indicate that the estimated impact of the bans on the fraction of overall option trading volume that originates from put options obtained in earlier estimations can be assigned mainly to options on U.S. stocks.

5.1.4.2 Exclusion of Options with High Trading Volume

As seen in Figure 4, option classes of a handful of stocks represent a major fraction of the aggregated average daily trading volume in our sample. The second subsample analysis, in which we exclude the ten largest options classes in terms of average daily number of traded contracts, allows us to check whether the results obtained when using the entire sample is simply not a reflection of the bans having a large impact on "large" option classes and a small or no impact on "small" option classes.[41] The resulting estimates of the coefficients of the ban dummies, shown in column 5, 6, 7 and 8 of Table 8, panel A, B and C, are quite interesting. First, the coefficients become considerably smaller, which is probably a reflection of the average trading volume being considerably smaller when the ten "largest" option classes are excluded from the sample. Maybe more interesting is that the statistical significance changes dramatically. As seen in column 5 and 6, panel A and B, the estimated impact of short sale bans (irrespective of type) on the daily number of traded put contracts and the open interest of puts is no longer statistically significant, nor is the estimated impact of naked bans. In column 7 and 8, Panel A, on the other hand, where the dependent variable is the daily number of traded call contracts, the coefficients of the *Ban* dummy and the *Covered ban* dummy are still significant, whereas naked bans are no longer estimated to have a significant impact on the number of traded call contracts. When it comes to open interest of calls (column 7 and 8, Panel B), it is only covered bans that are estimated to have a statistically significant impact. In column 3 and 4, panel C, where the dependent variable is the daily number of traded put contracts over the total daily number of traded option contracts (puts+calls) we see that the signs, magnitude and statistical significance of the coefficients of the ban dummies are unaffected.

From these results, we draw the conclusion that much of the results obtained in previous sections can be assigned to option classes with high trading volume. In a sense, these findings are bad news as they

[41] For put options, the following firms are excluded (in order of size): Citigroup Inc. (the U.S.), Allianz SE (Germany), Bank of America Corporation (the U.S.), General Electric CO. (the U.S.), ING Groep N.V. (the Netherlands), UBS AG (Switzerland), Bank AG (Germany), J. P. Morgan Chase & Co. (the U.S.), Schweiz Ruckversicherungs-Gesellschaft AG (Switzerland) and Wells Fargo & CO New (the U.S.). For call options, the following firms are excluded (in order of size): Citigroup Inc. (the U.S.), Bank of America Corporation (the U.S.), Allianz SE (Germany), General Electric CO. (the U.S.), ING Groep N.V. (the Netherlands), J. P. Morgan Chase & Co. (the U.S.), UBS AG (Switzerland), Schweiz Ruckversicherungs-Gesellschaft AG (Switzerland), Goldman Sachs Group Inc (the U.S.) and Deutsche Bank AG (Germany).

indicate that we have not managed to mitigate the potential endogeneity problem related to the adoption of the bans. The excluded option classes belong to stocks of several of the major financial institution around the world. It is likely that these firms were relatively more exposed to time varying variables related to the crisis (such as house prices, the degree of stability in the financial system, etc) and that potential crisis-related problems of these firms received relatively more publicity compared to firms remaining in the sample. Hence, it is likely that the excluded firms attracted relatively more speculators. To put in other words, the correlation between country specific time varying variables and option trading activity is likely to have been higher for the excluded firms than for the firms remaining in the sample. Furthermore, it is likely that financial regulators put a relatively high weight on the financial condition and speculative attacks of short sellers in the stocks of these firms when deciding whether and when they should impose bans on short sales. However, it may also be the case that the underlying stocks of the excluded option classes were more liquid than the underlying stocks of option classes remaining in the sample (even under the short sale bans), and hence market makers may have been more willing to provide liquidity in these options.

5.1.4.3 Shorter Sample Period

In a way to further test the robustness of our results, we re-estimate the standard specification for a sample period ranging from December 1, 2008 to February 28, 2011 (rather than from January 1, 2008 to February 28, 2011). In December, all countries in our sample had adopted short sale bans. This means that we rely solely on the removal of the bans in our difference-in-difference re-estimation, and hence this procedure eliminates the potential endogeneity problem related to the adoption of the bans by excluding the endogenous event.[42]

The results from this subsample analysis are shown in Table 9. We see that the signs of all estimated coefficients remain unchanged compared to when estimating the standard specification for the entire sample period (shown in Table 5, 6 and 7). The major difference is that the statistical significance of the estimated impacts is reduced. For trading volume of put options, the estimated effect of bans (irrespective of type) and bans on naked short sales are still significant, but only at the 10% level. The estimated impact of bans (irrespective of type) and covered bans on the trading volume of call options are also still significant, but again only at the 10% level. For open interest, the coefficients of the ban dummies all lose their significance for both puts and calls. Bans (irrespective of type), naked bans and covered bans are all still estimated to reduce the proportion of overall option trading volume that originates from put options, but the estimated coefficient of the *Naked ban* dummy is no longer significantly different from zero.

There are at least two potential explanations why the statistical significance declines. First of all, the results from the estimations where we use the entire sample period may be biased. The adoption of the bans may be correlated with some country specific time varying variable that is also correlated with

[42] This robustness check hinges on the assumption that the removal of the bans is more random than is the adoption of the bans.

option activity. Hence, in the results presented in Table 5 through 7, changes in option trading activity have been assigned to the bans, even though it may have changed due to some other reason. Secondly, when we exclude all observations belonging to the first 11 months of our sample, we exclude a substantial fraction of all non-ban observations and a major share of the number of changes from 0 to 1 and from 1 to 0 in the ban dummy variables. Hence, the difference-in-difference estimation is not as "sharp" as when all observations are included in the regressions.

5.1.4.4 Did Naked Bans and Covered Bans have Different Effects?

In an attempt to further investigate the surprising result that naked bans generally seem to have larger and more statistically significant impacts on option trading activity of put options than do covered bans, we rerun the regressions for another two subsamples. The first subsample only includes options classes of stocks listed in countries that only adopted bans on naked short sales, and the second subsample only includes option classes of stocks listed in countries that only adopted bans on all short sales (both naked and covered short sales). Option classes of stocks that were subject to both types of bans but at different points in time are excluded from both subsamples.[43]

The results from the re-estimations are shown in Table 10 and they provide (weak) support that naked bans and covered bans have different effects on option trading activity, especially for put options. The results obtained so far suggest that both naked and covered bans generally have a positive impact on option trading activity, but in several of the specifications (especially those focusing on puts) it is only the coefficient of the *Naked ban* dummy that is significantly different from zero. In contrast to most of the results obtained so far, covered bans are now estimated to have a *negative* impact on all measures of option trading activity (except for open interest of calls). However, the coefficient is only statistically significant for trading volume of puts. More specifically, covered bans are associated with a decrease in the daily number of traded puts by 274 contracts (see column 3, panel A). In line with previous results presented in table 5 through 7, the coefficient of the *Naked ban* dummy is positive in most specifications, but it is only significantly different from zero in column 1, panel A, where the dependent variable is the daily number of traded put contracts (significant at the 10 percent level).

In Table 10, panel C, where the dependent variable is the proportion of overall option trading volume that originates from put options, the coefficients of both ban dummies are estimated to be negative, which is in line with the results from the estimation of the standard specification in Table 7. However, it is only the coefficient of the *Covered ban* dummy that is estimated to be significantly different from zero.

Notice that a major fraction of the number of observations, option classes and countries are excluded from this subsample analysis as a majority of the stocks in our dataset were subject to both types of bans but at different points in time. This may partially explain the observed differences between the results

[43] Countries in our main sample that only adopted bans on naked short sales are Austria, France, Germany and Spain. Countries that only adopted bans on all short sales (both naked and covered short sales) include Canada and the U.K.

reported in this section and the results obtained when estimating the version of the standard specification that allows for different impacts of naked and covered bans.

6. Conclusion

In this thesis, we study the bans on short sales that were adopted around the world in connection with the financial crisis of 2007-2009. With few exceptions, the bans only prohibited short sales of stocks and not trading in options. Using option and stock data from 10 countries that adopted bans on short sales at some point in time from January 1, 2008 through February 28, 2011 and fixed effect panel data techniques, we examine how the short sale bans affected the option markets. More specifically, we examine whether short sellers used the option markets as a substitute for direct short selling during the ban periods, and whether they did so primarily by trading put options or by trading call options.

To the extent the bans induced short sellers to move to the option markets, we would expect to see an increase in the demand for options, and hence an increase in both option trading volume and open interest of options. However, it is also likely that the bans reduced option market makers' ability to hedge, increased their hedging costs and made it more expensive for them to carry inventory and trade with informed investors, and hence reduced their willingness to provide liquidity in options on banned stocks. The net effect of the bans on option trading activity is thus ambiguous.

In contrast to previous empirical work, we find that the short sale bans are associated with a statistically and economically significant increase in option trading volume and open interest for both put and call options, especially for options with high trading volume. When allowing for different effects of naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), we find that naked bans generally are associated with an increase in option trading activity, while the results for covered bans are more ambiguous. Furthermore, we find that the bans are associated with a decrease of the fraction of overall option trading volume that originates from put options, especially for options on stocks listed in the U.S.

