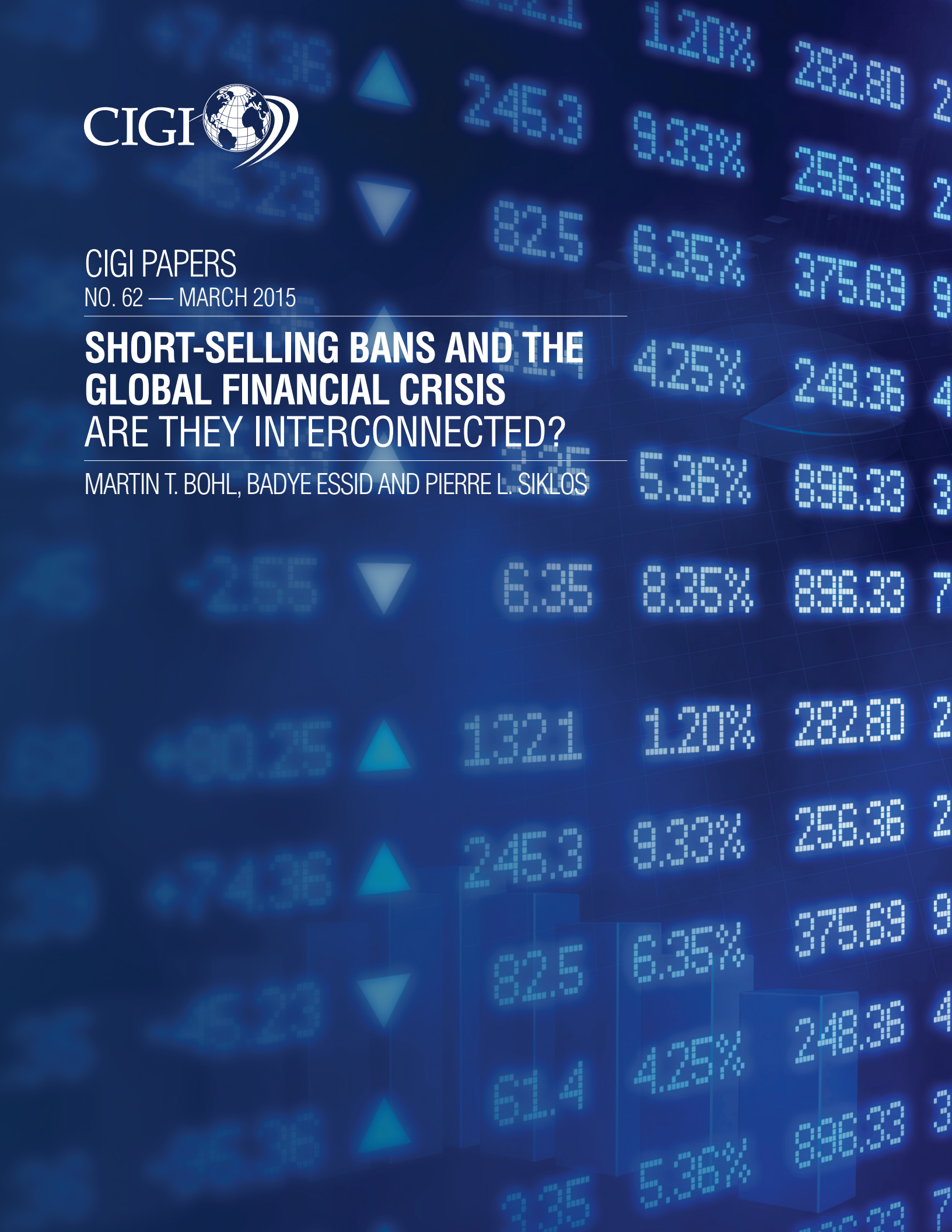

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SHORT-SELLING BANS AND THE GLOBAL FINANCIAL CRISIS
ARE THEY INTERCONNECTED?

MARTIN T. BOHL, BADYE ESSID AND PIERRE L. SIKLOS

SHORT-SELLING BANS AND THE GLOBAL FINANCIAL CRISIS: ARE THEY INTERCONNECTED?

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EXECUTIVE SUMMARY

This paper observes that short-selling bans spread globally beginning in 2007. We seek to empirically determine whether there were spillover effects over and above the domestic impact from the imposition of such bans. There is some evidence that the bans were unsuccessful, at least insofar as they did not take into account the global component a short-selling ban might have. In the individual countries we examine, the bans had relatively little impact. Nevertheless, our finding that equity returns do not appear to show a decline may be evidence that the bans stemmed further deterioration in stock prices that policy makers sought to avoid.

INTRODUCTION

Although the past few years have, for the most part, seen stock markets surge around the world, fears of an imminent correction in stock prices are always on the minds of investors, especially as central banks contemplate removing extraordinarily loose monetary policies, albeit gradually. Events such as the 2013 "taper tantrum," and signals that central bank quantitative easing will eventually end, have contributed to the impression that the recent upward trend stock market will be reversed. The Bank for International Settlements (BIS), among others, points out that the compression of yields in bonds and similar financial instruments has spilled over into stocks in advanced economies especially, and there is a likelihood that the recent stock price rises will be undone once monetary conditions tighten (BIS 2014, chapter 2). Indeed, emerging market economies (EMEs) have experienced a sell-off of stocks despite no evidence to date that a lasting downward movement in stock indices is underway. More importantly, recent events have contributed to raising not only overall uncertainty, but also regulatory uncertainty — in the event of a stock market downturn, loose monetary policies are viewed by financial markets as being withdrawn too early — since short-sale bans tend to be imposed without warning (Battalio and Schultz 2011).

