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Financial Contagion and Oil Risk

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Abstract

In this paper we test for the existence of equity market contagion originating from OECD monthly Stock Market Indices: United States, Canada, Finland, France, Germany, Ireland, Italy, Netherlands, Spain, Denmark, Norway, Sweden, Switzerland, United of Kingdom, Australia, Japan and New-Zealand. The data are collected over the period from January 1991 to May 2015. We apply an International Capital Asset Pricing Model (ICAPM) with currency risk. Our study offers the possibility to disentangle simple correlation due to fundamentals and contagion, which we define as the excess correlation that is not explained by fundamental factors. Our results show provides strong evidence of contagion effects originating in US equity markets to the OECD equity markets.

Keywords: Global financial crisis, financial contagion, Oil risk, ICAPM, GJR-DCC-GARCH.
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1. Introduction

Oil prices and the business cycle have fueled considerable debate in the macroeconomic literature, including variables such as real GDP, industrial production, unemployment, inflation and market uncertainty. To quantify the impact of oil on the economy, one can distinguish different research avenues. First, oil can be represented as the very pinnacle of cross-sectional financial asset prices. Second, price fluctuations due to dramatic market changes, but also political and regulatory decisions, seasonal variations, technology, may adversely impact a producer who uses oil as input. This creates the so called "oil risk", that is a commodity risk often hedged by major consumers.[5] Finally, the oil price fluctuations may spread off to other sectors in the economy, via contagion effects.

Whereas a large body of econometric models *à la* Fama-French typically accounts for the financial consequences of oil pricing, yet relatively few academic studies have focused on the concept of "oil risk" in a homogeneous framework. This lack of a comprehensive setting might be due to the fact that the notion of oil risk is multidimensional: it includes the sensitivity of oil and gas companies stock market value to oil price fluctuations, the exposure of importing and exporting countries to changes in the trade balance and oil security of supply, and the correlation effects between oil and stock markets. The concept of oil risk has been firstly raised by Sadorsky (2001) in its micro-economic component that is the negative impact of oil-gas price fluctuations on the stock value of Canadian firms. Since this seminal paper, a few applications have been made, enlarging the sample or the time span (see for instance El-Sharif et al. 2005; Boyer and Filion, D, 2007; Park and Ratti, 2008), or more recently looking at asymmetric effect of stock markets to increasing or decreasing oil prices (Ramos and Veiga, 2011). In a more aggregate perspective, countries exposure has been studied as well (Faff and Brailsford, 1999), distinguishing between oil importing countries (Gupta, 2008) or exporting ones (Demirer et al. 2015). With respect to these two strands of literature, this paper neglects the micro-economic aspect of companies exposure, but takes into account both importing and exporting countries, in a multifactor model, in a spirit that is close to Basher and Sadorsky (2006), who allow for both unconditional and conditional risk factors to investigate the relationship between oil price risk and emerging stock market returns, which is found to be significant and positive.

We also make a bridge with the broad literature on oil and stock markets, by studying the indirect or the direct effect of oil price fluctuations in an international CAPM market model. The paper closest to our, in this respect, is Broadstock et al. (2014), who show that additional oil price risk exposure is embedded in the traditional market beta, for the most important Asian countries. This paper takes his roots in previous studies, such as Scholtens and Wang (2008), who show the positive correlation between the oil price sensitivities and oil risk premia of NYSE-listed oil and gas firms' returns by applying the Fama-French factor model, or Mohanty and Nandha (2011) who estimate oil price risk exposures of the U.S. oil and gas sector using the Fama-French-Carhart's four-factor asset pricing model augmented with the oil price and interest rate factors. This latter paper finds that the market, book-to-market, and size factors, as well as momentum characteristics of stocks and changes in oil prices are significant determinants of oil returns. Finally, some recent papers focus on security pricing and the oil

[5] For instance, the IMF regularly publishes price outlook and risk assessment for oil and other selected commodities as depicted from futures and options markets.

risk premia,[6] a financial aspect of oil risk that we disregard in the present paper. This notion of oil price risk is inherent to securities and to the relationship between spot and future prices, two aspects that are not included in our analysis.

Regarding contagion, different measures have been tested. Forbes and Rigobon (2002) define contagion as a significant increase in cross-market linkages after a shock to one country (or group of countries). The spread of financial disturbances can therefore be tackled traditionally about the conception of correlation breakdown, or from several other methodological viewpoints. Kenourgios et al. (2011) report alternative tests of contagion under the frameworks of dynamic conditional correlation models (DCC, see for instance Chiang et al., 2007), regime-switching models (see Baele and Inghelbrecht, 2010), and copulas (see Rodriquez, 2007). Oil can also play a role in reinforcing contagion effects. Indeed, Malik and Ewing (2009) analyze the volatility transmission mechanism between five different U.S. sector indexes and oil prices. They document significant transmission of shocks and volatility between oil prices and some of the examined market sectors. Their findings support the idea of cross-market hedging and sharing of common information by investors.