Our results suggest that the short sale bans were accompanied by an increase in the demand for options on banned stocks, and that covered bans reduced the supply of liquidity in options on banned stocks more than did naked bans. Furthermore, the results indicate that the demand for calls increased more than the demand for puts, which in turn suggests that investors speculated on declining stock prices by writing calls rather than buying puts, or that investors viewed the low price levels prevailing during the ban periods as an opportunity to enter long positions at favorable prices. However, graphical results and various subsample re-estimations indicate that the adoption of the bans is endogenous. Hence, it is hard to assign the observed effects to an attempt by short sellers to circumvent the bans. The ambiguous results suggest that further studies focusing on the impact of short sale bans on the option markets are needed to enhance our understanding of the effectiveness of bans on short sales of stocks.

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8. Appendix

Table 1. Overview of the Dataset

Country	Ban start date: Naked	Ban lift date: Naked	Ban start date: Covered	Ban lift date: Covered	Scope of ban	Number of pairs of options classes	Fraction of total option classes	Duration (days)[6]	Obs.	Fraction of total obs.	Obs. with naked ban	Fraction of obs.	Obs. with covered ban	Fraction of obs.	Obs. with ban	Fraction of obs.
Austria	2008-10-27	still banned			fin	4	2.3%	855	3,147	2.7%	2,319	73.7%	0	0.0%	2,319	73.7%
Canada			2008-09-19	2008-10-08	fin	7	4.0%	20	4,844	4.2%	0	0.0%	98	2.0%	98	2.0%
France	2008-09-22	2011-02-01			fin	10	5.8%	863	7,408	6.4%	5,553	75.0%	0	0.0%	5,553	75.0%
Germany[1]	2008-09-20	2009-05-31			fin	7	4.0%	254	5,617	4.8%	2,600	46.3%	0	0.0%	2,600	46.3%
	2010-05-19	still banned			fin	7	4.0%	286								
Italy[2]	2008-09-23	2009-07-31	2008-10-01	2009-05-31	fin, then all	11	6.4%	312	8,474	7.3%	561	6.6%	1,813	21.4%	2,374	28.0%
Netherlands[3]	2008-09-22	2008-10-04	2008-10-05	2009-06-01	fin	4	2.3%	253	3,244	2.8%	40	1.2%	660	20.3%	700	21.6%
Spain	2008-09-24	still banned			all	8	4.6%	888	6,363	5.5%	4,944	77.7%	0	0.0%	4,944	77.7%
Switzerland[4]			2008-09-19	2009-01-16	fin	7	4.0%	120	8,655	7.4%	7,089	81.9%	480	5.5%	7,569	87.5%
	2008-09-19	still banned			all	17	9.8%	893								
UK			2008-09-19	2009-01-16	fin	12	6.9%	120	8,350	7.2%	0	0.0%	927	11.1%	927	11.1%
UnitedStates[5]	2008-07-21	2008-08-12			fin	7	4.6%	23	60,426	51.9%	52,302	86.6%	1,108	1.8%	53,410	88.4%
	2008-09-18	still banned	2008-09-19	2008-10-08	naked-all, covered-fin	91	52.6%	894								
Total						171			116,528	100.0%	75,408		5,086		80,494	

[1] The ban adopted in Germany on May 19, 2010, applied to the same stocks as the ban ranging from September 20, 2008 to May 31, 2009.

[2] In Italy, the financial regulator initially banned naked short sales in financial stocks (imposed on September 23, 2008). On October 1, the ban was extended to include also covered short sales in financial stocks, and on October 10, the ban was extended further to include all short sales (both naked and covered short sales) in all stocks. On December 31, 2008, the stringency of the ban on short sales in non-financial stocks was reduced to prohibit naked short sales only, and on May 31, 2009, the stringency of the ban on short sales in financial stocks was reduced to prohibit naked short sales only.

[3] On September 22, 2008, the Netherlands imposed a ban on naked short sales in financial stocks, and two weeks later, on October 4, the ban was replaced by a ban on all short sales (both naked and covered short sales).

[4] On September 19, 2008, Switzerland adopted a ban on all short sales (both naked and covered short sales) in specified financial stocks and a ban on naked short sales in all other stocks.

[5] The seven stocks in our sample that were subject to the ban on naked short sales in 19 financial stocks imposed by the U.S. SEC on July 21, 2008, were also subject to the ban on naked short sales in all U.S. stocks imposed on September 18, as well as the ban on all short sales (both naked and covered short sales) in 797 financial stocks imposed on September 19.

[6] Number of days with restrictions on short sales (no difference is made between naked and covered bans).

Table 2. Summary Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
vm_put	116,528	4,284	16,749	0	960,433
vm_call	116,528	5,622	36,447	0	4,924,182
vm_all	116,528	9,906	48,877	0	5,263,235
vm_put-to-option	106,075	0.444	0.279	0.000	1.000
oi_put	116,526	220,580	624,688	0	9,522,233
oi_call	116,526	256,321	951,197	0	20,100,000
oi_option	116,526	476,902	1,541,719	0	28,200,000
vm_undrl	116,528	17,241	62,681	0	3,772,638

Table 2 reports summary statistics for our dependent variables and the trading volume of the underlying stock. *Vm_put* is the trading volume of put options, measured as the daily number of traded put contracts (each contract covers 100 shares); *vm_call* is the trading volume of call options, measured as the daily number of traded call contracts; and *vm_all* is the sum of the two (i.e. *vm_put+vm_call). Vm_put-to-option* is a ratio of the daily number of traded put contracts over the daily number of traded option contracts (i.e. *vm_put/vm_all*). *Oi_put* is the open interest of put options, measured as the number of outstanding put contracts on a given day; *oi_call* is the open interest of call options, measured as the number of outstanding call contracts on a given day; and *oi_all* is the sum of the two (i.e. *oi_put+oi_call). Vm_undrl* is the trading volume of the underlying stock, measured in thousands of shares. Each option contract covers 100 shares.

Table 3. Daily Averages for an Average Option Class for the Ban Periods and Non-Ban Periods

	dummy	vm_put	vm_call	vm_all	vm_put-to-all	oi_put	oi_call	oi_all
Naked ban	0	3,969	4,155	8,125	0.468	211,285	220,584	431,869
	1	4,435	6,324	10,759	0.432	225,029	273,425	498,455
Covered ban	0	4,297	5,656	9,953	0.444	220,829	254,980	475,810
	1	4,001	4,882	8,883	0.450	215,124	285,703	500,827
Ban	0	3,920	4,064	7,984	0.469	208,481	211,060	419,542
	1	4,447	6,320	10,767	0.433	225,997	276,583	502,579

Table 3 reports the average daily trading volume, open interest and put-to-option ratio for an average option class separately for the ban periods and non-ban periods. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise. *Vm_put* is the trading volume of put options, measured as the daily number of traded put contracts (each contract covers 100 shares); *vm_call* is the trading volume of call options, measured as the daily number of traded call contracts; and *vm_all* is the sum of the two (i.e. *vm_put+vm_call). Vm_put-to-option* is a ratio of the daily number of traded put contracts over the daily number of traded option contracts (i.e. *vm_put/vm_all*). *Oi_put* is the open interest of put options, measured as the number of outstanding put contracts on a given day; *oi_call* is the open interest of call options, measured as the number of outstanding call contracts on a given day; and *oi_all* is the sum of the two (i.e. *oi_put+oi_call).* Each option contract covers 100 shares.

Table 4. Monthly Statistics

ym	vm_put	vm_call	vm_all	vm_put-to-option	oi_put	oi_call	oi_all	vm_undrl
2008m1	6,097	5,295	11,392	0.527	279,143	222,162	501,305	20,254
2008m2	5,996	5,635	11,631	0.498	313,703	268,440	582,143	17,692
2008m3	6,073	5,240	11,313	0.502	310,172	273,857	584,029	21,169
2008m4	5,205	5,922	11,127	0.463	306,477	263,011	569,488	15,548
2008m5	4,285	7,337	11,622	0.476	223,725	205,670	429,395	15,143
2008m6	4,326	3,226	7,552	0.515	185,223	177,134	362,357	13,291
2008m7	4,431	3,757	8,188	0.485	206,010	199,034	405,044	16,980
2008m8	3,149	2,966	6,115	0.466	225,460	219,635	445,095	11,024
2008m9	5,110	4,868	9,978	0.507	229,593	227,956	457,550	19,993
2008m10	4,483	4,648	9,131	0.484	224,457	236,322	460,778	17,654
2008m11	3,644	4,013	7,658	0.467	216,241	260,081	476,322	15,315
2008m12	2,601	3,052	5,653	0.434	182,252	234,179	416,431	11,456
2009m1	4,084	4,081	8,165	0.446	148,185	174,110	322,294	17,623
2009m2	4,350	4,733	9,083	0.460	170,849	184,871	355,720	21,745
2009m3	5,157	6,208	11,365	0.424	185,976	206,604	392,580	27,547
2009m4	5,000	6,794	11,794	0.408	208,221	232,516	440,737	23,500
2009m5	4,515	6,037	10,552	0.423	236,079	259,326	495,405	21,307
2009m6	3,797	4,331	8,128	0.431	235,245	263,851	499,096	16,293
2009m7	3,567	4,763	8,330	0.425	231,585	259,755	491,340	16,224
2009m8	4,124	6,738	10,861	0.412	242,639	281,779	524,418	23,337
2009m9	3,932	5,399	9,330	0.429	235,830	277,079	512,909	20,030
2009m10	3,726	5,193	8,919	0.440	232,087	276,892	508,979	16,914
2009m11	3,197	3,634	6,831	0.443	245,988	291,834	537,823	13,614
2009m12	2,692	3,579	6,271	0.434	219,621	266,454	486,075	15,585
2010m1	3,473	4,511	7,984	0.427	173,363	218,318	391,680	16,634
2010m2	3,142	3,603	6,745	0.447	184,459	220,088	404,548	16,224
2010m3	2,937	4,027	6,964	0.412	186,849	225,621	412,469	15,917
2010m4	3,447	5,597	9,044	0.417	197,988	249,859	447,847	20,469
2010m5	4,776	5,148	9,925	0.474	248,747	308,936	557,683	27,168
2010m6	2,660	2,998	5,658	0.441	240,537	303,914	544,452	18,607
2010m7	3,297	4,851	8,147	0.412	214,962	268,829	483,791	15,563
2010m8	4,503	5,796	10,299	0.442	231,318	300,850	532,168	12,481
2010m9	4,537	7,992	12,529	0.427	231,422	299,682	531,104	12,727
2010m10	6,415	8,630	15,045	0.435	238,874	307,064	545,938	14,492
2010m11	5,914	9,091	15,005	0.454	260,939	333,532	594,471	14,648
2010m12	4,889	10,658	15,548	0.414	239,428	324,576	564,004	13,776
2011m1	7,090	13,286	20,377	0.419	213,849	287,885	501,734	17,087
2011m2	5,892	9,183	15,075	0.414	208,313	268,675	476,988	14,856