The possibility of a significant downturn in stock markets implies that short-sellers may well lead the way. As a result, the perennial question concerning whether short-selling exacerbates a downturn in stock prices is always on the minds of policy makers and academics. History is replete with episodes of regulators banning short-selling. However, a distinguishing characteristic of the global financial crisis (GFC) that erupted in 2007 is that short-selling was banned almost simultaneously in many parts of the world. The combination of greater global financial market integration and loose monetary policies has led to a rise in the systemic component of risk (see International Monetary Fund 2014). Hence, policy makers who might otherwise not follow the lead of some and impose a short-

selling ban may well do so, even if observable economic conditions might not warrant such a step.

It is well-known that financial markets may appear coupled because comparable fundamentals drive asset price movements. Financial globalization has made it easier for investors to buy and sell stocks around the world. Technological changes have also reduced the transactions costs to trading in stocks. Finally, as news travels quickly, events in one part of the world may easily influence stock prices elsewhere, thereby providing another avenue for stock returns to be globally correlated. If a substantial amount of global decoupling in stock price movements is expected, perhaps in part because of home bias in portfolio investing, then contagion is another possibility that could drive stock returns in different parts of the globe to be significantly correlated with each other. Contagion is the phenomenon whereby shocks are transmitted for reasons that cannot be explained by fundamentals such as trade and common business cycle movements.

As a result, this paper begins with the observation, which we document, that short-selling bans spread globally beginning in 2007. We find some evidence that restrictions on short-selling were unsuccessful, at least insofar as the extant literature has not adequately taken into account the global component associated with the bans. In the individual countries we investigate, the bans had relatively little impact. Nevertheless, the fact that stock returns do not appear to show a decline may be seen as evidence that the bans stemmed further deterioration in stock prices that policy makers sought to avoid. We are also able to identify sharp changes in the dynamic conditional correlations across stock markets. Indeed, we report sharp increases in these correlations across markets that were previously uncorrelated. We attribute part of the increase to the global spread of short-selling bans during the financial crisis of 2008-2009. Since policy makers may well have reacted to what they perceived to be a common shock, the imposition of the bans also had a comparable impact across the globe. As a result, in future, it is likely that policy makers will have to consider more closely both the timing and the nature of the response to short-selling bans in view of the existence of the spillovers of the kind we have identified. Clearly, these findings add another element to the increasing prominence of financial system stability as a separate objective of policy.

As will be shown in the paper, short-selling bans spread quickly around the globe in the wake of the GFC (see Beber and Pagano 2013). This development was not unique to the events of 2007–2009 (see, for example, Bris, Goetzmann and Zhu 2007). Superficially, then, these phenomena suggest the possibility of interconnectedness in the response of policy makers to developments, especially negative ones, in their own stock market. Investigating the correlation in stock returns and the role played by the imposition of short-selling bans requires that the resulting estimates should be conditioned on other factors that could also explain co-movements in returns. Moreover, since stock returns are volatile and there is the potential for a large number of stock markets to investigate, researchers must be aware of the potential for the dimensionality of the problem to become large. For all these reasons, we empirically investigate the links arising from the imposition of short-selling bans using an econometric technique well-suited to handling the difficulties just described. Accordingly, we estimate a model of stock returns relying on the dynamic conditional correlations (DCC) approach combined with a Generalized AutoRegressive Conditional Heteroskedasticity (GARCH) model. The latter is typically the preferred methodology under the circumstances of estimating conditional volatilities.[1]

The paper provides a brief literature review in the following section, focusing on the nature and type of short-selling bans put in place in recent memory. Next, it describes the econometric methodology and the data employed, as well as a few stylized facts about the performance of stock markets globally. The empirical evidence is discussed before concluding.

LITERATURE REVIEW

Financial crises, especially the most recent ones, have prompted policy makers to impose short-selling bans. The often-stated fear is that large-scale shorting will drive down stock prices contributing to a massive loss of confidence in financial markets. One feature of the GFC that originated in the United States in 2007 is that short-selling restrictions proved "contagious." That is, several countries imposed restrictions of various durations and severity (Reuters 2009; Mackintosh, Mitchell and Fry 2009). S. N. Gruenewald, A. F. Wagner and R. H. Weber (2010a; 2010b) provide a descriptive overview of the legal aspects of the most recent bout of short-sale restrictions imposed around the world.[2]

There exists a rich and diverse literature assessing the impact of short-selling restrictions. Space constraints prevent a complete listing of the vast literature that explores various facets of the impact of imposing short-selling constraints. A. Bris, W. N. Goetzmann and N. Zhu (2007) provide many of the most important references on the topic. A. Beber and M. Pagano (2013) and A. Jain et al.

1 Note that other methodologies were experimented with (see footnote 13), but the relevant results are not reported as the conclusions are unaffected.

2 Recent comprehensive economic and statistical analyses of the impact of short-sale constraints are found in Bris, Goetzmann and Zhu (2007), and Charoenrook and Daouk (2009).

(2013) are recent studies that, like ours, take a global view of the short-sale restrictions during the 2007–2009 period.[3]

The onset of a crisis appears to whet the appetite of regulators in favour of banning short-selling opportunities. Their logic is that a downward movement in stock prices will be exacerbated by short sellers. Yet, as pointed out by J. E. Engelberg, A. V. Reed and M. C. Ruggenberg (2012), the evidence that short-selling bans of all types create a variety of distortions in stock markets is "overwhelming." Similarly, other observers (see, for example, Blinder 2013, 282) have suggested that "short-selling probably kept the housing and bond bubbles from blowing up even bigger than they did." In other words, the prospect that asset price increases, including stock prices, would be reversed at some point moderates the emergence of bubbles as there are investors willing to bet against future asset price increases.