Departing from previous studies, we analyze the three aspects of the economic implications of oil prices, namely *(i)* financial effects, *(ii)* oil price risk, and *(iii)* contagion spillovers in a unified and comprehensive framework. Considering an international setting, we deal with a multi-factorial model, the International Capital Asset Pricing Model (ICAPM). We consider the U.S. equity market return, which is considered as a world equity portfolio, a regional equity market return and an oil risk. We define contagion effects as an excess of correlations: the co-movements are explained by the common sources of risk. Contagion is the portion of risk that is not explained by the fundamentals part. The dimension of the correlation fluctuation depends on the factor loadings, and contagion is basically explained by the correlation of the residuals part. Our model is tested on OECD stock markets regrouped in four regions: the European Monetary Union (EMU), Asia-Pacific (AP), the Non-European Monetary Union (NEMU) and North America (NA).

Segmentation versus integration is an important process for the specification of our model. If individual stock markets and regions are perfectly integrated but unexpectedly experience their correlations coefficients rising during a sub-period of crisis, our test rejects the null hypothesis of no contagion. If, however, stock markets are strictly segmented, the increased correlations may basically be a consequence of increased factor volatility.

The model closest to ours is Harvey et al. (2005), who model the contagion effect as correlation

[6] For instance, Chiang et al. (2015) are among the first scholars to recognize that oil plays an important role in the pricing of securities, which is worth including in an empirical asset pricing model. Recently, Sim and Zhou (2015) have examined the effect that the quantiles of oil price shocks have on the quantiles of the U.S. stock returns. They highlight an asymmetric relationship whereby large, negative oil price shocks are found to affect U.S. equities positively when the U.S. market is bullish (whereas the converse empirical evidence is weak). Haugom et al. (2014) provide additional results on the information content of the CBOE Crude Oil Volatility Index (OVX) when forecasting realized volatility in the WTI futures market. In a similar vein, Hamilton and Wu (2014) document significant changes in oil futures risk premia since 2005, with the compensation to the long position smaller on average in more recent data. This observation is consistent with the claim that index-fund investing has become more important relative to commercial hedging in determining the structure of crude oil futures risk premia over time.

among the model residuals, using a multi-factor model allowing for time-varying expected returns risk prices. Their results show that there is no evidence of additional contagion caused by the Mexican crisis. However, they find economically meaningful increases in residual correlation, especially in Asia, during the Asian crisis.

The main contributions of our study to the literature on contagion effects are as follows: *(i)* it allows us to take into account the dynamics of oil risk. In fact, this is crucial in the case of international portfolio choice, *(ii)* the ICAPM makes clearly the difference between simple correlation due to fundamentals, or to contagion, and *(iii)* the model takes into account asymmetric effects and enables stock markets to vary through time.

The novelties of the paper to the literature on oil risks are twofold: *(i)* we introduce oil risks as an additional channel of contagion in the category of global/macro risks that has not been covered to date (even in the recent paper by Bekaert et al., 2014), and *(ii)* we extend Bekaert et al. (2005)'s specification to the case of the multivariate DCC setting, which has not been sought for either to our best knowledge (despite attempts to capture contagion through pure DCC or asymmetric DCC models as in the original contribution by Cappiello et al., 2006). Besides, we examine the sub-periods of crises and investigate whether our model can generate sudden increases in correlations. Our model provides a robust test for international, regional market and oil price risks. Last but not least, we test the time variation and cross-sectional patterns in intra-regional versus regional correlations.

The rest of the paper is organized as follows: in section 2, we present the multifactor model. Section 3 describes the data used. In section 4, we analyze the empirical results and finally in section 5 we conclude results.

2. Model

Our model is inspired from Bekaert et al. (2005) by retaining a three-factor model with time-varing loadings: the U.S. market return, the oil price and the regional equity portfolio return. Therefore, we take account in our framework of a local source, global and regional risk factors in addition to oil risk.

In this section, we present the international version of the conditional CAPM from a three factor setting.