Table 4 reports the average daily trading volume, open interest and put-to-option ratio for an average option class (stock) for each month throughout the sample period. *Vm_put* is the trading volume of put options, measured as the daily number of traded put contracts (each contract covers 100 shares); *vm_call* is the trading volume of call options, measured as the daily number of traded call contracts; and *vm_all* is the sum of the two (i.e. *vm_put+vm_call*). *Vm_put-to-option* is a ratio of the daily number of traded put contracts over the daily number of traded option contracts (i.e. *vm_put/vm_all*). *Oi_put* is the open interest of put options, measured as the number of outstanding put contracts on a given day; *oi_call* is the open interest of call options, measured as the number of outstanding call contracts on a given day; and *oi_all* is the sum of the two (i.e. *oi_put+oi_call*). *Vm_undrl* is the trading volume of the underlying stock, measured in thousands of shares. Each option contract covers 100 shares.

Table 5. The Effect of Short Sale Bans on Option Trading Volume

Regressions of option trading volume on different types of ban dummies and a number of control variables, using an international panel of daily option and stock data from 10 countries that adopted bans on short sales of stocks at some point in time between January 1, 2008 and February 28, 2011. All regressions include stock-level fixed effects and time fixed effects (weekly). In Panel A the dependent variable is the daily number of traded put option contracts; in panel B it is the daily number of traded call option contracts; and in panel C it is the total daily number of traded option contracts (puts+calls). Each option contract covers 100 shares. Daily number of traded put (call) contracts is measured as the sum of all put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. The regressions are run with two different functional forms: level-level and log-level. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise. *Return* is the daily return of the underlying stock, expressed in decimals. *Bid-ask spread* is the percentage bid-ask spread of the underlying stock at closing, expressed in decimals. *SD* is a moving standard deviation of the return of the underlying stock based on the previous 10 observations, expressed in decimals. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics.

Panel A. Dependent Variable: Number of Traded Put Contracts

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable	vm_put	vm_put	vm_put	vm_put	log(vm_put)	log(vm_put)	log(vm_put)	log(vm_put)
Naked ban	1579.827**		1552.421**		0.002		-0.005	
	(2.42)		(2.32)		(0.02)		(-0.05)	
Covered ban	18.191		43.413		-0.188		-0.165	
	(0.04)		(0.09)		(-1.60)		(-1.39)	
Ban		1117.187**		1093.729**		-0.054		-0.054
		(2.08)		(2.00)		(-0.59)		(-0.58)
Bid-ask spread			-3287.178***	-3461.145***			-3.306***	-3.324***
			(-3.29)	(-3.70)			(-16.40)	(-16.32)
Return			-1717.854**	-1704.784**			-0.327*	-0.326*
			(-2.24)	(-2.24)			(-1.86)	(-1.86)
SD			1721.654***	1727.334***			0.594***	0.594***
			(2.69)	(2.65)			(3.03)	(3.02)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,528	116,528	114,733	114,733	116,528	116,528	114,733	114,733
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel B. Dependent Variable: Number of Traded Call Contracts

Dependent variable	(1) vm_call	(2) vm_call	(3) vm_call	(4) vm_call	(5) log(vm_call)	(6) log(vm_call)	(7) log(vm_call)	(8) log(vm_call)
Naked ban	1819.904***		1513.072***		0.231**		0.220**	
	(2.63)		(2.74)		(2.28)		(2.17)	
Covered ban	1712.656*		1716.739*		0.120		0.155	
	(1.90)		(1.91)		(1.05)		(1.34)	
Ban		1788.131***		1574.980***		0.198**		0.200**
		(2.69)		(2.74)		(2.13)		(2.15)
Bid-ask spread			-7630.605***	-7607.125***			-4.345***	-4.352***
			(-2.93)	(-2.99)			(-22.33)	(-22.43)
Return			2956.836	2955.072			0.650*	0.651*
			(1.64)	(1.64)			(1.85)	(1.85)
SD			-435.793	-436.559			0.274*	0.275*
			(-0.31)	(-0.31)			(1.66)	(1.67)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,528	116,528	114,733	114,733	116,528	116,528	114,733	114,733
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel C. Dependent Variable: Total Number of Traded Option Contracts

Dependent variable	(1) vm_all	(2) vm_all	(3) vm_all	(4) vm_all	(5) log(vm_all)	(6) log(vm_all)	(7) log(vm_all)	(8) log(vm_all)
Naked ban	3399.730***		3065.492***		0.096		0.089	
	(2.82)		(2.79)		(1.00)		(0.93)	
Covered ban	1730.846		1760.152		-0.046		-0.016	
	(1.40)		(1.42)		(-0.43)		(-0.15)	
Ban		2905.319***		2668.709***		0.054		0.057
		(2.75)		(2.76)		(0.61)		(0.64)
Bid-ask spread			-10917.783***	-11068.270***			-3.986***	-3.998***
			(-3.16)	(-3.28)			(-20.53)	(-20.50)
Return			1238.981	1250.287			0.239*	0.240*
			(1.03)	(1.03)			(1.96)	(1.97)
SD			1285.861	1290.774			0.393***	0.394***
			(0.82)	(0.82)			(2.63)	(2.63)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,528	116,528	114,733	114,733	116,528	116,528	114,733	114,733
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 6. The Effect of Short Sale Bans on Open Interest of Options

Regressions of open interest on different types of ban dummies and a number of control variables, using an international panel of daily option and stock data from 10 countries that adopted bans on short sales of stocks at some point in time between January 1, 2008 and February 28, 2011. All regressions include stock-level fixed effects and time fixed effects (weekly). In Panel A the dependent variable is the daily open interest of put option contracts; in panel B it is the daily open interest of call option contracts; and in panel C it is the total daily open interest of all option contracts (puts+calls). Each option contract covers 100 shares. Open interest of put (call) contracts is measured as the sum of all outstanding put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace on a specific day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. The regressions are run with two different functional forms: level-level and log-level. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise. *Return* is the daily return of the underlying stock, expressed in decimals. *Bid-ask spread* is the percentage bid-ask spread of the underlying stock at closing, expressed in decimals. *SD* is a moving standard deviation of the return of the underlying stock based on the previous 10 observations, expressed in decimals. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics.

Panel A. Dependent Variable: Open Interest of Put Options

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable	oi_put	oi_put	oi_put	oi_put	log(oi_put)	log(oi_put)	log(oi_put)	log(oi_put)
Naked ban	48901.067**		47003.977**		0.136		0.134	
	(2.22)		(2.17)		(1.35)		(1.33)	
Covered ban	18273.366		18762.162		0.025		0.021	
	(0.81)		(0.84)		(0.21)		(0.18)	
Ban		39826.982**		38419.109**		0.103		0.099
		(2.00)		(1.98)		(1.03)		(0.99)
Bid-ask spread			-134334.51***	-137590.37***			-0.127	-0.140
			(-4.89)	(-5.34)			(-0.92)	(-0.96)
Return			1732.980	1977.607			-0.076***	-0.075***
			(0.16)	(0.19)			(-2.75)	(-2.76)
SD			-20377.313	-20271.248			0.253**	0.253**
			(-0.51)	(-0.51)			(2.13)	(2.13)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,526	116,526	114,732	114,732	116,526	116,526	114,732	114,732
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel B. Dependent Variable: Open Interest of Call Options

Dependent variable	(1) oi_call	(2) oi_call	(3) oi_call	(4) oi_call	(5) log(oi_call)	(6) log(oi_call)	(7) log(oi_call)	(8) log(oi_call)
Naked ban	55488.718**		52177.780**		0.149		0.147	
	(2.13)		(2.19)		(1.41)		(1.36)	
Covered ban	84725.804*		85848.307*		0.113		0.112	
	(1.69)		(1.73)		(1.10)		(1.09)	
Ban		64150.805**		62412.852**		0.139		0.137
		(2.00)		(2.04)		(1.39)		(1.35)
Bid-ask spread			-174859.43**	-170977.73**			-0.297	-0.301
			(-2.02)	(-2.07)			(-1.55)	(-1.54)
Return			19510.892	19219.242			-0.024	-0.024
			(0.71)	(0.70)			(-1.27)	(-1.26)
SD			-42340.788	-42467.241			0.175***	0.175***
			(-0.43)	(-0.43)			(2.70)	(2.70)
Option-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,526	116,526	114,732	114,732	116,526	116,526	114,732	114,732
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel C. Dependent Variable: Open Interest of All Options