Nevertheless, a theoretical case can be made that short-selling restrictions can increase the likelihood of stock market crashes, as these tend to follow stock market booms or bubbles (Abreu and Brunnermeier 2003; Scheinkmann and Xiong 2003). More recently, M. Brunnermeier and E. H. Omke (2013) made the interesting observation that, because financial institutions are special in that they face a more binding leverage constraint than other types of firms, short-selling bans in times of crisis can actually be destabilizing. Otherwise, healthy financial institutions are under threat from poor balance sheet positions that afflict weak banks.

Theoretical models also find that short-selling bans increase the prospect of stock market bubbles and lead to excessive stock market volatility. Of course, a short-selling ban is often introduced at a time when other economic conditions might also be expected to prompt policy makers to act to stem the downward movement in stock prices.[4] There is fairly broad agreement that stocks are more volatile in the presence of constraints on short selling, although the empirical evidence is inconclusive. Unsurprisingly, empirical work continues to investigate the issues.

Empirical evidence suggests that banning short selling distorts markets because it hinders the ability of markets to engage in price discovery (Boehmer and Wu 2009). In the absence of short-sale restrictions, stock prices ought to be determined according to underlying fundamentals. Instead, a ban will exclude relatively well-informed market participants, leading to the overpricing of equities (Miller 1977). Banning the shorting of stocks also impacts liquidity, which is reduced as informed investors withdraw from the market (see Boehmer, Jones and Zhang 2008). As a consequence, restrictions on this kind of activity produce less efficient stock pricing. Moreover, models of investor behaviour have implications for higher moments of the distribution of returns, reflected in the volatility and skewness of returns.[5] Yet, the fear that engaging in short selling increases the frequency of large negative returns (i.e., stock market crashes) is not supported by the available empirical evidence (Bris, Goetzmann and Zhu 2007; Saffi and Sigurdsson 2011). Indeed, short-selling bans may result in asymmetric effects in the behaviour of higher moments in the distribution of stock returns (see, for example, Bohl, Essid and Siklos 2012).

At the heart of the debate about imposing short-selling bans is how well informed the traders are that engage in short selling. The consensus is that short sellers are better informed and, therefore, play a valuable role. Indeed, for this reason, short-selling bans are believed to be efficiency-reducing policies since they result in the overpricing of equities (see Miller 1977) or influence the price discovery process, leading to higher bid-ask spreads (Diamond and Verrecchia 1987). A recent spate of papers on the subject confirms this view (see Chague et al. 2014; Bernal, Hendrickx and Szafarz 2014; Kelley and Tetlock 2013; Liu, McGuire and Swanson 2013; Lynch et al. 2014; Engelberg, Reed and Ruggenberg 2012 and references therein).[6] Much of the literature focuses on the US experience (see Boehmer, Jones and Zhang 2013; Bailey and Zheng 2013).

From the perspective of this study, we are also interested in the role of information as it pertains to the impact of short-selling bans. However, our focus is on the global impact of this type of policy, as well as on the influence that bans have

3 In particular, Table 1 and Figure 1 in Beber and Pagano (2013) and Table 1 in Jain et.al. (2013) contain details about the timing and type of short-selling bans around the world. For the cross-country evidence considered in this paper, we have also compiled comparable information that is relegated to an Appendix elsewhere. However, essential information about short-selling bans around the world during the GFC is in Table 1 of this paper.

4 Empirical evidence suggests that the volatility of stocks is higher in recessions (Hamilton and Lin 1996) or when returns are negative (Bekaert and Wu 2000). Changes in the volatility of stock returns have also been associated with increases in political tensions (Bittlingmayer 1998).

5 Bris, Goetzmann and Zhu (2007) report strong evidence that the removal of short-sale restrictions is associated with more negative skewness in returns, based on a large cross-section of countries, including China. In an equally large panel analysis, Charoenrook and Daouk (2009) find no significant impact on skewness from short-selling bans.

6 Dupuis and Kryzanowski (2014) are a recent exception. They claim that short-selling bans create "intangible" costs and propose a new taxonomy to understand this kind of policy. Their empirical investigation, relying on data from 2006 to 2010, rejects Miller's (1977) overpricing of equities hypothesis when a ban is imposed. Nevertheless, they also describe some of their evidence as mixed.

on aggregate equity prices.[7] As noted above, this aspect of the relevant literature has not received as much attention. Indeed, the usual approach has been to consider how individual stocks react to various types of short-selling bans. In this study, we consider the behaviour of aggregate equity indices around the world. Several authors have noted that macroeconomic information, as well as publicly available information, gives short sellers the advantage over other traders in equity markets (see Engelberg et al. 2012). Moreover, Jain et al. (2013) demonstrate that while research has typically considered that short-selling bans for some types of stocks (such as financial) can create arbitrage opportunities within a particular market, globalization in finance has also prompted regulators to consider that they must deal with spillovers into other markets.[8] This raises the issue of the reach of regulators, as well as the possibility that the global spread of short-selling bans during the GFC was no accident. It is with this in mind that we proceed to an empirical investigation of the international consequences arising from the imposition of various short-selling bans around the world since 2007.

DATA

We rely on daily stock price indices from January 2, 1995 to December 3, 2013 (inclusive), covering global stock markets from Europe, Asia and North America. A total of 18 stock markets are considered.[9] The choice of a long sample is to facilitate investigating the impact of short-selling bans before, during and after the policies were put into place. The data set includes FTSE 100 Index price indices for Australia, Japan, South Korea, India, Indonesia, Malaysia, Canada, the United States, Norway, Denmark, Germany, Austria, Switzerland, France, Portugal, Greece and Italy. All data were obtained from Thomson Reuters Datastream.