2.1 CAPM from a three-factor setting to capture unexpected return

We suppose that the Purchase Power Parity (PPP) is verified, and the U.S. market acts as benchmark for the international market. The model is expressed as follows:

$$r_{i,t} = \delta_i' Z_{i,t-1} + \beta_{i,t-1}^{us} \Re_{us,t-1} + \beta_{i,t-1}^{oil} \Re_{oil,t-1} + \beta_{i,t-1}^{reg} \Re_{reg,t-1} + \beta_{i,t-1}^{oil} e_{oil,t} + \beta_{i,t-1}^{us} e_{us,t} + \beta_{i,t-1}^{reg} e_{reg,t} + e_{i,t} \quad (1)$$

with $e_{i,t} \left| \Omega_{t-1} \sim N(0, \sigma_{i,t}^2) \right.$

where $r_{i,t} = E\left((R_{i,t} / \Omega_{t-1}) - R_{f,t} \right)$ is the conditional excess returns on the national equity index of country *i*, with $R_{i,t}$ is the returns in U.S. dollar of the market i, $R_{f,t}$ is the risk-free rate and Ω_{t-1} includes all the information available at time t – 1. $\Re_{us,t-1}$, $\Re_{oil,t}$ and $\Re_{reg,t-1}$ are respectively the conditional expected excess returns on the U.S., the oil price, and regional markets. $e_{i,t}$, $e_{us,t}$, $e_{oil,t}$ and $e_{reg,t}$ are, respectively, the residual of the estimated model for the market i , the unanticipated returns of the global market, oil prices and the regional market; $Z_{i,t-1}$ is the set of local information variables available until the date t - 1 and δ_i is the vector of coefficients to be estimated. Moreover, $\beta^{us}_{i,t-1}$,$\beta^{reg}_{i,t-1}$ and $\beta^{oil}_{i,t-1}$ are the sensitivities of the market i to the U.S. market the regional one and the oil prices.

The conditional expected excess return on market *i* is :

$$\begin{aligned}\Re_{i,t-1} = E\left[r_{i,t-1} \mid \Omega_{t-1}\right] = \delta_i' Z_{i,t-1} + \left[\beta^{us}_{i,t-1} + \beta^{oil}_{i,t-1}\beta^{us}_{oil,t-1} + \beta^{reg}_{i,t-1}\varphi_{reg,t-1}\right]\left(\delta_{us}' Z_{us,t-1}\right) \\ + \beta^{reg}_{i,t-1}\left(\delta_{reg}' Z_{reg,t-1}\right) \qquad (2)\end{aligned}$$

with $\varphi_{reg,t-1} = \beta^{us}_{reg,t-1} + \beta^{oil}_{reg,t-1}\beta^{us}_{oil,t-1}$

where $\beta^{us}_{oil,t-1}$ is the sensitivity of the oil prices to the U.S. market. $\beta^{oil}_{reg,t-1}$ and $\beta^{us}_{reg,t-1}$ are the sensitivities of the regional market to the oil prices and the U.S. market. $\delta_{us}, \delta_{oil}$ and δ_{reg} are the vector of coefficients to be estimated. $Z_{us,t-1}$ contains a set of world information variables, $Z_{reg,t-1}$ includes the regional factors.

We should notice that the expected excess returns on market *i* proposed by Bekaert et al. (2005) is special case of Eq. (2) with
$\beta^{oil}_{i,t-1} = 0$ and $\beta^{oil}_{reg,t-1} = 0$

The effect of world market information originating from the United States on market *i*'s expected return has three components:

i) a direct impact, as measured by $\beta^{us}_{i,t-1}$, that would translate into :

$$\boxed{\text{US}} \xrightarrow{\beta^{us}_{i,t-1}} \boxed{\text{Market } i}$$

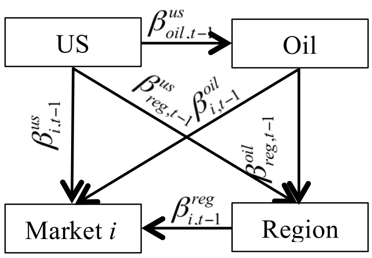
ii) indirect effect via its influence on the oil market, as measured by $(\beta^{oil}_{i,t-1} + \beta^{reg}_{i,t-1}\beta^{oil}_{reg,t-1})\beta^{us}_{oil,t-1}$, that could be presented as follows :

$$\begin{tikzcd} \boxed{\text{US}} \arrow[r, "\beta^{us}_{oil,t-1}"] & \boxed{\text{Oil}} \arrow[r, "\beta^{oil}_{reg,t-1}"] \arrow[rr, bend left, "\beta^{oil}_{i,t-1}"] & \boxed{\text{Region}} \arrow[r, "\beta^{reg}_{i,t-1}"] & \boxed{\text{Market } i} \end{tikzcd}$$

iii) a regional market effect – as measured by $\beta^{reg}_{i,t-1}\beta^{us}_{reg,t-1}$, that can be presented as follows:

$$\boxed{\text{US}} \xrightarrow{\beta^{oil}_{reg,t-1}} \boxed{\text{Region}} \xrightarrow{\beta^{us}_{i,t-1}} \boxed{\text{Market } i}$$