Dependent variable	(1) oi_all	(2) oi_all	(3) oi_all	(4) oi_all	(5) log(oi_all)	(6) log(oi_all)	(7) log(oi_all)	(8) log(oi_all)
Naked ban	104389.786**		99181.757**		0.120		0.115	
	(2.36)		(2.38)		(1.28)		(1.23)	
Covered ban	102999.170		104610.468		0.067		0.065	
	(1.47)		(1.50)		(0.68)		(0.65)	
Ban		103977.787**		100831.961**		0.104		0.100
		(2.08)		(2.11)		(1.16)		(1.10)
Bid-ask spread			-309193.94***	-308568.09***			-0.277**	-0.283**
			(-2.84)	(-2.94)			(-2.04)	(-2.03)
Return			21243.872	21196.849			-0.046**	-0.045**
			(0.56)	(0.56)			(-2.24)	(-2.24)
SD			-62718.100	-62738.488			0.190**	0.190**
			(-0.46)	(-0.46)			(2.48)	(2.48)
Option-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	116,526	116,526	114,732	114,732	116,526	116,526	114,732	114,732
Included option classes	All	All	All	All	All	All	All	All
Number of option classes	171	171	171	171	171	171	171	171

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 7. The Effect of Short Sale Bans on the Composition of Option Trading Volume

Regressions of the proportion of put contracts of overall option trading volume on different types of ban dummies and a number of control variables, using an international panel of daily option and stock data from 10 countries that adopted bans on short sales of stocks at some point in time between January 1, 2008 and February 28, 2011. All regressions include stock-level fixed effects and time fixed effects (weekly). The dependent variable is a ratio of the daily number of traded put option contracts over the total daily number of traded option contracts (puts+calls). Each option contract covers 100 shares. Daily number of traded put (call) contracts is measured as the sum of all put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise. *Return* is the daily return of the underlying stock, expressed in decimals. *Bid-ask spread* is the percentage bid-ask spread of the underlying stock at closing, expressed in decimals. *SD* is a moving standard deviation of the return of the underlying stock based on the previous 10 observations, expressed in decimals. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics. The reason that the number of observations in these regressions is lower than in table 5 and 6 is that for some option classes, the number of traded contracts equals zero in some days, resulting in the denominator being zero for those days.

Dependent Variable: Number of Traded Put Contracts over Total Number of Traded Option Contracts

	(1)	(2)	(3)	(4)
Dependent variable	vm_put-to-option	vm_put-to-option	vm_put-to-option	vm_put-to-option
Naked ban	-0.039***		-0.038***	
	(-4.16)		(-4.03)	
Covered ban	-0.058***		-0.059***	
	(-4.93)		(-5.09)	
Ban		-0.045***		-0.045***
		(-5.08)		(-5.05)
Bid-ask spread			0.163***	0.159***
			(9.38)	(8.84)
Return			-0.173*	-0.173*
			(-1.92)	(-1.92)
SD			0.053**	0.053**
			(2.17)	(2.16)
Stock-level fixed effcts	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly
Number of observations	106,075	106,075	104,568	104,568
Included option classes	All	All	All	All
Number of option classes	171	171	171	171

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 8. Subsample Analysis 1: Non-U.S. and Ten Largest Excluded

Regressions of various measures of option trading activity on different types of ban dummies for two subsamples of countries and option classes, using an international panel of daily option and stock data from countries that adopted bans on short sales of stocks at some point in time between January 1, 2008 and February 28, 2011. The first subsample, denoted "non-US", excludes all option classes of U.S. stocks. The second subsample, denoted "ten largest excluded", excludes the ten largest option classes in terms of average daily number of traded contracts. All regressions include stock-level fixed effects and time fixed effects (weekly). In panel A the dependent variable is the daily number of traded put (call) option contracts; in panel B it is the daily open interest of put (call) option contracts; and in panel C it is a ratio of the daily number of traded put option contracts over the total daily number of traded option contracts (puts+calls). Each option contract covers 100 shares. Daily number of traded put (call) contracts is measured as the sum of all put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. Open interest is measured as the sum of all of put (call) contracts outstanding on a specific day. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* are equal to one and zero otherwise. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics.

Panel A. Dependent Variable: Number of Traded Put and Call Contracts

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable	vm_put	vm_put	vm_call	vm_call	vm_put	vm_put	vm_call	vm_call
Naked ban	1529.741**		1114.464**		153.461		277.509	
	(2.48)		(2.27)		(0.97)		(1.22)	
Covered ban	-170.841		115.146		-182.291		786.157***	
	(-0.36)		(0.26)		(-0.86)		(2.66)	
Ban		1030.593**		821.148**		48.433		436.900**
		(2.14)		(2.41)		(0.37)		(2.43)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	56,102	56,102	56,102	56,102	109,015	109,015	109,015	109,015
Included option classes	non-US	non-US	non-US	non-US	ten largest excluded	ten largest excluded	ten largest excluded	ten largest excluded
Number of option classes	80	80	80	80	161	161	161	161

t-statistics in parentheses

* p<0.10, ** p<0.05, *** p<0.01

Panel B. Dependent Variable: Open Interest of Put and Call Options

Dependent variable	(1) oi_put	(2) oi_put	(3) oi_call	(4) oi_call	(5) oi_put	(6) oi_put	(7) oi_call	(8) oi_call
Naked ban	55420.734**		34145.397		5814.198		9806.787	
	(2.39)		(1.66)		(0.54)		(0.83)	
Covered ban	1263.448		22453.711		-175.468		28779.742**	
	(0.08)		(0.49)		(-0.02)		(1.98)	
Ban		39523.383**		30713.415		3940.419		15752.565
		(2.11)		(1.20)		(0.39)		(1.37)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	56,100	56,100	56,100	56,100	109,013	109,013	109,013	109,013
Included option classes	non-US	non-US	non-US	non-US	ten largest excluded	ten largest excluded	ten largest excluded	ten largest excluded
Number of option classes	80	80	80	80	161	161	161	161

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel C. Number of Traded Put Contracts over Total Number of Traded Option Contracts

Dependent variable	(1) vm_put-to-opt	(2) vm_put-to-opt	(3) vm_put-to-opt	(4) vm_put-to-opt
Naked ban	-0.005		-0.045***	
	(-0.44)		(-4.39)	
Covered ban	-0.017		-0.060***	
	(-1.23)		(-4.94)	
Ban		-0.009		-0.050***
		(-0.82)		(-5.33)
Stock-level fixed effcts	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly
Number of observations	46,886	46,886	98,562	98,562
Included option classes	non-US	non-US	ten largest excluded	ten largest excluded
Number of option classes	80	80	161	161

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 9. Subsample Analysis 2: Shorter Sample Period

Regressions of various measures of option trading activity on different types of ban dummies, using an international panel of daily option and stock data from 10 countries that adopted bans on short sales of stocks at some point in time between December, 2008 and February 28, 2011. All regressions include stock-level fixed effects and time fixed effects (weekly). In Panel A the dependent variable is the daily number of traded put (call) contracts; in panel B it is the daily open interest of put (call) contracts; and in panel C it is a ratio of the daily number of traded put option contracts over the total daily number of traded option contracts (puts+calls). Each option contract covers 100 shares. Daily number of traded put (call) contracts is measured as the sum of all put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. Open interest is measured as the sum of all of put (call) contracts outstanding on a specific day. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics.

Panel A. Dependent Variable: Number of Traded Option Contracts

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable	vm_put	vm_put	vm_call	vm_call	vm_all	vm_all
Naked ban	1556.808*		619.538		2176.346	
	(1.76)		(0.86)		(1.43)	
Covered ban	20.809		2198.986*		2219.795	
	(0.04)		(1.70)		(1.27)	
Ban		816.686*		1380.595*		2197.282**
		(1.70)		(1.93)		(2.19)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	90,125	90,125	90,125	90,125	90,125	90,125
Included option classes	All	All	All	All	All	All
Number of option classes	170	170	170	170	170	170

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel B. Dependent Variable: Open Interest

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable	oi_put	oi_put	oi_call	oi_call	oi_all	oi_all
Naked ban	21276.036		10630.511		31906.547	
	(1.09)		(0.40)		(0.74)	
Covered ban	14694.238		104162.255		118856.493	
	(0.47)		(1.46)		(1.18)	
Ban		18104.345		55702.350		73806.695
		(0.98)		(1.27)		(1.23)
Stock-level fixed effcts	Yes	Yes	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly	Weekly	Weekly
Number of observations	90,124	90,124	90,124	90,124	90,124	90,124
Included option classes	All	All	All	All	All	All
Number of option classes	170	170	170	170	170	170

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel C. Dependent Variable: Number of Traded Put Contracts over Total Number of Traded Option Contracts

	(1)	(2)
Dependent variable	vm_put-to-option	vm_put-to-option
Naked ban	-0.012	
	(-1.04)	
Covered ban	-0.067***	
	(-4.68)	
Ban		-0.039***
		(-3.51)
Stock-level fixed effcts	Yes	Yes
Time-fixed effects	Weekly	Weekly
Number of observations	81,764	81,764
Included option classes	All	All
Number of option classes	170	170

t-statistics in parentheses

* p<0.10, ** p<0.05, *** p<0.01

Table 10. Subsample Analysis 3: Naked and Covered Bans

Regressions of various measures of option trading activity on different types of ban dummies for two subsamples of countries and option classes, using an international panel of daily option and stock data from countries that adopted bans on short sales of stocks at some point in time between January 1, 2008 and February 28, 2011. The first subsample, denoted "naked regimes", only includes option classes of stocks listed in countries that only adopted bans prohibiting naked short sales. The second subsample, denoted "covered regimes", only includes option classes of stocks listed in countries that only adopted bans prohibiting all short sales (both naked and covered short sales). Option classes of stocks that were subject to both types of bans at different point in time are excluded from both subsamples. All regressions include stock-level fixed effects and time fixed effects (weekly). In panel A the dependent variable is the daily number of traded put (call) option contracts; in panel B it is the daily open interest of put (call) option contracts; and in panel C it is a ratio of the daily number of traded put option contracts over the total daily number of traded option contracts (puts+calls). Each option contract covers 100 shares. Daily number of traded put (call) contracts is measured as the sum of all put (call) contracts of all American put (call) options on the underlying stock listed on the domestic derivatives marketplace that are traded during the day, irrespective of the expiry dates and strike prices of the individual options. Hence, we only have one observation per underlying stock (firm) per day. Open interest is measured as the sum of all put (call) contracts outstanding on a specific day. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if all short sales (both naked and covered short sales) are forbidden and zero otherwise. The standard errors are robust for heteroskedasticity and clustered at the stock level. The numbers in parenthesis below the coefficient estimates are t-statistics.