All of the countries included in our set imposed some sort of a ban on short selling. More often than not the ban was limited to financial stocks. Occasionally, the short-selling ban extended to all stocks. In the cases of Japan, Germany and Portugal there were also bans on naked short-selling. We considered these on the same footing as ordinary short-selling bans. For our purposes — and the previous section's literature review provides some support for this view — we make no distinction between the two types of short-selling ban.[10]

Table 1 provides a list of the markets in our sample and provides some information about the timing of short-selling bans. Beber and Pagano (2013) and Jain et al. (2013) are two other sources for the dates when short-selling bans were imposed. The dates are virtually the same across the sources examined.[11] Countries in the euro zone tend to have banned financial stocks, with most retaining a ban on naked short-selling well after the GFC. Of course, the sovereign debt crisis in the euro zone continues to linger. In a few other countries (such as Canada, the United Kingdom and the United States) the bans were short lived. With only two exceptions (Malaysia and India) the bans were introduced on the heels of the GFC, which erupted in the United States in 2008. Bans are equally distributed in the dataset between ones that were applied to all stocks versus a ban of financial stocks.[12]

7 An issue that is occasionally raised is whether the type of short-sale ban can make a difference. Typically, bans come in two forms. The most prevalent is the covered short-selling ban, followed by naked short-selling bans (i.e., sale of securities without borrowing them for delivery to the buyer). Short-sale bans can also be influenced by the regulators' disclosure requirements. Although the differences might matter, empirical evidence (see Bernal, Hendrickx and Szafarz 2014; Liu, McGuire and Swanson 2013) suggests that these distinctions matter less for our proposed empirical study. Finally, the claim has been made that, in the event of a ban on short selling, investors turn to options. Battalio and Schultz (2011) find strong evidence against this claim (also see Bohl, Essid and Siklos [2012] for a similar conclusion in the case of Taiwan).

8 Their empirical investigation is based on a large global sample of American Depository Receipts during the period of November 2007 to December 2010.

9 There is the risk that our sample is selectively biased. However, this is unlikely to pose a difficulty for four reasons. First, the timing, duration and precise details of the bans (see Table 1) differ across countries; second, there are other regulatory constraints (such as limitations of capital mobility) that may also play a role in how returns are internationally correlated; third, our hypothesis does not rest specifically on fundamentals to explain changes in correlation of returns; and finally, our sample includes periods when there were no short-selling bans at all. Hence, we can compare ban and no ban samples. Nevertheless, as will be emphasized below, our evidence is suggestive, not causal, and we cannot exclude the possibility that other latent factors are also at play, especially during the GFC.

10 Attempts to distinguish among types of bans did not alter our conclusions.

11 Beber and Pagano's (2013) dataset also includes Belgium, Czech Republic, Finland, Hong Kong (no ban imposed), Hungary, Ireland, Israel (no ban imposed), Luxembourg, Netherlands, Poland, Singapore (no ban imposed), Slovenia and Spain. Ten of the 13 countries not examined in our study are in Europe and the other three did not impose a short-selling ban.

12 We also collected data on financial sector stock indexes for those countries that imposed a ban on financial stocks. There does not seem to be any difference between focusing on aggregate- versus sector-specific indexes. Hence, in what follows, the evidence reported is based on market-wide equity returns.

Table 1: Short-selling Bans around the World

Country	Authority	Ban Introduced	Ban Repealed	Nature of the Ban
		MM/DD/YYYY		
Australia	ASIC	09/21/2008	05/25/2009	SS Ban — ALL
Austria	WB	10/26/2008	In effect	SS Ban/NSS — FIN
Canada	OSC	09/19/2008	10/08/2008	SS Ban — ALL
Denmark	Finanstilsynet	10/13/2008	11/01/2012	SS Ban/NSS — FIN*
France	AMF	22/09/2008	In effect	SS Ban/NSS — FIN
Germany	BaFin	09/20/2008	03/31/2011	NSS Ban — FIN
Greece	HCMC	10/10/2008	06/01/2009	SS Ban — ALL
India	BSE	05/04/2009	In effect	SS Ban — ALL
Indonesia	IDX	10/01/2008	04/30/2009	SS Ban — ALL
Italy	Consob	09/22/2008	In effect	SS Ban/NSS Ban — ALL
Japan	FSA	10/20/2008	07/31/2010	NSS Ban — ALL
Malaysia	Bank Negara	08/08/1997	03/24/2006	SS Ban — ALL
Norway	Kredittilsynet	10/08/2008	10/09/2008	SS Ban — FIN
Portugal	CMVM	09/23/2008	In effect	NSS Ban — FIN
South Korea	FSC	10/01/2008	06/01/2009	SS Ban — ALL*
Switzerland	SIX	09/19/2008	01/16/2009	SS Ban — FIN
UK	FSA	09/19/2008	01/16/2009	SS Ban — FIN
US	SEC	09/19/2008	10/08/2008	SS Ban — FIN

Note: * unless for hedging; some ban on naked short-selling still in place. ALL means all stocks; FIN means financial stocks; SS ban refers to a ban on short selling; and NSS refers to a ban on naked short selling.