Our model can thus be summarized by the following graph:



In addition, the unexpected portion of the market return *i* is driven by the shocks from the local market, and also by three foreign shocks originating in the United States of America, oil risks and the region risks given by:

$$\varepsilon_{i,t} = \beta^{us}_{i,t-1} e_{us,t} + \beta^{oil}_{i,t-1} e_{oil,t} + \beta^{reg}_{i,t-1} e_{reg,t} + e_{i,t} \tag{3}$$

where $\varepsilon_{i,t}$ denotes the return residual of market *i*.

Similarly to Bekaert et al. (2005, 2011), we decompose the total variance:[7]

$$\begin{aligned}
h_{i,t} &= E\left(\varepsilon^2_{i,t} \mid \Omega_{t-1}\right) = (\beta^{us}_{i,t-1})^2 \sigma^2_{us,t} + (\beta^{oil}_{i,t-1})^2 \sigma^2_{oil,t} + (\beta^{reg}_{i,t-1})^2 \sigma^2_{reg,t} + \sigma^2_{i,t} \\
h_{i,us,t} &= E(\varepsilon_{i,t}\varepsilon_{i,t} / \Omega_{t-1}) = \beta^{us}_{i,t-1}\sigma^2_{us,t} \\
h_{i,oil,t} &= E(\varepsilon_{i,t}\varepsilon_{oil,t} / \Omega_{t-1}) = \beta^{us}_{i,t-1}\beta^{us}_{oil,t-1}\sigma^2_{us,t} + \beta^{oil}_{i,t-1}\sigma^2_{oil,t} \\
h_{i,reg,t} &= E(\varepsilon_{i,t}\varepsilon_{reg,t} / \Omega_{t-1}) = \beta^{us}_{i,t-1}\beta^{us}_{reg,t-1}\sigma^2_{us,t} + \beta^{oil}_{i,t-1}\beta^{oil}_{reg,t-1}\sigma^2_{oil,t} + \beta^{reg}_{i,t-1}\sigma^2_{reg,t} \\
h_{i,j,t} &= E(\varepsilon_{i,t}\varepsilon_{j,t} / \Omega_{t-1}) = \beta^{us}_{i,t-1}\beta^{us}_{j,t-1}\sigma^2_{us,t} + \beta^{oil}_{i,t-1}\beta^{oil}_{j,t-1}\sigma^2_{oil,t} + \beta^{reg}_{i,t-1}\beta^{reg}_{i,t-1}\sigma^2_{reg,t}
\end{aligned} \tag{4}$$

Additionally, we analyze the shares of the total variance explained respectively by the global market (VR^{us}_i), the regional one (VR^{reg}_i) and the exchange market for the country k, (VR^k_i), calculated as follows:

$$\begin{aligned}
VR^{us}_{i,t} &= \frac{(\beta^{us}_{i,t-1})^2 \sigma^2_{us,t}}{h_{i,t}} \\
VR^{oil}_{i,t} &= \frac{(\beta^{oil}_{i,t-1})^2 \sigma^2_{oil,t}}{h_{i,t}} \\
VR^{reg}_{i,t} &= \frac{(\beta^{reg}_{i,t-1})^2 \sigma^2_{reg,t}}{h_{i,t}}
\end{aligned} \tag{5}$$

Those variance ratios are proportional to the increase of the factor variances.

This preliminary analysis is necessary to investigate when returns are excessive, as a pre-condition for detecting contagion effects. Insomuch as we are interested in the crisis periods, we will investigate whether the model can generate sudden increases in correlations across markets in the aftermath of a crisis.

2.2 Contagion setting

[7] We assume that the idiosyncratic shocks of the United States, the oil price, the regional market and country *i* are uncorrelated.

As in Bekaert et al. (2005), we estimate the unexplained returns of the various markets to study contagion effects. We test the hypothesis of contagion by modeling the unexpected returns as follows:

$$
\begin{aligned}
\hat{e}_{i,t} &= \pi_i + \lambda_{i,t}\hat{e}_{m,t} + \phi_{i,t} \\
\lambda_{i,t} &= p + q_1 D_{1t} + q_2 D_{2t} \\
e_{m,t} &= e_{US,t}, e_{reg,t}, e_{oil,t}
\end{aligned}
\tag{6}
$$

We estimate the system of Eq. (6) by resorting to panel data econometrics. We consider four regions: North America, Asia-Pacific, the European Monetary Union and the Non European Monetary Union. The estimation of the model allows us to retrieve the coefficient $\lambda_{i,t}$.

