Does Shift Contagion Exist Between OECD Stock Markets During The Financial Crisis?

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ABSTRACT

This study tests whether contagion effects exist, during the financial crisis between the U.S stock market and the OECD ones. We define shift-contagion as a significant increase in correlations in stock returns after a shock. The identification of the break point, the financial crisis, is made by the structural break test of Bai-Perron (2003). Then, time-varying correlation coefficients are estimated by the Dynamic Conditional Correlation (DCC) Multivariate GARCH Model. In order to recognize the contagion effects, we test whether the mean of the DCC coefficients in post-crisis period differs from that in the pre-crisis stable period.

Empirical findings show that the OECD stock markets have displayed a significant increase in the means of correlation coefficients between the pre-crisis and post-crisis periods. This proves the existence of contagion between the U.S and the studied markets.

Keywords: Financial Crisis; OECD Stock Markets; Shift Contagion; Multivariate DCC-GARCH Model

1. INTRODUCTION

During these last years, multiple financial crises occurred and lead to devastating effects incurred by financial markets worldwide. The recent subprime crisis, which leads to the last financial crisis, resulted in catastrophic losses. Several studies have tried to explain the reasons of these financial setbacks and the mechanisms of their spread across the international financial market. In fact, one can see, in the negative effects induced by the mortgage crisis and incurred by financial markets worldwide, a looming sign, and may wonder about the existence of a contagion phenomenon across different financial markets worldwide. The global financial Crisis of 2007-present is generally recognized as one of the most severe since the Great Depression of 1929, and it will be-well narrated in books of history and finance. Former Chief Economist at, described the global financial crisis as "a once in a 50-year event". This International Monetary Fund and Professor of Economics and Public Policy at Harvard University, Kenneth Rogoff tsunami of financial catastrophe could be traced back to the beginning of the US housing boom and to the inevitable burst (also known as the Subprime Crisis). Like other crises in history, the seeds for the Subprime Crisis were also sown in good times. The Federal Reserve reduced the Fed funds rate from 6.5 in May 2000 to 1.75% in December 2001. This led to a flood of liquidity, and the money washed through the economy like water rushing through a broken dam (2011 Commission). Lower interest rates supported by large inflows of foreign capital created easy credit conditions, which helped fuel the boom. On the one hand, the bankers and other lenders were busy in lending to anyone in search of a mortgage loan; and on the other hand, these lenders were busy in repackaging these loans into securities (CBOs and MBOs) and reselling to investors around the world. This included securitization firms and investment banks such as Merrill Lynch, Bear Stearns, and Lehman Brothers; and commercial banks and thrifts such as Citibank, Wells Fargo, and Washington Mutual. In October 2004, the Securities Exchange Commission reduced the capital requirement for 5 investment banks including Lehman Brothers, Bear Stearns and Morgan Stanley which helped these banks to leverage their investments by 30 to 40 times. These hey days came to an end when the Fed Reserve Bank decided to raise the Fed funds rate on 30 June, 2004. Till mid-2006, this rate reached a level of 5.25%. Down-turns in the housing industry can cause ripple effects almost everywhere. But this is what was not predicted, as in the words of Warren Buffet: "very, very few people could appreciate the bubble" which he called "a mass delusion" shared by "300 million Americans" (2011 Commission). By early 2004, the subprime crisis started showing signs in the form of declining housing prices, higher interest rates, and many of the mortgage loan borrowers were in no position to pay for their liabilities and started to default on their loans. Consequently, in the year 2007, subprime lenders filled bankruptcy applications. This severely affected banks and other financial institutions all over the world. Largest banks around the globe started writing down their holdings of subprime mortgage-backed securities. And ultimately, this housing bubble burst in August 2007, and the Northern Rock failed in UK, which gave birth to the global financial crisis. Equity markets play an important role in the economic growth of any nation. These markets are generally recognized as the barometer of the economic health of any nation. The country's equity markets readily indicate problems with the underlying economic factors.

The scope of this study is to contribute to the literature on international contagion effects of U.S. shocks by exploring the dynamic volatility process and detecting contagion effects during the global financial crisis.

This paper investigates the two principal questions: Is there evidence of contagion effects on the OECD countries during the global financial crisis (2007-2009)? And what are the OECD countries the most impacted?

The objective of our research is to investigate whether shift-contagion effects exist or not among the OECD stock markets during the U.S. Financial turmoil period (2007-2009). Our study extends the research into volatility spillover between stock markets. Specifically, we look at the OECD countries and the United States during the last financial crisis. So, the novelty of approach comparing to the previous ones is in the way of detecting contagion effects on OECD stock markets during the global financial crisis.

We use the Bai-Perron test (2003) for the identification of the structural break and the distinction of the period before and after the crisis. To achieve our task of identification of the shift-contagion effects, we use Engle's (2002) Dynamic Conditional Correlation (DCC) MVGARCH Model[1] for estimating time-varying correlation coefficients. Then, we test whether there is shift-contagion effect of the financial crisis on OECD stock markets or whether there are only interdependencies.

In this paper, we contribute to the literature dealing with the contagion effects of the global financial crisis on OECD stock markets. More specifically, we focus on the dynamics of the correlations between markets, and analyze whether those correlations evolve according to the stock market trend, bullish or bearish.

We pay a particular attention to the recent 2007-2009 financial crisis to detect which OECD stock markets have been impacted by its negative effects.