METHODOLOGY

Our aim is to investigate the interdependencies between bans around the world. For this reason we estimate DCC multivariate GARCH (MGARCH) models developed by Engle (2002).[13] The multivariate DCC-MGARCH model provides all possible correlations for the index returns included in our set. Therefore, we are able to study the behaviour of the returns during periods of particular interest. Our empirical estimation of the cross-country spillovers in the adoption of short-selling restrictions on stock markets is based on the following specification:

$$r_t^i = \alpha_{i,0} + \sum_{k=1}^{4} \alpha_{i,k} D_{kt} + \beta_{i,1} r_{t-1}^i + \beta_{i,2} r_{t-1}^{UK} + \beta_{i,3} r_{t-1}^i D_t^{UK} + \beta_{i,4} D_t^i + \varepsilon_{i,t} \quad (1)$$

$$h_{ii,t} = \omega_i + \gamma_{i,1} \varepsilon_{i,t-1}^2 + \gamma_{i,2} h_{ii,t-1} \quad (2)$$

$$q_{ij,t} = \bar{\rho}_{ij}(1 - a - b) + b q_{ij,t-1} + a \xi_{i,t-1} \xi_{j,t-1} \quad (3)$$

$$\rho_{ij,t} = \frac{q_{ij,t}}{\sqrt{q_{ii,t}}\sqrt{q_{jj,t}}} \text{ for } i \neq j \quad (4)$$

Equation (1) is the mean equation; equations (2) and (3) are the variance equations; and (4) represents the DCC. Returns for country i (r_t^i) are defined as 100 times the logarithmic difference in the levels of the indices $r_t^i = \ln P_t^i - \ln P_{t-1}^i$, and the residual term in equation (1), $\varepsilon_{i,t} = N(0, h_{ii,t})$ denotes the unpredictable component of stock index returns. $\sum_{k=1}^{4} D_{kt}$ is a day of the week dummy variable for Tuesday, Wednesday, Thursday and Friday ($k = 1,\ldots,4$). D_t^i is a dummy that captures the period of ban on short-selling for country i. Hence, $\beta_{i,4}$ measures the own-country impact of the short-selling ban. The dummy variable takes on the value of 1 for days of ban and 0 otherwise.

To estimate interdependence in the imposition of short-selling bans, we interact the dummy variable for bans in the UK with the returns of each country. We consider the transmission of bans from the UK to other countries, given that UK is among the first countries to impose a ban around the world and the fact that its financial market is one of the

13 Empirical findings on exponential GARCH, MGARCH and DCC models are available on request, largely because none of the conclusions reported below are affected. Moreover, it is well-known that MGARCH models easily become over-parameterized, and this is especially the case when the investigator wishes to allow for asymmetric effects. Also, there are a number of other outstanding statistical issues around the estimation of such models that remain unanswered (Silvennoinen and Teräsvirta 2008).

most important.[14] Therefore, we are especially interested in estimates for $\beta_{i,3}$ since this coefficient captures return behaviour in country i, conditional on the UK's decision to ban short selling. In other words, the interaction term represents a proxy for the cross-country spillover effects from the imposition of the short-selling bans. The UK is a global financial centre, second only to the US, and the first jurisdiction to have imposed a ban.[15]

We also add a variable to capture the purely domestic impact of short-sale restrictions. Therefore, $\beta_{i,4}$ measures the effect of local short-selling bans in market i on local returns. The addition of lagged stock index returns, r_{t-1}, serves to detect autocorrelation of returns.

While the DCC approach has the virtue of simplicity and can deal with the curse of dimensionality in a straightforward manner, which is critical when estimating relatively large systems of equations, it is no panacea. G. P. Aielli (2013) points out that DCCs can be inconsistently estimated, while M. Caporin and M. McAleer (2012) suggest that the MGARCH approach is preferable, at least in small systems, because the standardization employed in typical DCC estimation is not unique (Caporin and McAleer 2013 highlight other problems). Nevertheless, the literature seems to strongly support the DCC technique as a useful diagnostic tool (see also footnote 13).

EMPIRICAL RESULTS

For ease of exposition we present the results of the multivariate DCC-MGARCH model estimated for a variety of country groupings chosen according to geographical areas.[16] As noted previously, all data are daily and cover the sample from December 1995 to December 2013. Although the focus of the analysis centres on the GFC — dated from June 7, 2007 to July 15, 2010 — there were other crises that preceded it. The most notable of these are listed in Figure 1.

Because of events in Europe since 2010, we also further subdivide European countries according to whether they were directly impacted economically by the sovereign debt crisis in that continent. The groupings are: Europe 1 (Norway, Denmark and Germany) and Europe 2 (Portugal, Greece and Italy). The Asia-Pacific is also subdivided by geography and proximity as follows: Asia 1 (Australia, Japan and South Korea) and Asia 2 (India, Indonesia and Malaysia). Finally, the North American continent is defined here for convenience as consisting only of Canada and the United States. Experimentation with larger country groupings, as well as different combinations of countries, did not impact the conclusions (not shown).

Table 2 presents the coefficient estimates for the spillover effects from the imposition of short-selling bans ($\beta_{i,3}$). The coefficient estimates shown consider the impact of excluding day-of-the-week dummies to determine how sensitive the results are to changes in the specification. These are intended to determine the robustness of our results. As Table 1 makes clear, the bans originated in advanced economies.

We also consider the possibility that there is an interaction effect between the ban in the UK, where this policy was first introduced, and the impact on lagged UK returns from the imposition of short-selling bans. Of course, our specification also controls for lagged UK returns to avoid confounding the impact of the ban and any other changes related to developments in UK financial markets. To the extent that too much weight might be given to the role of the UK in the global spread of short-selling bans, we also estimate a variety of factor models. These factor models (results not shown) are used to identify a "global" element in returns.[17] Hence, we provide an alternative proxy for the possibility that there exists an international component in the spread of short-selling bans that builds on the first move taken by UK authorities. Alternatively, one may view the factor model as seeking to capture the bandwagon effect of the short-sale bans as these spread across the globe. Finally, we also present estimates for subsamples that exclude or include the period of the GFC, as well as full sample estimates where UK returns are replaced by the first principal component from all the returns in the dataset. The subsample estimates are meant to address the possibility that spillover effects may have been more intensive or more likely to have become significant during the height of the financial crisis (that is, beginning in September 2008) than when at least one of the widely adopted chronologies dates the financial crisis as having been well underway by June 2007[18] (that is, before any short-sale bans were put into place).