To differentiate the “tranquil” period from the “turmoil” one, we use two dummy variables: D_{1t} and D_{2t}. These dummy variables allow for a change in the coefficients during the crisis. In concordance with Bekaert et al. (2005), we use such model to study the contagion phenomenon. Our model tries to uncover the sources of contagion through the various p and q coefficients.

These residual correlations are corrected for heteroskedasticity as suggested by Forbes and Rigobon (2002). The joint significance test of the parameters p and q is interpreted as a test of contagion over the entire period. In particular, testing the significance of the parameter q is interpreted as a test of increasing contagion effects in times of crisis. By adding the crisis dummy, we allow a dynamic movement to the λ coefficients during tranquil and crisis periods. If there is evidence for such a change, we call this phenomenon contagion.

The subprimes crisis was detected on the 2007:M3 and the financial one was found on the 2008:M5 using the Bai and Perron test (2002).

2.3. Estimation Method

The model in Eq. (1) can be expressed in a multivariate setting as follows:

$$r_t = \omega_{t-1}\Psi_{t-1} + \beta_{t-1}e_t$$

$$
\text{with } \omega_{t-1} = \begin{pmatrix}
1 & 0 & 0 & 0 & \cdots & 0 \\
\beta^{us}_{oil,t-1} & 1 & 0 & 0 & \cdots & 0 \\
\varphi_{reg,t-1} & \beta^{oil}_{reg,t-1} & 1 & 0 & \cdots & 0 \\
\psi_{1,t-1} & \beta^{oil}_{1,t-1} & \beta^{reg}_{1,t-1} & 1 & \cdots & 0 \\
\vdots & \vdots & \vdots & \vdots & \ddots & \vdots \\
\psi_{N,t-1} & \beta^{oil}_{N,t-1} & \beta^{reg}_{N,t-1} & 0 & \cdots & 1
\end{pmatrix}
\tag{7}
$$

$$\beta_{t-1} = \begin{pmatrix} 1 & 0 & 0 & 0 & \mathrm{L} & 0 \\ \beta^{us}_{oil,t-1} & 1 & 0 & 0 & \mathrm{L} & 0 \\ \beta^{us}_{reg,t-1} & \beta^{oil}_{reg,t-1} & 1 & 0 & \mathrm{L} & 0 \\ \beta^{us}_{1,t-1} & \beta^{oil}_{1,t-1} & \beta^{reg}_{1,t-1} & 1 & \mathrm{L} & 0 \\ \mathrm{M} & \mathrm{M} & \mathrm{M} & \mathrm{M} & \mathrm{O} & \mathrm{M} \\ \beta^{US}_{N,t-1} & \beta^{oil}_{N,t-1} & \beta^{reg}_{N,t-1} & 0 & \mathrm{L} & 1 \end{pmatrix}$$

$$r_t = \left[r_{us,t}, r_{oil,t}, r_{reg,t}, r_{1,t}, \ldots, r_{N,t} \right]'$$

$$\Psi_{t-1} = \left[\delta'_{us} Z_{us,t-1}, 0, \delta'_{reg} Z_{reg,t-1}, \delta'_1 Z_{1,t-1}, \ldots, \delta'_N Z_{N,t-1} \right]'$$

$$e_t = \left[e_{us,t}, e_{oil,t}, e_{reg,t}, e_{1,t}, \ldots, e_{N,t} \right]' \; ; \; e_t \mid \Omega_{t-1} \sim N\left(0, \Phi_t\right)$$

N is the number of countries within the particular

where $\varphi_{reg,t-1} = \beta^{us}_{reg,t-1} + \beta^{oil}_{reg,t-1}\beta^{us}_{oil,t-1}$ and $\psi_{i,t-1} = \beta^{us}_{i,t-1} + \beta^{oil}_{i,t-1}\beta^{us}_{oil,t-1} + \beta^{reg}_{i,t-1}\varphi_{reg,t-1}$.

The entire estimation process is carried out in four stages. Firstly, we estimate the parameter of the international market:

Step one:

$$r_{us,t} = \delta'_{us} Z_{us,t-1} + e_{us,t}$$

$$e_{us,t} \mid \Omega_{t-1} : \; N\left(0, \sigma_{us,t}{}^2\right) \tag{8}$$

The conditional volatility of the return series, $\sigma_{i,t}{}^2$, is modeled by a GARCH model.