Policymakers pay a particular attention to stock markets and their volatility since it is a central issue for the world economy, as notably illustrated by the meetings of the G20. Moreover, analyzing the links between stock markets is of particular interest for financial players. Portfolio managers look at stock markets fluctuations to infer the trend of each market and make diversification decisions.

Comparing the impact of the financial crisis on equities prices dynamic volatility provides useful information about possible substitution strategies between stock classes. In particular, volatility plays a key role regarding hedging possibilities, and impacts asset allocation and their risk-return trade-off.

[1] The GARCH-type approach has received a special interest from almost all previous papers dealing with the issue of volatility modeling. When the objective is to investigate volatility interdependence and transmission mechanisms among different time-series, multivariate settings such as the CCC-MGARCH model of Bollerslev (1990), or the DCC-MGARCH model of Engle (2002) are more relevant than univariate models. Empirical results reported in Hassan and Malik (2007), Agnolucci (2009), and Kang et *al*. (2010), among others, confirm the superiority of these models and show that they satisfactorily capture the stylized facts of the commodity-price conditional volatility and the dynamics of volatility interaction.

The rest of the paper is organized as follows: Section 2 presents literature review on stock market contagion and empirical studies. Section 3 gives the methodology to estimate both structural change and time-varying correlation. Section 4 presents the data and the empirical analysis. Finally, section 5 provides conclusions.

2. LITERATURE REVIEW

Our study aims to bridge two strands of the literature: detecting contagion effects, in general, and investigating stock linkages on major OECD stock markets, in particular.

The studies on contagion can be broadly classified in several categories, as there is not only one definition of contagion. Surprisingly, the economists are not unanimous on a single definition of contagion[2]. For instance, that of Favero and Giavazzi (2002) who focus on financial linkages[3] and transmission process. The focus on relations between the transmissions of shocks through fundamental linkages has been primarily studied by Masson (1999), and called "Pure contagion". They used Vector Auto-regressions (VAR) models developed by Sims (1980) to offer an alternative to simultaneous equation models, and to detect contagion effects. Initially, Sims had emphasized the use of unrestricted VAR models as a means of modeling economic relationships. Another restrictive definition given by Forbes and Rigobon (2002). Contagion is defined as the increase in cross-market comovement during a crisis, as the case in our study, the use of correlation coefficients provides a straightforward method to test for the presence of contagion. This approach has been gradually improved to account for macroeconomic fundamentals and exogenous global shocks. Forbes and Rigobon (2002) explicate the interdependence effect via different channels: trade linkages, policy coordination, market reevaluation and global shocks. They further exemplify multiple equilibriums, endogenous liquidity, political economy, and other non-pre-hypothesized channels to illustrate the transmissions through non-existent channels in stable times. The presence of heteroskedasticity can lead to an increase in correlation in the crisis period, even when transmission remains unchanged. This implies that a marked increase in correlation is an insufficient proof of contagion.

Kleimeier and *al*. (2003) look at the contagion during the Asian crisis. They find little evidence of a change in the transmission mechanism from Thailand to any other country in the sample, but they do find evidence of contagion from the Hong Kong stock market. This result is in contrast with that of Forbes and Rigobon (2002). They find no evidence of contagion, but only of interdependence during the Hong Kong stock market crash and likewise during the Mexican crisis and the U.S. stock market crash.

Rigobon (2003) identifies trade links as an important channel of crisis transmission (Mexico 1994, Asia 1997 and Russia 1998). The results show that an increase in the correlation between stock markets does not result from instability in the mechanisms of propagation, but that it is rather the consequence of a strong interdependence during the crisis periods as well as during the stability periods. Although the conclusions of Rigobon (2003) are interesting, they have been considered as not robust since the size of the crisis window has an important influence on the sensitivity of the results (Dungey and Zhumabekova, 2001; and Billio and Pelizzon, 2003).

Yang (2005), applying the multivariate DCC model to daily stock index data from 1990 to 2003, investigates the conditional correlations between Japan and Hong Kong, Taiwan, South Korea, and Singapore and finds increases in the correlations, particularly when high volatilities are observed during the financial crisis.

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[2] For a more complete review, the reader can refer to the study of Dungey and *al*. (2005). They compare the correlation analysis approach popularized in this literature by Forbes and Rigobon (2002), the VAR approach of Favero and Giavazzi (2002), the probability model of Eichengreen and *al*. (1995, 1996) and the approach of Bae and *al*. (2003). They showed that the different definitions used to test for contagion are minor, and under certain conditions, are even equivalent.

[3] The relationship among national stock markets has been analyzed since the seminal work of Grubel (1968), which explained the benefits of international portfolio diversification, in a series of studies such as Granger and Morgenstern (1970), Ripley (1973), Lessard (1974, 1976) and Panton, Lessig and Joy (1976) among others. Following by the seminal works of Engle and Granger (1987) on cointegration analysis. Many studies as Johansen (1988), Johansen & Juselius (1990), Taylor & Tonks (1989), Kasa (1992) and, also, Chowdhry and *al*. (2007), among several others, have used the cointegration hypothesis to assess the international integration of financial markets.

Gravelle and *al*. (2006) study time-varying transmission mechanisms using a multivariate Markov switching approach. They employ Markov Switching VAR (MS-VAR) models aiming at identifying periods of contagion and studying the evolution of regime-dependent impulse responses. They point out the subjective and arbitrary choice of the structural change points which define the beginning and the end of the financial crisis.