Panel A. Dependent Variable: Number of Traded Put and Call Contracts

	(1)	(2)	(3)	(4)
Dependent variable	vm_put	vm_call	vm_put	vm_call
Naked ban	3806.489*	4355.219		
	(1.81)	(1.60)		
Covered ban			-274.319**	-146.023
			(-2.27)	(-0.67)
Stock-level fixed effcts	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly
Number of observations	22,535	22,535	13,194	13,194
Included options classes	naked regimes	naked regimes	covered regimes	covered regimes
Number of option classes	29	29	19	19

t-statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel B. Dependent Variable: Open Interest of Put and Call Options

	(1)	(2)	(3)	(4)
Dependent variable	oi_put	oi_call	oi_put	oi_call
Naked ban	43703.501	75425.198		
	(1.24)	(1.65)		
Covered ban			-8539.863	1901.545
			(-0.88)	(0.15)
Stock-level fixed effcts	Yes	Yes	Yes	Yes
Time-fixed effects	Weekly	Weekly	Weekly	Weekly
Number of observations	22,533	22,533	13,194	13,194
Included options classes	naked regimes	naked regimes	covered regimes	covered regimes
Number of option classes	29	29	19	19

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Panel C. Number of Traded Put Contracts over Total Number of Traded Option Contracts

	(1)	(2)
Dependent variable	vm_put-to-option	vm_put-to-option
Naked ban	-0.024	
	(-1.35)	
Covered ban		-0.065***
		(-2.99)
Stock-level fixed effcts	Yes	Yes
Time-fixed effects	Weekly	Weekly
Number of observations	16,448	12,438
Included options classes	naked regimes	covered regimes
Number of option classes	29	19

t statistics in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Figure 1. Expected Effects of Short Sale Bans on Option Trading Activity

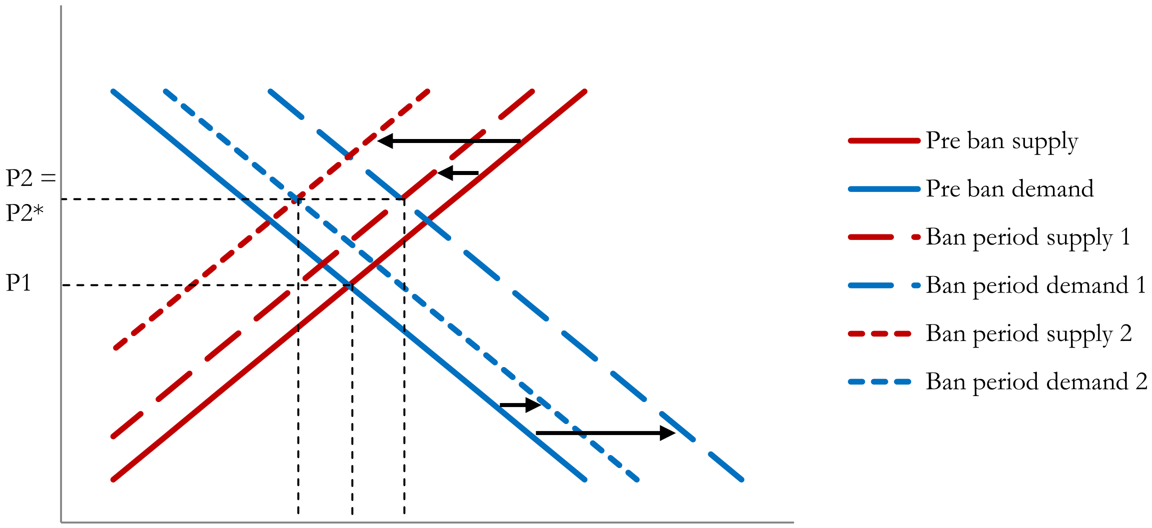


Figure 5 illustrates the expected effects of a short sale ban on option trading activity in options on stocks that are subject to the ban. For simplicity, let's use option trading volume as a measure for option trading activity. The quantity of options is on the x-axis and the price of option liquidity is on the y-axis. If short sale bans induce investors to switch to the option markets, we would expect to see an increase in the demand for options (for long put positions and short call positions). This increase in demand is illustrated by a shift of the blue solid demand curve denoted "Pre ban demand" to the right. The ban is also likely to increase the adverse selection costs and inventory holding costs of option market makers, which reduces their willingness to provide liquidity. This is illustrated by a shift of the red solid supply curve denoted "Pre ban supply" to the left. As we can see, the net effect of the ban on option trading volume is ambiguous. If the demand increases more than the corresponding decrease of the supply, we would expect option trading volume to increase from Q1 to Q2. On the other hand, if the supply decreases more than the corresponding increase in the demand, we would expect option trading volume to decrease from Q1 to Q2*. In both scenarios, the equilibrium price paid by the party demanding liquidity to the party supplying liquidity (mainly option market makers) increases (shown by the increase of the price from P1 to P2=P2*). On the option markets, this increase would be reflected by an increase in the bid-ask spreads of options on banned stocks (increased bid-ask spreads of options on banned stocks are documented by Battalio and Schultz (2009) and Grundy et al. (2010)).

Figure 2: Ban Periods around the World, January 2008 – February 2011

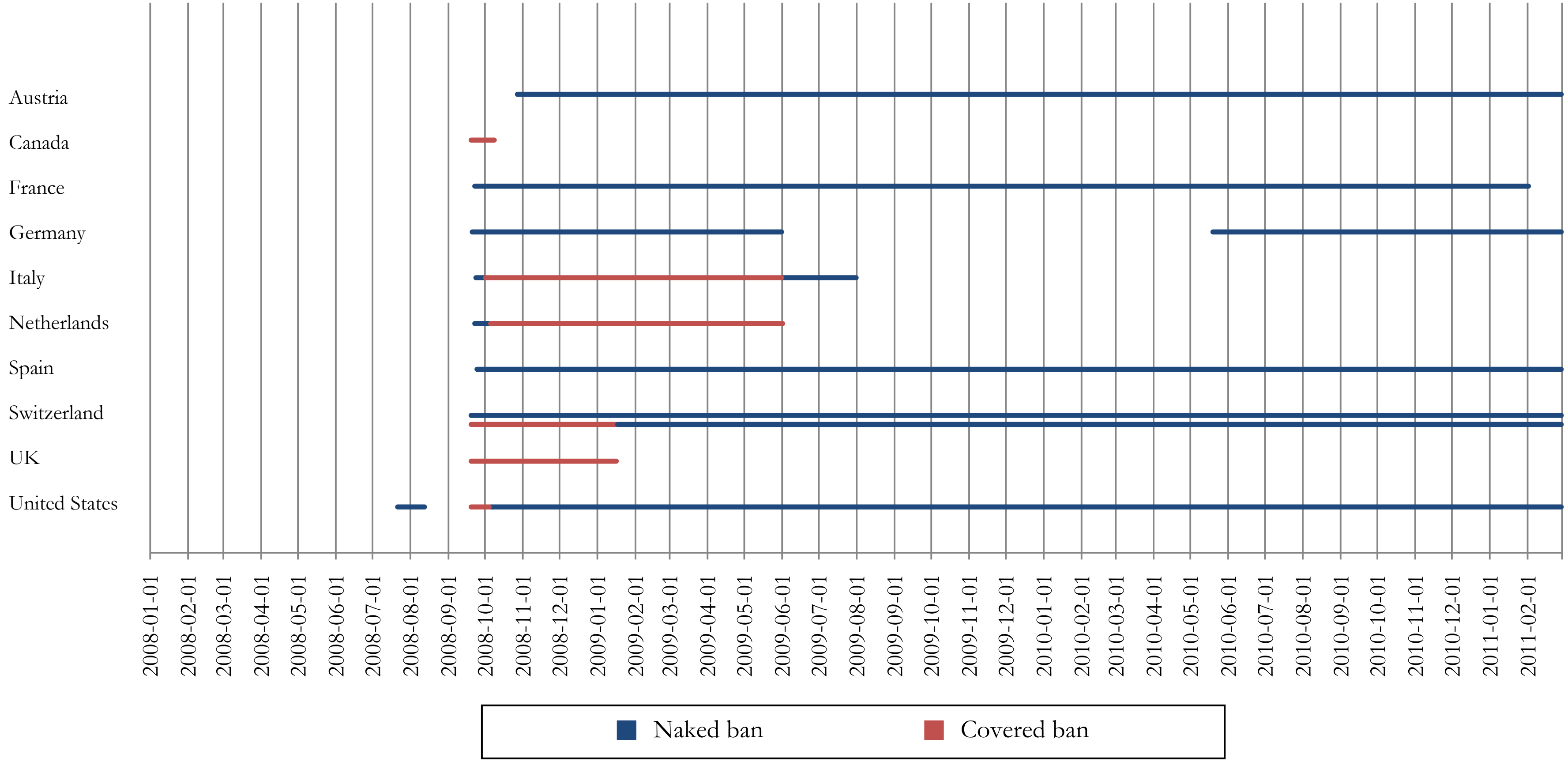


Figure 2 illustrates the introduction and removal of short sale bans of varying stringency around the world. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise.