Step two:

Based on the estimation results of stage 1, we estimate the model for the oil price. Conditional on Ω_{t-1} and $r_{us,t}$, we estimate the following equation:

$$r_{oil,t} = \beta^{us}_{oil,t-1}\hat{\Re}_{us,t-1} + \beta^{us}_{oil,t-1}\hat{e}_{us,t} + e_{oil,t}$$

$$e_{oil,t} \mid \Omega_{t-1} : \; N\left(0, \sigma_{oil,t}{}^2\right) \tag{9}$$

where $\hat{\Re}_{us,t-1}$ and $\hat{e}_{us,t}$ are the conditional expected excess return and residual of the U.S. market [8].

Step three:

We estimate the model for the regional market portfolio. Conditional on Ω_{t-1}, $r_{us,t}$ and $r_{oil,t}$, we estimate the following system of three equations of excess returns for each region:

[8] As the same of the first stage, the conditional volatility of the return series, $\sigma_{i,t}{}^2$, is modeled by one of the three most commonly used specifications of the GARCH family models (GARCH, EGARCH and TGARCH).

$$r_{reg,t} = \gamma_{reg,t-1} + \beta_{reg,t-1}\hat{\Re}_{t-1} + \beta_{reg,t-1}\hat{e}_t + e_{reg,t} \tag{10}$$

where $r_{reg,t} = \left(r_{reg_{EMU},t}, r_{reg_{NEMU},t}, r_{reg_{AP},t}\right)'$ refers to the (3, 1) vector of excess returns of the three emerging market regions (European Monetary Union "EMU", Non-European Monetary Union "NEMU" and Asia-Pacific "AP") which are assumed to be normally distributed.

For clarity, the regional index used is equal to the weighted average of all regional markets

$$r_{reg_i,t} = \sum_{l}^{n} \alpha_l r_{l,t} / \sum_{l}^{n} \alpha_l \tag{11}$$

where n is the number of markets in the region reg_i with $i = EMU, NEMU$ and AP and α is the market capitalization.

Step four:

The Eurodollar rate at one month, considered as the risk-free rate, is subtracted from the index returns for getting returns in excess of the risk-free rate. Data are expressed in U.S. dollars.

Also $\gamma_{reg,t-1} = \left(\delta'_{reg_{EMU}} Z_{reg_{EMU},t-1}, \delta'_{reg_{NEMU}} Z_{reg_{NEMU},t-1}, \delta'_{reg_{AP}} Z_{reg_{AP},t-1}\right)'$

$\beta_{reg,t-1} = \left(\beta^{us}_{reg,t-1}, \beta^{oil}_{reg,t-1}\right)$ where $\beta^{j}_{reg,t-1} = \left(\beta^{j}_{reg_{EMU},t}, \beta^{j}_{reg_{NEMU},t}, \beta^{j}_{reg_{AP},t}\right)'$ with j = *us, oil* , $\hat{\Re}_{t-1} = \left(\hat{\Re}_{us,t-1}, \hat{\Re}_{oil,t-1}\right)'$ and $\hat{e}_t = \left(\hat{e}_{us,t}, \hat{e}_{oil,t}\right)'$ with $\hat{\Re}_{oil,t-1}$ and $\hat{e}_{oil,t}$ are the conditional expected excess return and residual of the U.S. market.

We estimate the model in Eq. (1) for all markets, conditioning on the U.S. and regional markets model estimates:

$$r_t = \gamma_{t-1} + \beta_{t-1}\hat{\Re}_{t-1} + \beta_{t-1}\hat{e}_t + e_t \tag{12}$$

where $r_t = \left(r_{1,t}, \ldots, r_{n,t}\right)'$ refers to the *(n ,1)* vector of excess returns of markets in the region reg_i with $i = EMU, NEMU$ and AP, $\gamma_{t-1} = \left(\delta'_1 Z_{1,t-1}, \ldots, \delta'_n Z_{n,t-1}\right)'$, $\beta_{t-1} = \left(\beta^{us}_{t-1}, \beta^{oil}_{t-1}, \beta^{reg}_{t-1}\right)$ where $\beta^{j}_{t-1} = \left(\beta^{j}_{1,t}, \ldots, \beta^{j}_{n,t}\right)'$ with j = *us, oil, reg* , $\hat{\Re}_{t-1} = \left(\hat{\Re}_{us,t-1}, \hat{\Re}_{oil,t-1}, \hat{\Re}_{reg,t-1}\right)'$ and $\hat{e}_t = \left(\hat{e}_{us,t}, \hat{e}_{oil,t}, \hat{e}_{reg,t}\right)'$.

$\hat{\Re}_{reg,t-1}$ and $\hat{e}_{reg,t}$ are the conditional expected excess return and residual on the region market. The region market is calculated by the value-weighted average of all emerging regional markets, excluding the country under consideration.