Kuper and Lestano (2007) analyze the dynamic correlations of daily stock returns, exchange rates and interest rates between Indonesia and Thailand. Their results reveal that the correlations first decline at the inception of the Asian financial crisis before abrupt jumps, thereby indicating that contagion across countries may take some time.

Cheung and *al.* (2008), studying weekly stock returns in the US, East Asia, and Pacific region using the DCC model, find significant contagion effects within the East Asia and Pacific region during the international financial crisis period; however, they identify no evidence of contagion between the US and each country in the region.

Yiu and *al*. (2010), using the M-DCC model for weekly stock index data, examine the dynamic correlations between the US and eleven Asian nations and suggest the existence of contagion from the US market during the global financial crisis period.

Aloui and *al.* (2011) examines the extent of the current global crisis and its contagion effects investigating extreme financial interdependences of some selected emerging markets with the US. So, using copula on daily return data from Brazil, Russia, India, China (BRIC) and the US, show strong evidence of time-varying dependence between each of the BRIC markets and the US markets.

Guesmi and Nguyen (2011) show that correlations of international stock markets vary over time and detect an increase in correlations during periods of falling markets and a reduction in the correlation in periods of rising markets. Also, Baur (2012) study the spread of the Global Financial Crisis of 2007–2009 from the financial sector to the real economy by examining ten sectors in 25 major stock markets. They find evidence of increase co-movements in returns among financial sector stocks across countries and between financial sector stocks and real economy stocks.

Kaabia and Abid (2013) propose, so as to better characterize the transmission channels in the case of studying contagion effects during the subprime crisis, to extract constrained Bayesian factors (the transmission channels) and introduce them in FAVAR models to study the contagion process. They find that the interest rate shock appears to play an important role in the spillover mechanism from the United States to the considered OECD countries.

In our paper, we formalize the idea of contagion by testing if there is a significant shift in the degree of comovements between asset returns; we apply Engle’s (2002) Dynamic Conditional Correlation (DCC) Multivariate GARCH Model to daily stock price data (2002-2009).

In order to recognize the contagion effects, we test whether the mean of the DCC coefficients in the crisis period differs from that in the pre-crisis stable period. The identification of break point due to the crisis is made by the Bai-Perron (2003) Structural Break Test.

3. METHODOLOGY

In this section, we present the different econometric tools used in our analysis. First, we address the issue of estimating the break date in the daily U.S. stock index, the NASDAQ 100. More precisely, the Bai-Perron test (1998) is based upon an information criterion in the context of a sequential procedure, and allows finding the number of breaks implied by the data, as well as estimating the timing of the breaks and the parameters of the processes between the breaks.

The standard linear regression model is the following:

$$y_t = x_t' \beta_j + u_t \quad for \quad t = T_{j-1}+1, \ldots, T_j \quad and \quad j = 1, \ldots, m+1 \tag{1}$$

with y_t is the observation of the dependent variable, x_t is a $k \times 1$ vector of regressors, β_j is the $k \times 1$ vector of regression coefficients and u_t is the error term. The parameter m is the number of breaks.

Note that in this structural change model, all the coefficients are subject to change over time.

The hypothesis that the regression coefficients remain constant is:

$H_0 : \beta_i = \beta_0$ for $i = 1, \ldots, n$ against the alternative that at least one coefficient varies over time.

The break points $(T_1, ..., T_m)$ are explicitly treated as unknown and for $i = 1, ..., m$, we have $\lambda_i = T_i / T$ with $0 < \lambda_1 < \cdots < \lambda_m < 1$.

The purpose is to estimate the unknown regression coefficients and the break dates $(\beta_1, ..., \beta_{m+1}, T_1, ..., T_m)$ when T observations on (y_t, x_t) are available.

Bai and Perron (1998) impose some restrictions on the possible values of the break dates. Indeed, they define the following set for some arbitrary small positive number ε as the following:

$$\lambda_\varepsilon = \{(\lambda_1, \ldots, \lambda_m); |\lambda_{i+1} - \lambda_i| \geq \varepsilon, \lambda_1 \geq \varepsilon, \lambda_m \geq 1 - \varepsilon\} \tag{2}$$

This condition is made to restrict each break date to be asymptotically distinct and bounded from the boundaries of the sample.

The estimation method considered by Bai and Perron (1998) is based on the least squares. For each $m-$ partition $(T_1, ..., T_m)$, the associated least-squares estimate of β_j , noted $\hat{\beta}(T_1, ..., T_m)$ are obtained by minimizing the sum of squared residuals noted S_T. Then, the estimated break dates $(\hat{T}_1, ..., \hat{T}_m)$ are obtained as given below:

$$(\hat{T}_1, ..., \hat{T}_m) = \arg \min_{(T_1, ..., T_m)} S_T(T_1, ..., T_m) \tag{3}$$

After identification of the break date, we apply Engle's (2002) Dynamic Conditional Correlation (DCC) Multivariate GARCH Model to daily stock price data.