Figure 3: Fraction of Stocks Subject to a Ban

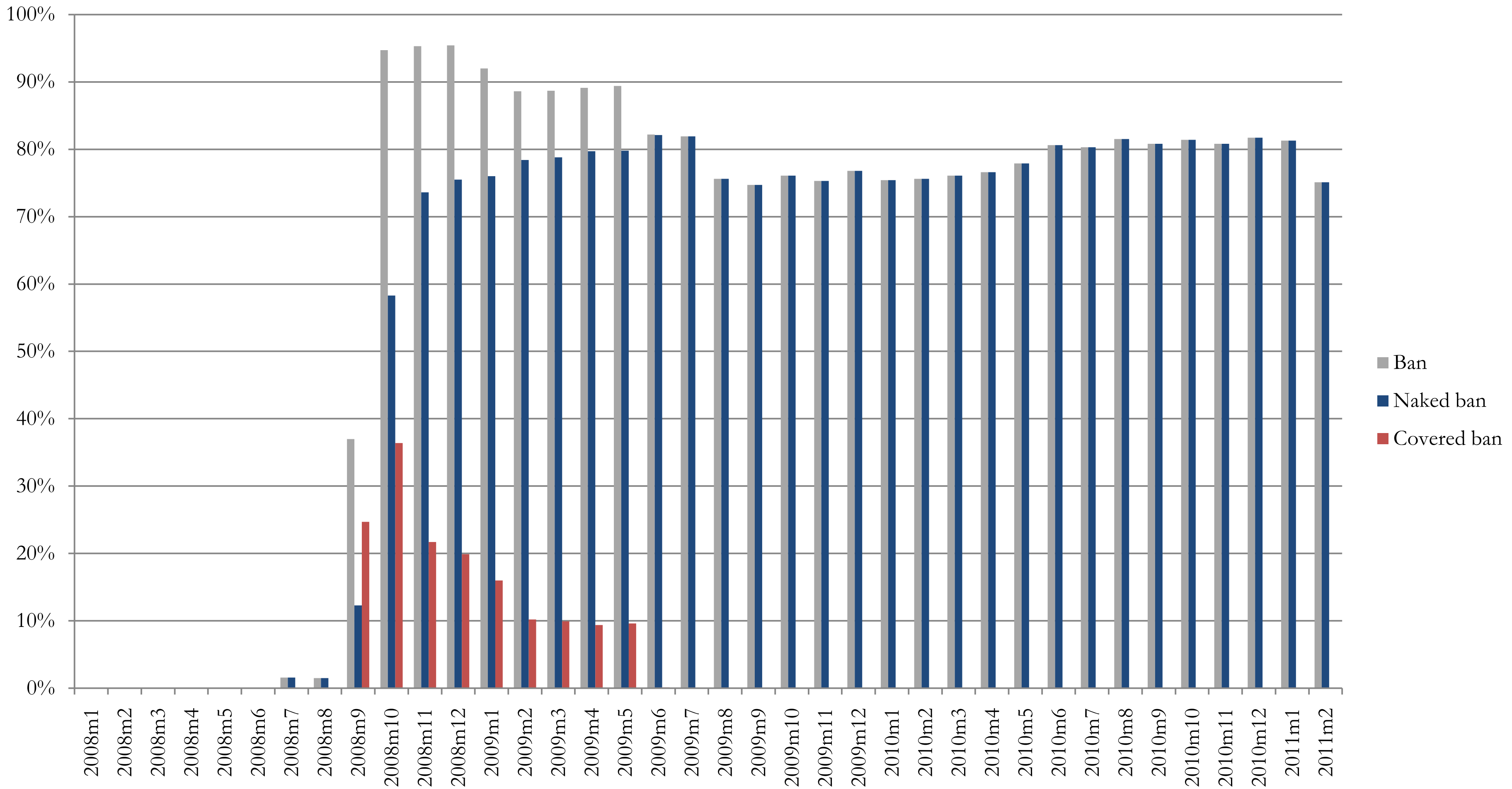


Figure 3 illustrates the fraction of stocks in the sample that is subject to a ban over time. *Naked ban* is a dummy variable that equals one if naked short sales are forbidden and covered short sales are allowed and zero otherwise. *Covered ban* is a dummy variable that equals one if even covered short sales are forbidden and zero otherwise. *Ban* is a dummy variable that equals one if short sales (either naked short sales *or* all short sales – both naked and covered) are forbidden. Hence, *Ban* equals one when one of the two dummy variables *Naked ban* and *Covered ban* is equal to one and zero otherwise.

Figure 4. Cumulative Distribution of Average Daily Trading Volume

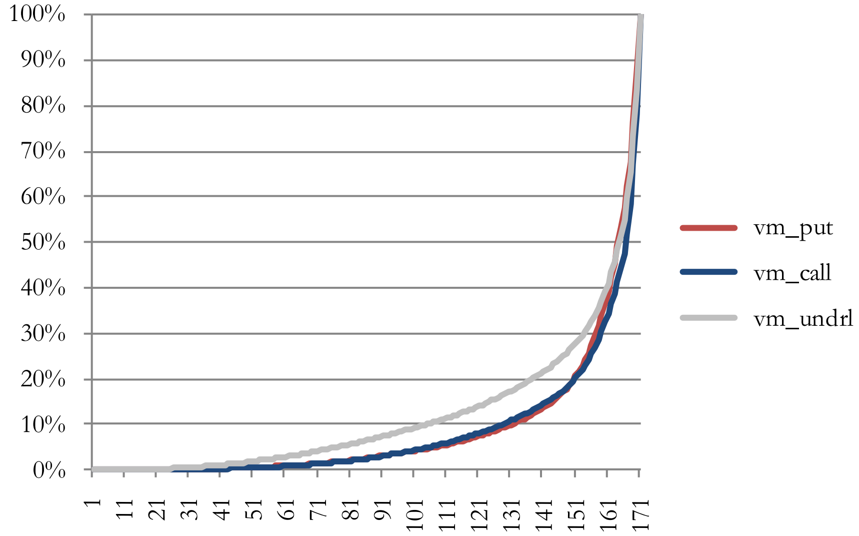
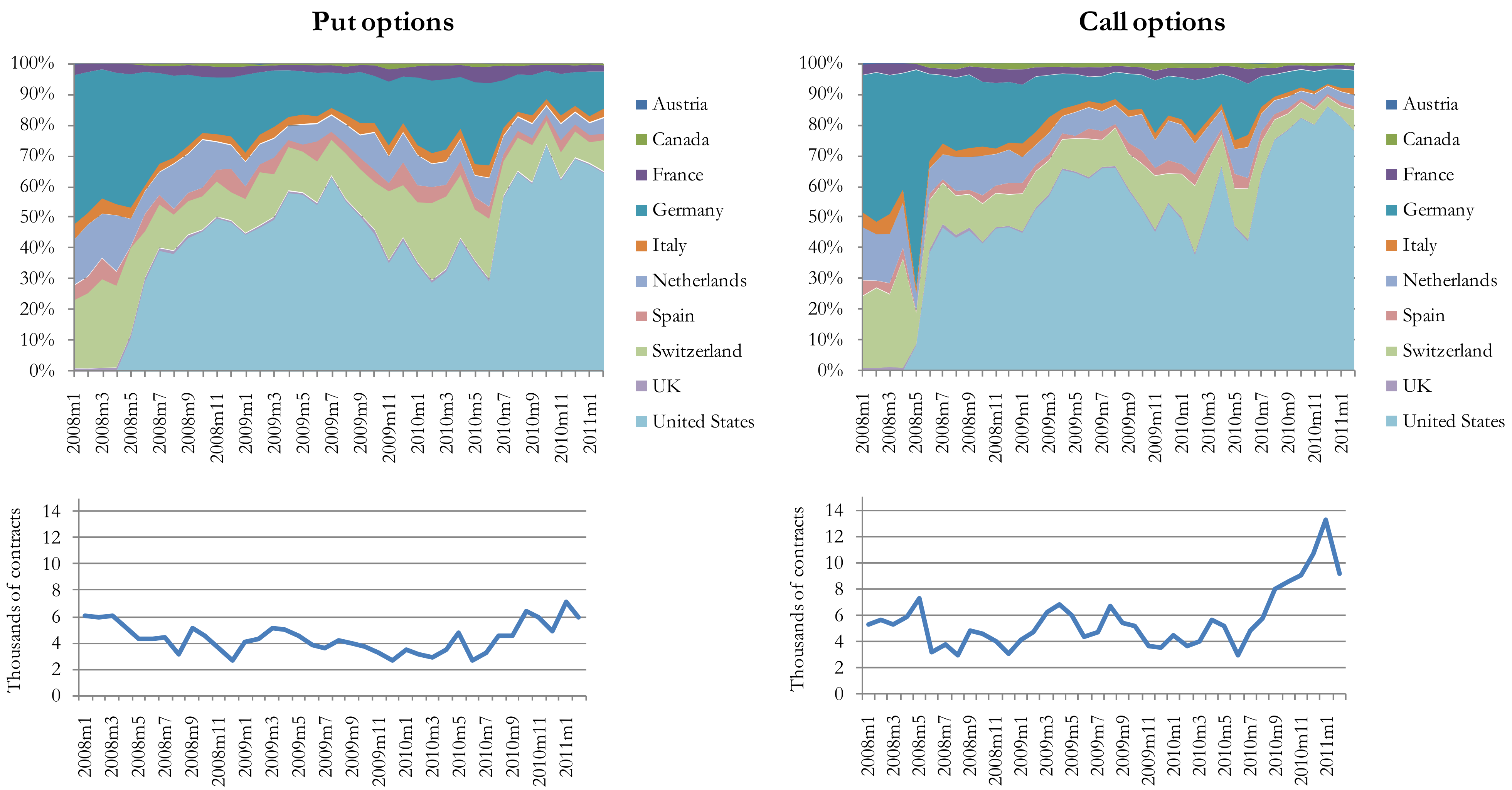
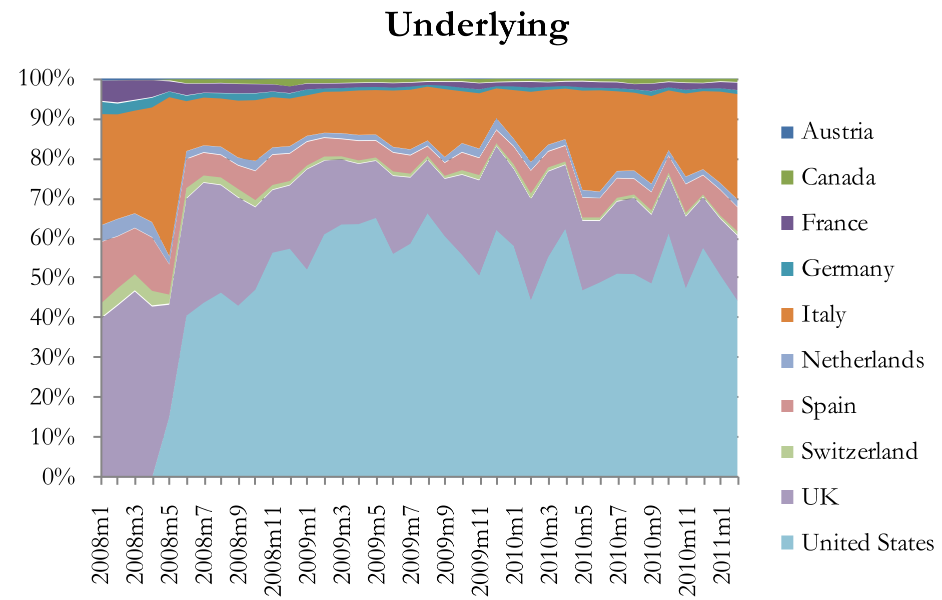