Besides, $e_{reg,t} = \left(e_{reg_{EMU},t}, e_{reg_{NEMU},t}, e_{reg_{AP},t} / \Omega_{t-1}\right)' : N\left(0, H_t\right)$ is a vector of unexpected excess returns given the set of information, Ω_{t-1}, and H_t is a conditional variance-covariance matrix of excess returns following a multivariate DCC-GJR-GARCH.

3. Data

3.1 Return Series

The dataset includes 17 OECD monthly Stock Market Indices: United States, Canada, Finland, France, Germany, Ireland, Italy, Netherlands, Spain, Denmark, Norway, Sweden, Switzerland, United of Kingdom, Australia, Japan and New-Zealand. The data are collected over the period from January 1991 to May 2015. Hence, our sample period is longer than the dataset recently used by Bekaert et al. (2014) who study the time frame from 01/01/1995 to 15/03/2009. All the data are extracted from Thomson Datastream International. Moreover, we divide the OECD stock markets in four regions: North America (NA: U.S.A and Canada), European Monetary Union (EMU: Finland, France, Germany, Spain, Ireland, Italy and the Netherland), Non-European Monetary Union (NEMU: U.K, Norway, Sweden, Switzerland and Denmark) and Asia-Pacific (AP: Japan, Australia and the New-Zealand).

3.2 Local and global Instrumental Variables

The vector $Z_{us,t-1}$ contains a set of world information variables (including a constant, the world market dividend yield, the difference between the U.S. 10-year Treasury bond yield and the 3-month bill yield, and the change in the 90-day Treasury bond yield). $Z_{reg,t-1}$ includes respectively a constant, the dividend yield of the regional market portfolio, the return in excess of the regional market of the risk-free rate, and the monthly change in inflation.

3.3 Statistical Properties

Table 1: Return Series Descriptive Statistics

Countries	Mean	Std. Dev	Skewness	Kurtosis	J.B	$\rho_{i,US,t}$
USA	0.0058	0.0467	-0.7679	5.1381	76.8098	---
Canada	0.0055	0.0491	-0.8836	5.6347	111.5521	0.7581
Germany	0.0039	0.0565	-0.7239	3.6330	27.6754	0.7502
Australia	0.0044	0.0412	-0.4863	3.2424	11.1355	0.7014
Denmark	0.0065	0.0550	-0.6623	4.2482	36.7153	0.6508
Finland	0.0059	0.0899	-0.1550	3.9711	11.5174]	0.5920
Spain	0.0034	0.0555	-0.5402	3.4424	15.1076	0.7506
France	0.0034	0.0555	-0.5402	3.4424	15.1076	0.7506
United Kingdom	0.0039	0.0451	-0.5234	3.8179	19.5583	0.8057
Italy	0.0008	0.0645	0.0530	3.5618	3.6225	0.6090
Sweden	0.0063	0.0685	-0.2893	3.6164	7.9225	0.6577

Switzerland	0.0051	0.0480	-0.8209	5.5834	103.8454	0.7177
New Zealand	0.0013	0.0450	-0.3840	4.3138	25.6673	0.5185
Norway	0.0058	0.0698	-1.1197	5.8562	146.0017	0.6589
Netherlands	0.0034	0.0560	-1.2546	6.1808	181.9162	0.7816
Japan	-0.0041	0.0592	-0.3770	4.6396	36.0943	0.4742
Ireland	0.0032	0.0673	-1.0031	5.3738	107.0630	0.7035
Brent Crude Oil	0.003	0.047	-0.935	6.467	117.07	0.889

Note: We report the basic statistics of sample data over the period from 01/01/1990 to 01/12/2012.

In Table 1, the skewness coefficients are negative, showing that the tail on the right side is smaller than the left one. The values of Kurtosis exceed 3 in all cases meaning the non-normality of the return series. The rejection of the null hypothesis of normality is confirmed by the Jarque-Bera (JB) test. The Engle ARCH shows the presence of ARCH effects in the return series. The equity market returns distributions are typically non-normal and display volatility clustering and fat tail. The stylized facts of the equity returns justify our choice of using GARCH processes to model their conditional volatility.

3.4 Specification Tests and Model Estimation

The test for sign bias is based on the significance of $\phi_{1,i}$ for each market i in:[9]

$$\hat{e}_{i,t}^2 = \phi_{0,i} + \phi_{1,i} S_{i,t-1}^- + \zeta_{i,t} \tag{13}$$

where

$$S_{i,t-1}^- = \begin{cases} 1 \text{ if } \hat{e}_{i,t-1} < 0 \\ 0 \text{ otherwise} \end{cases}$$ and $\zeta_{i,t}$ is an independent and identically distributed error term.