The multivariate model is defined as follows:

$$\begin{aligned} X_t &= \mu_t + H_t^{1/2} \varepsilon_t \\ H_t &= D_t R_t D_t' \\ R_t &= (diag(Q_t))^{-1/2} Q_t (diag(Q_t))^{-1/2} \\ D_t &= diag(\sqrt{h_{11,t}}, \sqrt{h_{22,t}}, \ldots, \sqrt{h_{NN,t}}) \end{aligned} \tag{4}$$

with $X_t = (X_{1t}, X_{2t}, \ldots, X_{Nt})$ is the vector of the past observations, $\mu_t = (\mu_{1t}, \mu_{2t}, \ldots, \mu_{Nt})$ is the vector of the conditional returns, $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}, \ldots, \varepsilon_{Nt})$ is the vector of the standardized residuals, R_t is a $(N \times N)$ symmetric dynamic correlations matrix, and D_t is a diagonal matrix of conditional standard deviations for each of the return series, with $h_{ii,t} = w_i + \alpha_i \varepsilon_{ii,t-1}^2 + \beta_i h_{ii,t-1}$. The DCC-MGARCH uses a two-stage estimation procedure. The first stage is the conventional univariate GARCH parameter estimation for each zero mean series. The residuals from the first stage are then standardized and used in the estimation of the correlation parameters in the second stage.

$$R_t = Q_t^{*-1} Q_t Q_t^{*-1} \tag{5}$$

The covariance structure is specified by a GARCH type process as below:

$$Q_t = (1 - \lambda_1 - \mu_1)\overline{Q} + \lambda_1(\eta_{t-1}\eta'_{t-1}) + \mu_1 Q_{t-1} \tag{6}$$

where the covariance matrix, $\overline{Q}$, is calculated as a weighted average of $\overline{Q}$, the unconditional covariance of the standardized residuals; $\eta_{t-1}\eta'_{t-1}$ a lagged function of the standardized residuals; and Q_{t-1} the past realization of the conditional covariance. In the DCC specification, only the first lagged realization of the covariance of the standardized residuals and the conditional covariance are used. This requires the estimation of two additional parameters, λ_1 and μ_1. Q_t^* is a diagonal matrix whose elements are the square root of the diagonal elements of Q_t .For a pair of markets i and j , the conditional correlation at time t can be written as:

$$\rho_{ij,t} = \frac{(1-\theta_1-\theta_2)\overline{q}_{ij} + \theta_1 u_{i,t-1}u_{j,t-1} + \theta_2 q_{ij,t-1}}{\left[(1-\theta_1-\theta_2)\overline{q}_{ii} + \theta_1 u_{i,t-1}^2 + \theta_2 q_{ii,t-1}\right]^{\frac{1}{2}}\left[(1-\theta_1-\theta_2)\overline{q}_{jj} + \theta_1 u_{j,t-1}^2 + \theta_2 q_{jj,t-1}\right]^{\frac{1}{2}}} \tag{7}$$

where q_{ij} is the element on the i^{th} line and i^{th} column of the matrix Q_t .

The parameters are estimated using Quasi-Maximum Likelihood Estimation (QMLE) introduced by Bollerslev and Wooldridge (1992). So, for each variable, we can obtain the conditional variance and the conditional covariance. Under the Gaussian assumption, the likelihood function can be rewritten as:

$$L(\theta) = -\frac{1}{2}\sum_{t=1}^{T}(n\log(2\pi) + 2\log|D_t| + \log|R_t| + u_t' R_t^{-1} u_t) \tag{8}$$

with $u_t = \varepsilon_t / \sqrt{h_t} = D_t^{-1}\varepsilon_t$

The estimation of the vector of unknown parameters (θ) is carried out by QMLE method which was introduced by Bollerslev and Wooldridge (1992).

4. DATA AND MAIN RESULTS

The dataset includes daily data for 17 OECD countries: USA (NASDAQ 100), Canada (TSX), Finland (Helsinki General), France (CAC 40), Germany (DAX 30), Ireland (ISEQ), Italy (Milan MIB), the Netherlands (AEX), Spain (Madrid General Index), Denmark (KFX Copenhagen), Norway (Oslo Stock Exchange), Sweden (Stockholm Index), Switzerland (Zurich Swiss Market Index), the UK (FTSE 100), Australia (All ordinaries Index), Japan (Nikkei 225), New-Zealand (New Zealand Stock Exchange 50) from 02/01/2002 to 01/06/2009. We compute the growth rates[4] and remove the mean from each series. Our dataset is primarily drawn from *EcoWin* database.

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[4] Stock return, $r_{i,t}$, is computed as the logarithmic difference of closing stock price index, $P_{i,t}$ as follows: $r_{i,t} = \log(p_{i,t} / p_{i,t-1}) \times 100$

In the presence of multiple breaks, the estimate of the break fraction will converge to one of the true break fractions, the one that is dominant in the sense that taking it into account allows the greatest reduction in the sum of squared residuals. The break date found is 01/10/2007. This break point corresponds to the financial crisis (2007-2009). This date is consistent with the speech, of the president of Federal Reserve, on October 15, 2007, admitting that the small US subprime crisis was having a large impact on global financial markets.

We divide our sample into two periods. The first period contains the observations before the crisis and the second one during the financial crisis.

Table 1 presents descriptive statistics for the stock market returns under study.

In the case of the whole period, we remark that the Denmark market displays the highest average return (0.031%), followed by the Canadian (0.011%) and the Norwegian (0.017%) stock markets.

However, if we consider separately the pre-crisis period and the crisis one, we notice that the average returns are respectively positive and negative. We observe that the mean in the OECD indices returns decreases during the crisis period compared to the pre-crisis and the entire period under study. This result is interesting and shows that all returns fall during the financial crisis.