Figure 4 illustrates the cumulative distribution of average daily trading volume across option classes/stocks. Each variable is sorted independently from smallest to largest. *Vm_put* is the trading volume of put options, measured as the daily number of traded put contracts (each contract covers 100 shares); *vm_call* is the trading volume of call options, measured as the daily number of traded call contracts; and v*m_undrl* is the trading volume of the underlying stock, measured in thousands of shares.

Figure 5. Distribution of Trading Volume across Countries





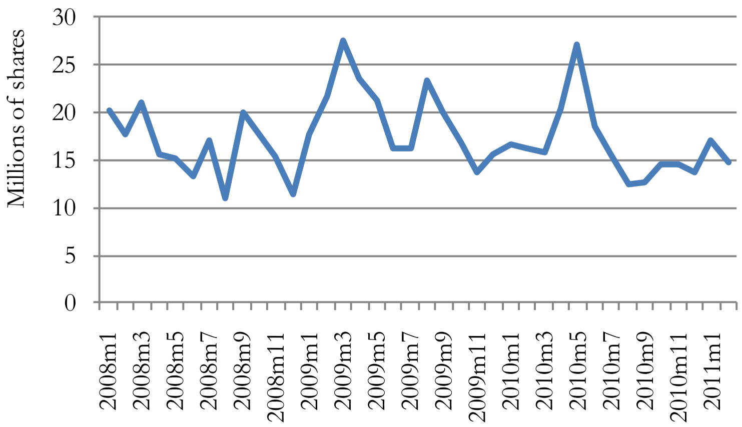


Figure 5 illustrates the distribution of the monthly cumulative number of traded put and call contracts, as well as the monthly cumulative trading volume in the underlying stock across countries over time. In addition, average daily trading volume for an average option class (stock) is illustrated below. Note that each option contract covers 100 shares.

NB: Stocks and pairs of option classes of some countries lack observations for some time periods, the most obvious example being that we have no data for U.S. options and U.S. stocks until May 2008.

Figure 6. Trading Volume of Put Options around the Introduction of the Bans

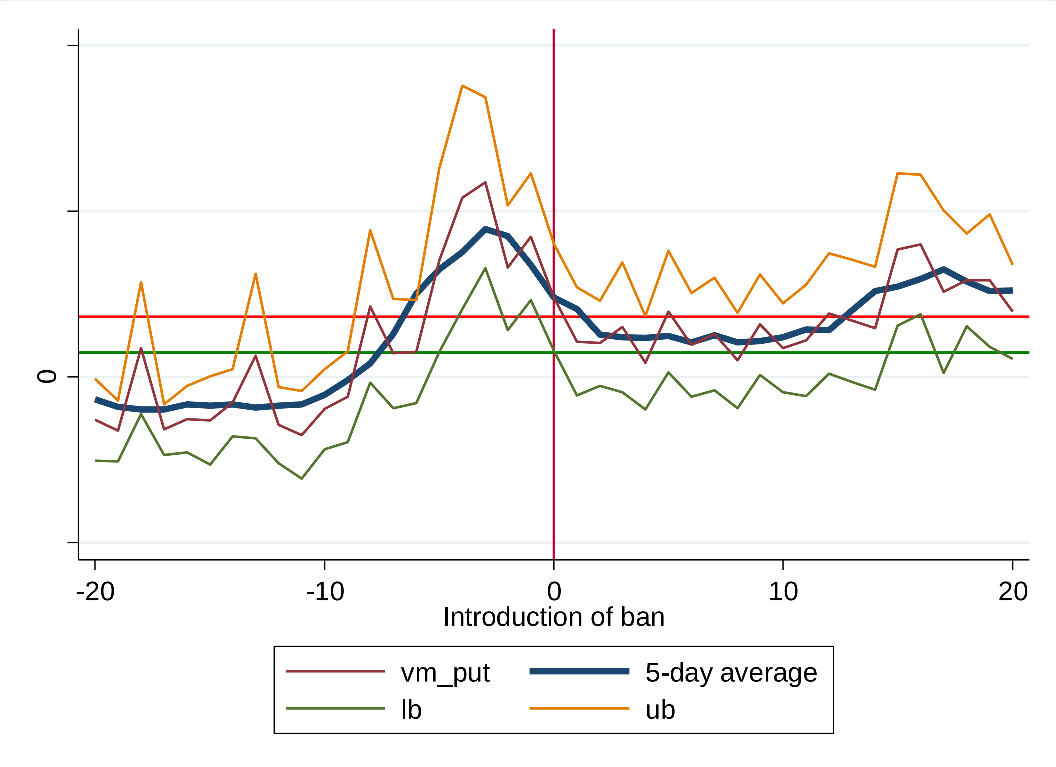


Figure 6 compare the average level of trading volume of puts for each day during a 41 day window ranging from 20 days before the introduction date up to 20 days after the introduction date with the average level of the measure for all non-ban observations except those belonging to the 20 days preceding the introduction of the bans. See the methodology section for the specification used. We make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), and thus we use the *Ban* dummy in order to decide on the introduction date. The second ban period is used for the U.S. (the ban period starting on September 18, 2008) and the first ban period is used for Germany (the ban period starting on September 20, 2008). Observations belonging to the first ban period in the U.S. and the second ban period in Germany are dropped as they shall not be part of the control group. "lb" and "ub" are the lower and upper bound of a 90 % confidence interval for the estimated coefficients. The green horizontal line is the average of the coefficients for the 20 days preceding the introduction of the bans, and the red horizontal line is the average of the coefficients for the 20 days following the introduction date.