In general, estimates of spillover effects are insensitive to the inclusion of day-of-the-week dummies (basic case shown in Table 2). Moreover, with the possible exception of India, whether or not spillover effects are found are

14 An obvious alternative is to use the United States as a benchmark. However, the US ban was imposed for such a short time (19 days) that it is doubtful the tests conducted here would be able to pick up any impact. In addition, the UK serves as a good benchmark since half the countries in our sample (nine) are in Europe. Note also that the sample in Beber and Pagano (2013) is also heavily represented by European markets.

15 The US, Canada and Switzerland also imposed bans on the same day. The US would have been the natural choice for capturing spillover effects, but the ban lasted only a few days and regulators made it clear that the ban was to be temporary. UK regulators were less clear on the length of time the ban would be in place.

16 This also simplifies problems arising from different closing hours on markets around the world.

17 This is a common approach used in the literature, especially when dealing with the effects of the GFC on financial markets. See, for example, Rogers, Scotti and Wright (2014).

18 For example, see www.stlouisfed.org/Financial-Crisis.

Table 2: Spillover Effects from Short-selling Bans — Global Evidence

Period	Basic			Subsamples		Alternative
	Jan. 1995 – Dec. 2013: Full Sample			Sept. 2008 – Jul. 2010 (GFC)	Jun. 2007 – Dec. 2013	Full Sample
	No D_{kt} (day-of-the-week dummy)	$\beta_{i,2}$ (for European Union only week dummy)	With D_{kt} (day-of-the-week dummy)	With SS ban dummy	With GFC dummy	Global factor
Country						
Canada	0.082 (.074)	-0.083 (.074)	0.083 (.071)	0.098 (.073)	0.098 (.073)	0.012 (.062)
US	-0.155**(.075)	-0.155** (.075)	-0.155** (.074)	-0.088 (.075)	-0.088 (.075)	-0.203*** (.070)
Japan	-0.120 (.077)	-0.120 (.077)	-0.123* (.068)	-0.084 (.066)	-0.085 (.066)	-0.098 (.078)
Australia	-0.161** (.085)	-0.161** (.082)	-0.162** (.077)	-0.164*** (.066)	-0.164*** (.066)	-0.149** (.073)
South Korea	-0.150* (.085)	-0.150* (.085)	-0.150* (.081)	-0.121* (.064)	-0.121* (.063)	-0.142* (.077)
India	-0.172* (.093)	-0.172* (.093)	-0.169* (.095)	-0.151 (.097)	-0.151 (.097)	-0.114 (.085)
Indonesia	-0.092 (.099)	-0.092 (.099)	-0.099 (.107)	0.050*** (.010)	0.049*** (.010)	0.098 (.097)
Malaysia	-0.175** (.083)	-0.175** (.083)	-0.174* (.091)	-0.168*** (.076)	-0.168** (.076)	-0.054 (.068)
Norway	0.052 (.065)	-0.018 (.052)	-0.018 (.054)	0.024 (.043)	0.024 (.043)	-0.008 (.050)
Germany	-0.016 (.073)	-0.132*** (.043)	-0.134*** (.044)	-0.133*** (.046)	-0.133*** (.046)	-0.084*** (.033)
Denmark	0.123** (.064)	0.076 (.049)	0.076 (.048)	0.133*** (.043)	0.133*** (.043)	0.092** (.040)
Austria	0.049 (.061)	0.063 (.052)	0.061 (.055)	0.090** (.044)	0.101* (.054)	0.075 (.056)
Switzerland	-0.073 (.063)	-0.019 (.036)	-0.020(.037)	-0.005 (.036)	0.001 (.034)	-0.035 (.046)
France	-0.074 (.059)	-0.030 (.027)	-0.031 (.026)	-0.024 (.028)	-0.021 (.029)	-0.061** (.029)
Portugal	-0.028 (.070)	0.014 (.051)	0.012 (.047)	0.029 (.051)	0.031 (.051)	-0.004 (.045)
Greece	-0.023 (.070)	0.078 (.067)	0.078 (.061)	0.123 (.072)	0.112* (.071)	0.048 (.065)
Italy	-0.011 (.069)	0.101*** (.036)	0.098*** (.035)	0.071* (.037)	0.061 (.039)	0.059 (.040)

Note: Standard errors in parenthesis. *** Means statistically significant at the 1% (** — 5%; * — 10%) level. The various dummies are described in the text.

insensitive to if UK returns are used or a single global factor is used to represent spillovers from international stock markets (alternative case shown in Table 2). Finally, there is relatively little impact from dating the sample as beginning in June 2007 versus September 2008. Only for the United States, India, Denmark and Austria are the spillover effects sensitive to sample choice (subsample case shown in Table 2). In the case of the United States, the brevity of the period of the short sale may be part of the explanation. India's equity markets are likely divorced from those in the other parts of the world included in our sample in large part because of capital controls (see Hutchison, Pasricha and Singh 2011). Therefore, estimates that focus on the period around the GFC would likely see few spillovers into India from short-selling bans in advanced economies. There is no obvious explanation for the results for Austria or Denmark. The former country is in the euro zone while Denmark remains outside the euro zone, even if its currency regime is linked to developments in the euro zone.

Elsewhere we find that the short-selling ban reduces stock returns in several economies, including Australia, South Korea, Malaysia and Germany. However, a positive response is also found, but only in a few European Union economies, namely Denmark, Austria and Italy. Since the results are sensitive to the choice of the sample, no conclusive answers can be drawn. As Table 1 shows, the type and length of the short-sale ban in these three countries also differ. Overall, there is considerable evidence of some spillover effects, which supports our contention that there was a significant global dimension to the imposition of a short-sale ban.