Besides, during the financial crisis, the most impacted stock market is the Irish one with (-0.024%) followed by the Danish (-0.0214%). Also, the most resistant stock market is the Canadian one with only a fall of 0.053%.

Moreover, during the whole period, we observe that the Japan stock market is the most volatile one, while the New Zealand is the least volatile.

Also, in the financial crisis, the Japanese stock market is the most volatile with 12.7% followed by the Norwegian one with 8.67%.

So the variance increases significantly during the financial crisis compared to the pre-crisis and the entire period under study.

The coefficients of asymmetry (skewness) are positive only for Switzerland, Sweden, France, Germany and the USA. They are significantly different from zero for all stock markets, indicating the presence of asymmetry. So we can say that left-skewed distributions are predominant. This negative asymmetry denotes a potential non-linearity in the process generating the returns. In addition, all the return series are characterized by statistically significant coefficients of kurtosis greater than 3, indicating that the distribution tails are thicker than the ones of the normal distribution. All the OECD stock returns are in a leptokurtic distribution, which is a common characteristic of financial variables.

According to the mean and variance analysis, we notice a possible existence of contagion effects during the financial crisis.

To study this assumption, we estimate a multivariate DCC-GARCH model. The coefficients of GARCH (1.1) in table 2 are observed to be significant and positive; they clearly exhibit that the GARCH model captures the volatility. All the estimated parameters are statistically significant at 5% significance level. The GARCH error parameter α[5] measures the reaction of conditional volatility to market shocks. In our case, α is above 0.1 for most countries, except for the USA, Canada, and Italy. The GARCH lag parameter β[6] measures the persistence in conditional volatility irrespective of anything happening in the market. In our case, β for all the countries is equivalent or very close to 0.9 except for Japan and the Netherlands.

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[5] When α is relatively large (e.g. above 0.1), then volatility is very sensitive to market events (Carol Alexander 2008).

[6] When β is relatively large (e.g. above 0.9), then volatility takes a long time to die out following a crisis in the market (Carol Alexander 2008).

Moreover, to examine the evolution of the different dynamic correlations, and analyze their ability to track what happened during the financial crisis, we plot in Figure 1, the time-varying conditional correlations of the U.S stock market index versus one of the OCDE stock market indices under study. We assume that the contagion source is the United States and so plot each time the DCCs between the US and one of the considered OECD stock markets.

The break-point date corresponding to the financial crisis is represented by a vertical red line. These graphs clearly show variation in the dynamic conditional correlations over time. As found by the Bai-Perron test, a shift is observed in the final quarter of 2007, and reaches its peak by the end of 2008. This phenomenon further strengthens the identification of the structural break in the final quarter of 2007.

Also, the increases in DCCs beyond the break point for most countries under study are obvious. So, it appears that most of the stock markets respond to the financial crisis.

Besides, we compute the unconditional correlations and the mean of DCC coefficients in the pre-crisis and crisis periods for comparison purposes as detailed in table 3. For all countries, the unconditional correlations and the mean of DCC coefficients increase in the crisis period compared to the correlations in the pre-crisis period, as expected. This result is effectively in agreement with the ones of previous studies, including Forbes and Rigobon (2002), Pukthuanthong and Roll (2009) and Guesmi and Nguyen (2011) showing that correlations of international stock markets vary over time and detect an increase in correlations during periods of falling markets, and a reduction in the correlation in periods of rising markets.

With respect to dynamic conditional correlations, it is observed that the highest correlation exists between the US and Canada for the period before and during crisis of around 0.589 and 0.685, respectively, whereas the lowest correlation exists between the U.S and Japan for the period before and during crisis of around -0.0543 and -0.039, respectively. We also remark that for most countries, dynamic conditional correlations better predict the contagion effect than traditional correlation, as the difference between the crisis and pre-crisis is found to be greater using DCC for all the countries, except for Germany, Spain, Italy, the U.K and Japan. All the DCCs are positive, except for Japan. This denotes that the Japanese stock market may behave idiosyncratically during the financial crisis.

Also, based on the increase in the DCC mean values in percentage terms (see column 7 of Table 3), the two mostly influenced by the contagion effects are Spain and Ireland. The least impacted stock market is the German one.

To check the existence of contagion, we employ t-tests for the mean difference.

We apply the t-test to statistically verify if the dynamic conditional correlation coefficients are the same during the crisis and the pre-crisis periods. The null hypothesis tests for zero hypothesized mean difference.

We observe that the t-test fails to support the null hypothesis of zero hypothesized mean difference considering one tail at 5% significance level for all the countries, except for Germany and Italy. Thus the null hypothesis of no contagion is rejected for most of the stock markets, except for Germany and Italy.

In case of a two-tail t-test at 5% significance level, the test rejects the null hypothesis of zero hypothesized mean difference for all countries, except for Germany, Italy and the U.K.

Therefore, we support the phenomenon of a contagion effect of the global financial crisis on most of the OECD countries.

These results provide another insight into the overall contagion effects during the financial crisis. The tests demonstrate the presence of contagion effects arising from the financial crisis for most of the OECD stock markets, except for Germany, Italy and the U.K where there were only interdependencies.