Figure 7. Open Interest of Put Options around the Introduction of the Bans

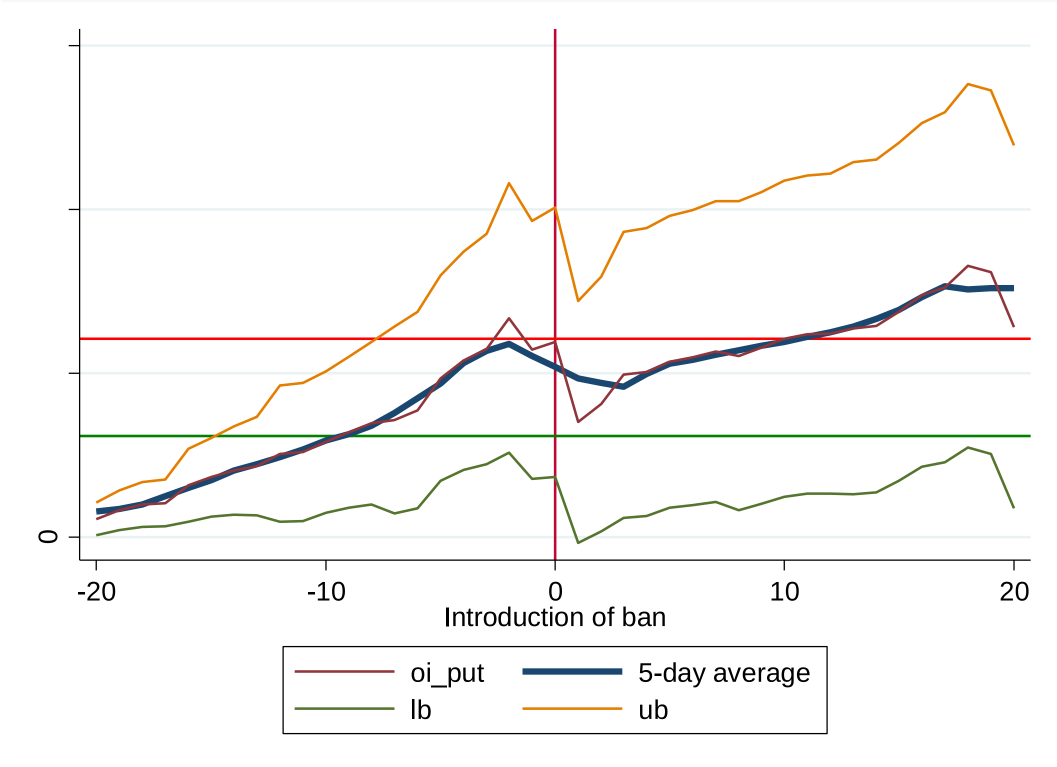


Figure 7 compare the average level of open interest of puts for each day during a 41 day window ranging from 20 days before the introduction date up to 20 days after the introduction date with the average level of the measure for all non-ban observations except those belonging to the 20 days preceding the introduction of the bans. See the methodology section for the specification used. We make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), and thus we use the *Ban* dummy in order to decide on the introduction date. The second ban period is used for the U.S. (the ban period starting on September 18, 2008) and the first ban period is used for Germany (the ban period starting on September 20, 2008). Observations belonging to the first ban period in the U.S. and the second ban period in Germany are dropped as they shall not be part of the control group. "lb" and "ub" are the lower and upper bound of a 90 % confidence interval for the estimated coefficients. The green horizontal line is the average of the coefficients for the 20 days preceding the introduction of the bans, and the red horizontal line is the average of the coefficients for the 20 days following the introduction date.

Figure 8. Trading Volume of Put Options around the Removal of the Bans

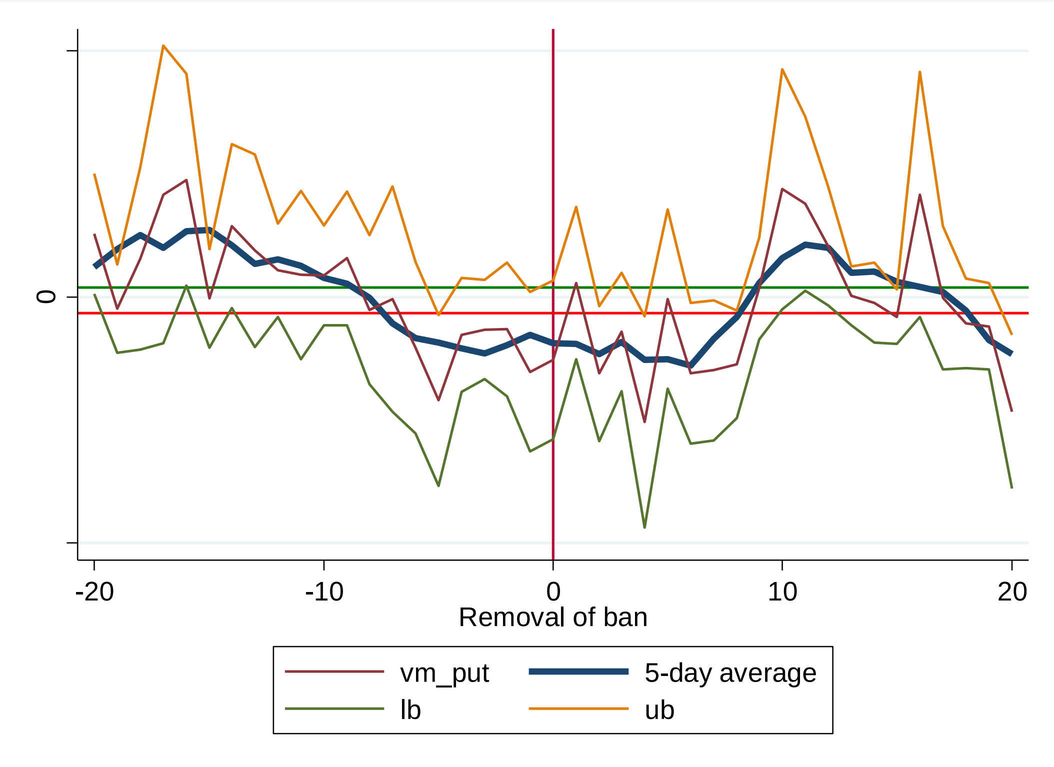


Figure 8 compare the average level of trading volume of puts for each day during a 41 day window ranging from 20 days before the removal date up to 20 days after the removal date with the average level of the measure for all non-ban observations except those belonging to the 20 days following the removal of the bans. See the methodology section for the specification used. We make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), and thus we use the *Ban* dummy in order to decide on the removal date. In some countries (e.g. Austria, Spain, Switzerland and the U.S), the bans are still in effect, and observations belonging to option classes of stocks listed in these countries are thus dropped for this graphical analysis. The first ban period is used for Germany (the ban period starting on September 20, 2008 and ending on May 31, 2009). Observations belonging to the second ban period in Germany are dropped as they shall not be part of the control group. "lb" and "ub" are the lower and upper bound of a 90 % confidence interval for the estimated coefficients. The green horizontal line is the average of the coefficients for the 20 days preceding the removal of the bans, and the red horizontal line is the average of the coefficients for the 20 days following the removal date.

Figure 9. Open Interest of Put Options around the Removal of the Bans

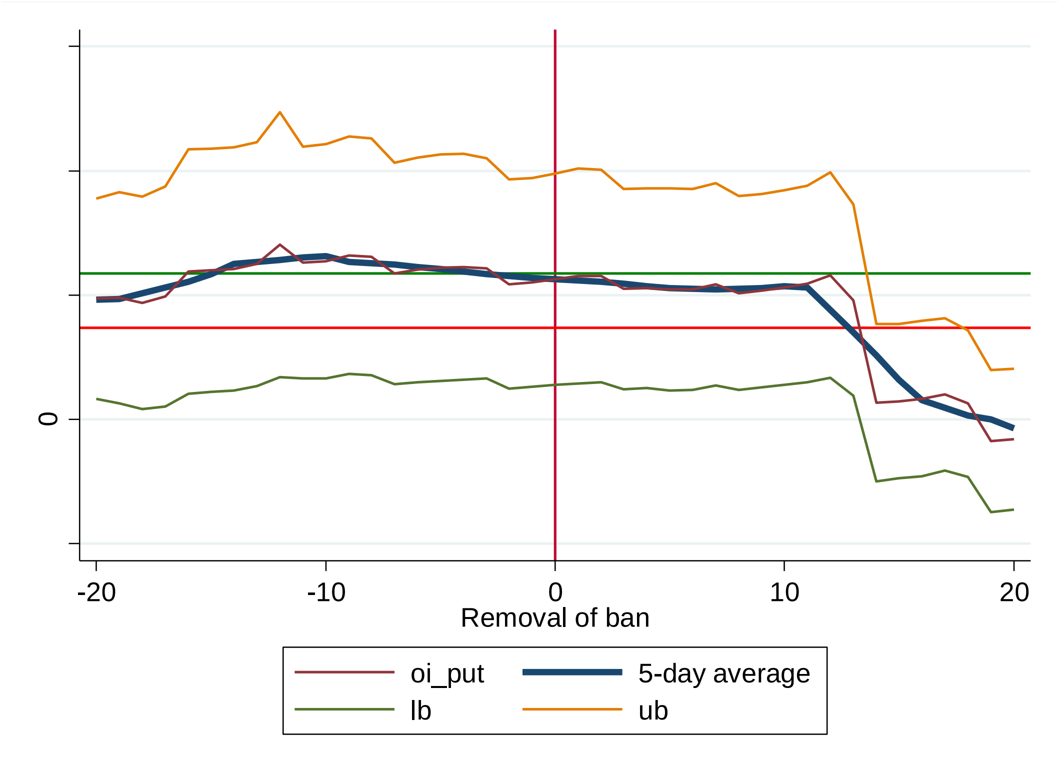


Figure 9 compare the average level of open interest of puts for each day during a 41 day window ranging from 20 days before the removal date up to 20 days after the removal date with the average level of the measure for all non-ban observations except those belonging to the 20 days following the removal of the bans. See the methodology section for the specification used. We make no difference between naked bans (bans on naked short sales) and covered bans (bans on all short sales – both naked and covered), and thus we use the *Ban* dummy in order to decide on the removal date. In some countries (e.g. Austria, Spain, Switzerland and the U.S), the bans are still in effect, and observations belonging to option classes of stocks listed in these countries are thus dropped for this graphical analysis. The first ban period is used for Germany (the ban period starting on September 20, 2008 and ending on May 31, 2009). Observations belonging to the second ban period in Germany are dropped as they shall not be part of the control group. "lb" and "ub" are the lower and upper bound of a 90 % confidence interval for the estimated coefficients. The green horizontal line is the average of the coefficients for the 20 days preceding the removal of the bans, and the red horizontal line is the average of the coefficients for the 20 days following the removal date.