Table 3 asks whether, conditional on spillover effects from abroad, there are domestic repercussions to the imposition of a ban on short selling (the coefficient shown is $\beta_{i,4}$). Only for Germany is the local ban found to further depress domestic stock returns beyond the global effects of the short-sale ban first imposed by UK authorities. In the case of Korea, the results are highly sensitive to sample choice and, hence, do not appear to be reliable. In the case of Malaysia, since the results are significant for samples that focus on the period of the GFC when Malaysian authorities did not impose a short-sale ban (see Table 1), it is conceivable that the Malaysian market reflects a safe haven of sorts from the ban on short selling. Finally, for Japan and the United States, the two largest economies in the dataset, the ban is only seen to have a small return-reducing effect, either when the GFC period is assumed to

Table 3: Own-country Impact of Short-selling Bans

Country	Full	GFC	SS Ban	Global Factor
Canada	-0.176 (.950)	-0.542 (1.039)	-0.490 (.995)	-0.719 (.631)
US	-1.133 (.871)	-1.060 (.930)	-1.017 (.919)	-1.213 (.676)*
Japan	-0.009 (.048)	-0.055 (.043)	-0.077 (.042)*	-0.020 (.053)
Australia	0.045 (.126)	0.037 (.142)	0.032 (.140)	0.022 (.130)
South Korea	0.018 (.034)	0.036 (.065)	0.100 (.037)***	0.019 (.027)
India	0.057 (.035)*	-0.047 (.083)	0.124 (.116)	-0.071 (.040)*
Indonesia	0.123 (.184)	0.177 (.193)	0.128 (.202)	0.236 (.150)
Malaysia	-0.029 (.019)	0.766 (.000)***	0.201 (.000)***	-0.036 (.022)*
Norway	1.609 (2.152)	3.393 (2.353)	3.284 (1.840)*	1.493 (1.833)
Germany	-0.037 (.030)	-0.057 (.029)**	-0.051 (.042)	-0.009 (.022)
Denmark	0.015 (.022)	0.020 (.039)	0.021 (.042)	0.025 (.022)
Austria	0.006 (.034)	-0.0003 (.046)	-0.026 (.082)	-0.003 (.028)
Switzerland	-0.043 (.065)	-0.004 (.115)	-0.106 (.117)	-0.069 (.131)
France	-0.013 (.018)	-0.026 (.032)	-0.027 (.044)	-0.006 (.016)
Portugal	-0.005 (.026)	0.035 (.058)	NA	0.007 (.028)
Greece	0.009 (.074)	0.040 (.077)	NA	-0.001 (.073)
Italy	-0.017 (.026)	0.020 (.005)	NA	-0.008 (.025)

Note: * Standard errors in parenthesis. *** Means statistically significant at the 1% (** — 5%; * — 10%) level. The coefficient shown is $\beta_{i,4}$. See Table 2 for sample definitions.

Table 4: Time-varying Co-movements and Volatility Persistence

Country	a	b	DCC(1)	DCC(2)
US	0.074 (.005)	0.893 (.011)		
Canada	0.065 (.005)	0.917 (.009)	0.028 (.004)	0.967 (.005)
Australia	0.065 (.006)	0.889 (.003)		
Japan	0.063 (.006)	0.840 (.027)		
South Korea	0.063 (.005)	0.886 (.014)	0.014 (.002)	0.984 (.002)
India	0.092 (.005)	0.896 (.012)		
Indonesia	0.096 (.008)	0.906 (.011)		
Malaysia	0.084 (.007)	0.868 (.023)	0.013 (.002)	0.986 (.002)
Norway	0.051 (.006)	0.958 (.007)		
Germany	0.075 (.007)	0.908 (.013)		
Denmark	0.067 (.008)	0.855 (.025)	0.010 (.0001)	0.990 (.002)
Austria	0.051 (.004)	0.927 (.010)		
France	0.046 (.005)	0.911 (.016)		
Switzerland	0.043 (.005)	0.925 (.023)	0.021 (.002)	0.978 (.002)
Portugal	0.046 (.006)	0.895 (.022)		
Greece	0.063 (.007)	0.926 (.010)		
Italy	0.056 (.006)	0.890 (.015)	0.009 (.002)	0.990 (.002)

Note: * All estimates shown above are statistically significant at the 1% level of significance. The coefficients are from $q_{ij,t} = \bar{\rho}_{ij}(1 - a - b) + bq_{ij,t-1} + a\xi_{i,t-1}\xi_{j,t-1}$. See the text for more details.

extend back to 2007 or when global returns are considered as the mechanism through which external returns influence domestic returns.