5. CONCLUSION

The aim of this paper was to investigate empirically the co-movements between US stock market and those of the other sixteen OECD countries over the period of 2002-2009 so as to study the contagion effect in the case of the last global financial crisis.

For that, we have characterized contagion as a pandemic process which happens once a local shock originating from the United States stock market spreads out to other OECD stock markets. We refer to Forbes and Rigobon (2002) by defining contagion as a positive shift in the degree of comovements between OECD asset returns during the financial crisis (2007-2009).

First, we use a DCC-MVGARCH model to study the dynamic correlations for a panel of 17 OECD countries observed over the period 02/01/2002 to 01/06/2009. We use the Bai-Perron test to estimate the break point found equal to 01/10/2007. This break point reflects the financial crisis (2007-2009).Then we estimate a DCC Multivariate GARCH (1.1) model for the period of study, and also for the pre-crisis and the financial crisis periods. The obtained coefficients were economically significant.

As pointed out in our empirical findings, there is an upward trend in the dynamic conditional correlations since October 2007 and onward in all the sample markets. This evidence is strengthened by the fact that most of the cross-market correlation coefficients exceed by 50% during the global financial crisis.

Further, the presence of frequent structural breaks in the time-path of cross-market correlation series, as evidenced in our results, encourages assessment and follow up of major stock markets and the stock market comovements in implementing an investment strategy in the US and around the world. Finally, high comovements of stock markets in times of crisis evidence contagion effects as confirmed by a number of previous studies.

Our results demonstrate the presence of shift-contagion effects arising from the financial crisis to most of the OECD stock markets, except to Germany, Italy, the U.K and to a certain extent to Japan; where there are only interdependencies. The other OECD stock markets are significantly impacted by the shift-contagion during the financial crisis (2007-2009).

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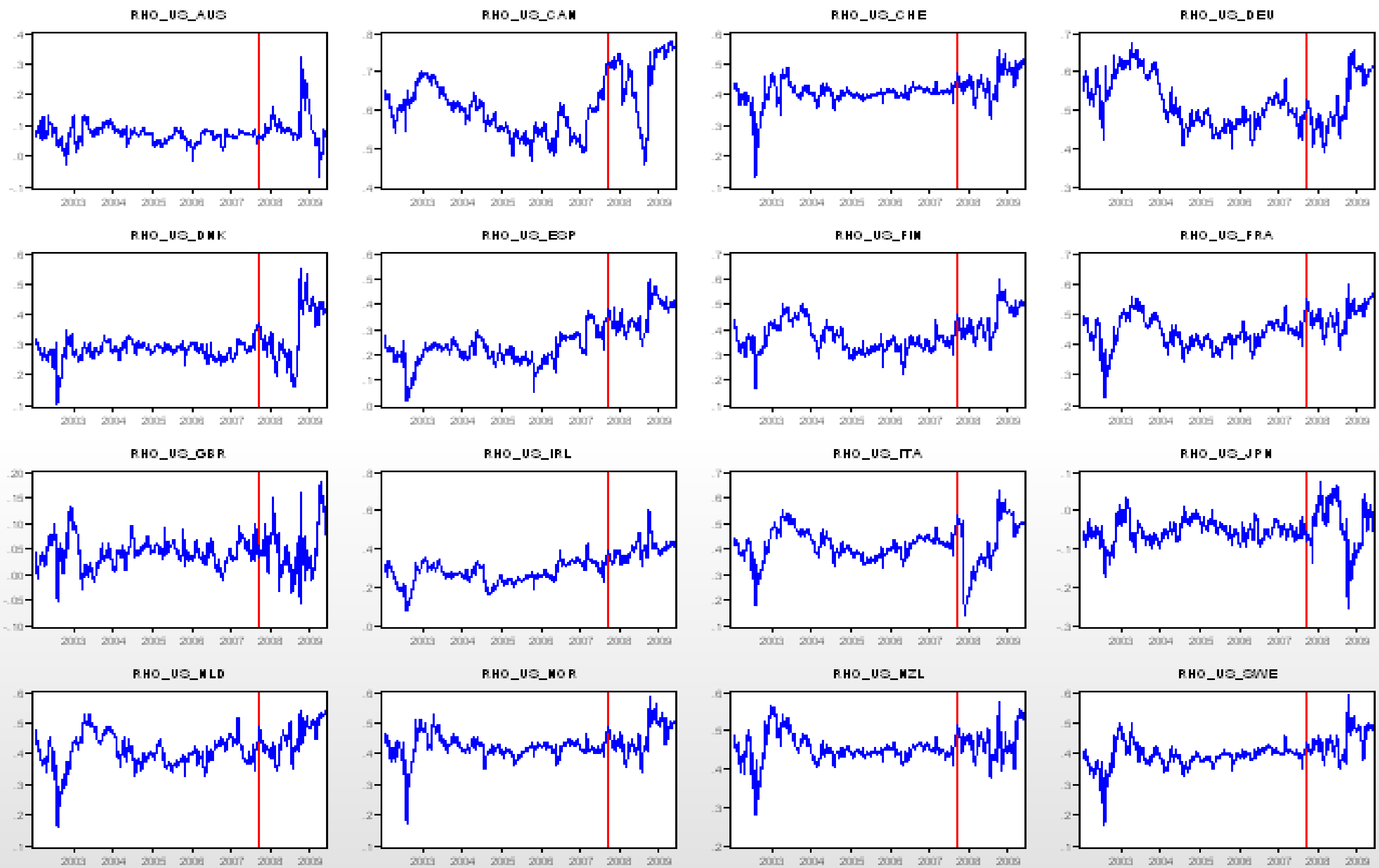

Figure 1: Dynamic Conditional Correlations between U.S and each OECD Country

Table 1: Descriptive Statistics

	Mean	Mean Pre-crisis	Mean Crisis	Variance	Var. Pre-crisis	Var. Crisis	Skew.	Skew. Pre-crisis	Skew. Crisis	Kur.	Kur. Pre-crisis	Kur. Crisis
USA	-0.0045	0.0174	-0.0797	3.0094	2.2672	5.5613	0.1271	0.2217	0.0831	7.7224	6.0142	6.6143
CAN	0.0199	0.0483	-0.0537	1.6956	0.6995	5.1181	-0.8505	-0.2839	-0.5872	13.2627	4.4500	6.1092
FIN	0.0040	0.0483	-0.1482	1.6398	0.9364	4.0328	-0.1344	-0.3887	0.1699	7.7122	5.2525	4.7258
FRA	-0.0157	0.0148	-0.1205	2.4969	1.7813	4.9497	0.0864	-0.0204	0.2433	9.2322	7.0527	7.3388
DEU	-0.0116	0.0186	-0.1152	2.7807	2.2664	4.5422	0.1185	-0.0484	0.3842	8.1152	6.3491	8.1495
IRL	-0.0381	0.0223	-0.2452	2.3948	1.0055	7.1226	-0.6847	-0.6976	-0.2655	12.4689	7.5611	5.6437
ITA	-0.0231	0.0146	-0.1525	2.0383	1.1869	4.9492	-0.0109	-0.1339	0.1743	10.5733	6.2088	6.7716
NLD	-0.0251	0.0123	-0.1537	5.6403	1.6605	5.9948	-0.2153	-0.0975	-0.1168	12.5643	7.8050	8.7419
ESP	0.0099	0.0444	-0.1087	1.8477	1.0808	4.4712	-0.0743	-0.1289	0.0929	10.5675	5.6403	6.8982
DNK	0.0319	0.1037	-0.2147	0.8933	0.4021	2.5058	-1.1224	-1.2789	-0.03828	15.7549	11.2786	7.5670
NOR	0.0175	0.0598	-0.1275	3.1404	1.5254	8.6745	-0.6503	-0.3829	-0.4106	10.2982	5.6158	5.3162
SWE	0.0031	0.0340	-0.1030	2.1514	1.4195	4.6590	0.0685	-0.0648	0.2357	7.5531	6.6292	5.0790
CHE	-0.0061	0.0262	-0.1170	1.6345	1.1609	3.2512	0.0178	-0.0986	0.2194	9.1047	7.4383	7.0060
GBR	-0.0076	0.0143	-0.0829	1.8293	1.1259	4.2454	-0.1099	-0.1901	0.0318	10.7046	7.8197	6.9483
AUS	0.0072	0.0444	-0.1207	1.1081	0.4854	3.2308	-0.6506	-0.2221	-0.3545	11.1478	6.1283	5.2286
JPN	-0.0041	0.0315	-0.1263	1.1779	1.0282	12.7092	-0.7904	-1.2135	0.5546	9.9614	6.7076	9.3949
NZL	-0.0038	0.0294	-0.1180	0.5058	0.3044	1.1829	-0.4065	-0.3417	-0.1167	8.1763	4.0996	5.6156

Note : Var., Skew. And Kur. Indicate respectively Variance, Skewness and Kurtosis.

Table 2: Estimation Results for GARCH (1,1)

Country	*C*	*A*	*β*
USA (US)	0.01105*	0.05438*	0.94129*
	(0.003949)	(0.006847)	(0.007822)
CANADA (CAN)	0.01215*	0.088438*	0.90242*
	(0.003972)	(0.010689)	(0.012329)
AUSTRALIA (AUS)	0.00860*	0.10523*	0.88893*
	(0.002078)	(0.009204)	(0.010289)
SWITZERLAND (CHE)	0.00860*	0.10523*	0.88894*
	(0.002078)	(0.009204)	(0.010289)
GERMANY(DEU)	0.02106*	0.09895*	0.89401*
	(0.004444)	(0.010234)	(0.010678)
DENMARK (DNK)	0.02440*	0.13209*	0.86173*
	(0.003383)	(0.012246)	(0.011036)
SPAIN (ESP)	0.01954*	0.11772*	0.87189*
	(0.003865)	(0.011691)	(0.012962)
FINLAND (FIN)	0.01419*	0.09488*	0.89742*
	(0.003721)	(0.010657)	(0.011053)
FRANCE (FRA)	0.01791*	0.09990*	0.89335*
	(0.004331)	(0.010385)	(0.010882)
UNITED KINGDOM (UK)	0.00951*	0.10979*	0.88768*
	(0.002722)	(0.011338)	(0.010411)
IRELAND (IRL)	0.02440*	0.13209*	0.86173*
	(0.003383)	(0.012246)	(0.011036)
ITALY (ITA)	0.01208*	0.08802*	0.90593*
	(0.002708)	(0.008747)	(0.009211)
JAPAN (JPN)	2.84557*	0.25587*	0.52745*
	(0.077479)	(0.025205)	(0.014525)
NETHERLANDS (NLD)	0.02134*	0.16501*	0.84333*
	(0.004391)	(0.012653)	(0.012539)
NORWAY (NOR)	0.04694*	0.11964*	0.86133*
	(0.010470)	(0.013484)	(0.014888)
NEW ZEALAND (NZL)	0.01485*	0.09596*	0.87142*
	(0.003327)	(0.012591)	(0.017031)
SWEDEN (SWE)	0.02243*	0.10657*	0.88377*
	(0.004773)	(0.010945)	(0.010979)

Notes: This table presents the estimation results of GARCH (1, 1) from January 2, 2002 to June 1, 2009. The numbers in parentheses represent associated standard errors. * Indicate that the coefficients are significant at the 5% level.