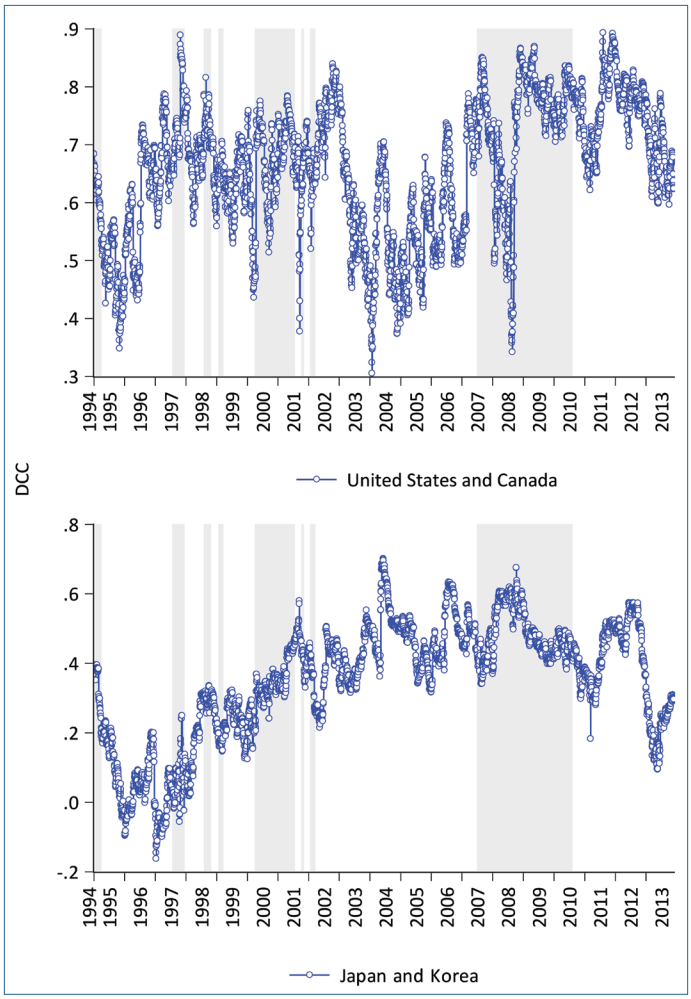
Table 4 presents a selection of DCC model estimation results. These confirm substantial time-varying co-movements in conditional volatility. Indeed, estimates of parameters a and b in equation (3), as well as the DCC estimates shown, suggest a high degree of volatility persistence. The fact that the estimates are comparable across the different grouping of countries considered also suggests that larger country groupings, which were examined (not shown), do not have a significant impact on the conclusions discussed above.

Finally, Figure 1 illustrates the additional insights obtained by estimating the DCC by focusing on two interesting cases: Canada and the United States, and Japan and Korea. Canada and the United States represent two highly integrated economies and have been so, both financially and economically, for a considerable period of time. Nevertheless, it is well known that Canada did not experience any crisis in its financial system following the events of 2008-2009, although it was pulled into a brief but relatively milder recession than its neighbour to the south. In the case of Japan and Korea, the former has been mired in a mild deflation and continues to suffer from the aftermath of bubbles that burst two decades ago, while Korea is a rapidly growing economy hard hit by the GFC.

The dynamic correlations between the United States and Canada are seen to remain high throughout the entire sample. Only during the height of the financial crisis are there signs of a sharp fall in the correlation of returns. In other words, the high degree of economic integration between the two economies, but different experiences in the fallout from the financial crisis in their respective financial sectors, did result in some decoupling. Thereafter, the historical pattern of dynamic correlations resumes.

Turning to the case of Japan and Korea, we see the rapid rise in dynamic correlations during the early 2000s as financial globalization gains pace and in spite of the dot-com financial crisis in the US. By the mid-2000s, the correlations reach levels already attained for some time between Canada and the US. It is also notable that the dynamic correlations rose during the early phases of the GFC, only to be reversed. Hence, and in spite of the differential impact of the GFC on the real and financial sectors of both economies, both equity markets are influenced by the global component of movement in stock returns. It can also be noted that Korea was one of the countries that benefitted from swap arrangements with the Fed and this, as well as the fact that Asia was less directly affected by the crisis, may also have contributed to the behaviour of dynamic correlations during the GFC.

Figure 1: DCC, Financial Crises and Short-selling Bans



Note: The vertical lines represent various major events thorough the sample, including the GFC (June 7, 2007–July 15, 2010). The others include the dot-com bubble, September 2011, the Long-Term Capital Management crisis, and Russian and other defaults.

Nevertheless, as in the Canada-US example, a complete decoupling is not evident. These two illustrations confirm that the spread of short-selling bans did not succeed in decoupling equity markets. Of course, our estimates are unable to determine the counterfactual — namely, whether stock returns might have declined even more had short-sale bans not spread globally.

CONCLUSIONS

To the extent that regulators around the globe sought to prevent further reductions in stock returns through the imposition of a ban on short selling, there is some evidence that they were unsuccessful, at least insofar as this did not factor in the possibility that there is a global component to the impact that such a ban might have had. Domestically, the bans had less or little impact, although the fact that returns do not appear to show a decline may be seen as evidence that the bans stemmed further deterioration in stock prices that policy makers sought to avoid. Whether the bans themselves can take the credit is unclear, although one might have expected further reductions in returns as a result of deteriorating financial and economic conditions after 2007.

Our results add to the analysis of short-selling bans by drawing attention to and empirically measuring the spillover effects of such bans. Even if the GFC had different economic and financial effects around the globe, regulators reacted in a similar fashion by showing a tendency to ban the short selling of stocks. Nevertheless, there were differences in both the kind of stocks banned from short-sale trades, as well as the length of time the bans were in place. It is likely that future financial crises, combined with a sharp rise in DCC across equity markets in recent years, may raise more persistent spillovers, leading to authorities reacting jointly to perceived threats in stock markets.

The proximate explanation for this development is, of course, the concern to maintain financial system stability. What is unclear is whether the global response of policy makers, who more or less simultaneously imposed short-selling bans, had unintended consequences. Based on the extant literature and the results reported in this study, it is unlikely that the degree of coupling or decoupling of global financial markets is affected. In contrast, regulatory uncertainty will have increased. Future research ought to attempt to more precisely pin down the sources, if any, of spillovers in the quality (such as price effects or bid-ask spreads) from the global imposition of bans. Moreover, in view of the potential difficutlties with DCC estimation, alternative estimation approaches are robustness tests that could be applied, in addition to the ones already considered in this paper.

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