Table 3: Comparative Analysis of Unconditional Correlation and DCC

UC and DCC for the period 02-01-2002 TO 01-06-2009							
	Unconditional Correlation			***Dynamic Conditional Correlation***			
	Pre-crisis	Crisis	% Difference	Pre-Crisis	Crisis	% Difference	DCC is greater than UC
CANADA	0.5803	0.663	14.2340	0.5892	0.685	16.23	YES
FINLAND	0.3422	0.375	9.4681	0.3622	0.437	20.76	YES
FRANCE	0.4232	0.427	0.8507	0.436	0.491	12.71	YES
GERMANY	0.5406	0.585	8.1946	0.5188	0.521	0.35	NO
IRELAND	0.2333	0.231	-1.0716	0.2713	0.393	45.01	YES
ITALY	0.4452	0.463	3.8859	0.4142	0.418	0.89	NO
NETHERLANDS	0.4043	0.406	0.4699	0.4059	0.452	11.41	YES
SPAIN	0.383	0.380	-0.8355	0.2174	0.352	62.05	NO
DENMARK	0.1658	0.141	-14.8372	0.2816	0.335	19.11	YES
NORWAY	0.2004	0.209	4.1916	0.4213	0.453	7.52	YES
SWEDEN	0.3847	0.408	5.9787	0.391	0.439	12.35	YES
SWITZERLAND	0.3455	0.353	2.2287	0.4028	0.445	10.38	YES
UK	0.3529	0.344	-2.4936	0.0423	0.047	9.93	NO
AUSTRALIA	0.0376	0.035	-6.6489	0.0691	0.094	36.47	YES
JAPAN	0.0473	0.071	49.0486	-0.0543	-0.039	-27.26	NO
NEW ZEALAND	0.0523	-0.077	-247.6099	0.4525	0.466	3.03	YES

Table 4: T-Test Estimation - Two-Sample Assuming Unequal Variances

	Mean	*Variance*	*Observations*	H_0	*t Stat*
BC_RHO_US_CAN	0.59	0.00	1497.00	0.00	-22.13**
C_RHO_US_CAN	0.68	0.01	436.00		
BC_RHO_US_FIN	0.36	0.00	1497.00	0.00	-23.04**
C_RHO_US_FIN	0.44	0.00	436.00		
BC_RHO_US_FRA	0.44	0.00	1497.00	0.00	-23.44**
C_RHO_US_FRA	0.49	0.00	436.00		
BC_RHO_US_DEU	0.52	0.00	1497.00	0.00	-0.46
C_RHO_US_DEU	0.52	0.01	436.00		
BC_RHO_US_IRL	0.27	0.00	1497.00	0.00	-46.49**
C_RHO_US_IRL	0.39	0.00	436.00		
BC_RHO_US_ITA	0.41	0.00	1497.00	0.00	-0.66
C_RHO_US_ITA	0.42	0.01	436.00		
BC_RHO_US_NLD	0.41	0.00	1497.00	0.00	-17.24**
C_RHO_US_NLD	0.45	0.00	436.00		
BC_RHO_US_ESP	0.22	0.00	1497.00	0.00	-42.94**
C_RHO_US_ESP	0.35	0.00	436.00		
BC_RHO_US_DNK	0.28	0.00	1497.00	0.00	-11.91**
C_RHO_US_DNK	0.34	0.01	436.00		
BC_RHO_US_NOR	0.42	0.00	1497.00	0.00	-12.47**
C_RHO_US_NOR	0.45	0.00	436.00		
BC_RHO_US_SWE	0.39	0.00	1497.00	0.00	-21.34**
C_RHO_US_SWE	0.44	0.00	436.00		
BC_RHO_US_CHE	0.40	0.00	1497.00	0.00	-17.76**
C_RHO_US_CHE	0.44	0.00	436.00		
BC_RHO_US_GBR	0.04	0.00	1497.00	0.00	-1.84*
C_RHO_US_GBR	0.05	0.00	436.00		
BC_RHO_US_AUS	0.07	0.00	1497.00	0.00	-8.98**
C_RHO_US_AUS	0.09	0.00	436.00		
BC_RHO_US_JPN	-0.05	0.00	1497.00	0.00	-4.61**
C_RHO_US_JPN	-0.04	0.00	436.00		
BC_RHO_US_NZL	0.45	0.00	1497.00	0.00	-6.67**
C_RHO_US_NZL	0.47	0.00	436.00		

Notes: * Indicates that the t-stat is significant at 5% confidence level for one tail critical value ($\pm$ 1.65). ** Indicate that the t-stat is significant at 5% confidence level for one tail and two tail critical values of $\pm$ 1.65 and $\pm$ 1.96 respectively.