If $\phi_{1,i}$ is statistically significant, it infers that negative and positive shocks to $\hat{e}_{i,t-1}$ impact differently upon the conditional time varying variance. We use three stages to test Eq. (14): *(i)* Estimation of the vector of unknown parameters for the U.S. market, *(ii)* Identification of the density function of the different regions' markets returns, *(iii)* Estimation of the multivariate system, conditioning on the US and regional markets model estimates.

Previous empirical works *à la* Bekaert et al. (2005) estimate the second and third steps in a univariate setting. In this paper, we consider a multivariate framework that appears more accurate when considering interactions between return series. To estimate the time-varying betas and reduce the estimation steps, the betas parameters can then be modelled as follows:

$$\beta_{i,t-1}^{us} = \frac{h_{i,us,t-1}}{h_{us,us,t-1}}, \ \beta_{i,t-1}^{reg} = \frac{h_{i,reg,t-1}}{h_{reg,reg,t-1}}, \beta_{i,t-1}^{oil} = \frac{h_{i,oil,t-1}}{h_{oil,oil,t-1}} \tag{14}$$

[9] See Engle and Ng (1993).

4. Empirical Results

4.1 Cross-patterns Co-movements during the Whole Period

First, let us focus on Table 2 that reports the estimated coefficients that measure the OECD equity markets' sensitivities to global, regional, and oil factors. For Canada, which represents with the USA the North American region, the beta with respect to the U.S. market is positive.

Table 2: Estimation Results of the loadings

Country Group	β_i^{US}	β_i^{Reg}	β_i^{Oil}
		1,018	0,191
USA	---	(0,001)	(0,011)
	0,717	0,756	0,555
Canada	(0,051)	(0,047)	(0,181)
	0,911	1,221	0,373
Finland	(0,233)	(0,285)	(0,014)
	0,185	0,883	0,103
France	(0,139)	(0,115)	(0,176)
	0,724	0,940	0,338
Germany	(0,125)	(0,146)	(0,188)
	0,825	0,818	0,049
Ireland	(0,189)	(0,216)	(0,245)
	0,677	1,108	0,266
Italy	(0,123)	(0,131)	(0,239)
	0,716	0,868	0,313
Netherland	(0,071)	(0,081)	(0,153)
	0,702	0,964	0,258
Spain	(0,076)	(0,078)	(0,166)
	0,574	0,875	0,030
Denmark	(0,102)	(0,080)	(0,042)
	0,816	1,166	0,058
Norway	(0,112)	(0,095)	(0,06)
	0,864	1,353	0,034
Sweden	(0,162)	(0,153)	(0,076)
	0,553	0,752	0,020
Switzerland	(0,068)	(0,060)	(0,036)
	0,636	0,874	0,047
United Kingdom	(0,042)	(0,030)	(0,039)
	0,503	0,288	0,234
Australia	(0,113)	(0,09)	(0,216)
	0,592	1,004	0,366
Japan	(0,101)	(0,001)	(0,296)
	0,404	0,270	0,187
New-Zealand	(0,130)	(0,107)	(0,222)

Note: Standard errors are given between parentheses.

The betas of Asian equity markets are positive and significant, varying between 0.404 and 0.592 respectively for New-Zealand and Japan, denoting that the Asian region is sensitive to the U.S. equity market. Therefore, the U.S. and the Asian-pacific factors do matter in the Asian return shocks.

Concerning the European region, the betas with respect to the U.S. market are positive and relatively high, ranging from 0.185 for France to 0.911 for Finland. Betas with respect to the regional market (the European index) are positive and very high, reaching 1.353 for Sweden.

Next, we analyze the variances ratios reported in Table 3. According to Fratzscher (2002) and Hardouvelis et al. (2006), an increase in correlations over time may result from increased volatility and/ or any change in cross-country linkages. Forbes and Rigobon (2002) show that higher volatility in one country's stock market will automatically increase the unconditional correlation in returns with another country. If volatility in one country increases, even if the transmission mechanism between the two countries is constant, a larger share of the return in the second country will be driven by the larger, idiosyncratic shocks in the first country. For Canada, the relative proportion of the conditional return variance that is accounted by the United States is positive, significant, and is the highest one. In North America, the amount of variance explained by the oil is also higher and significant. In the Asian-Pacific region, the amount of variance explained by the Asian-Pacific region is more pronounced than the one explained by the U.S. market. To give some values, 9.119%, 7.019% and 5.156% are the conditional return variances respectively for Australia, Japan, and New-Zealand and can be attributed to U.S. shocks. Moreover, the amount of variance attributed to oil is nearly the same that one explained by the Asian-Pacific region.

Unsurprisingly, the regional and oil factors account for the total variation of return shocks in Asia. The same finding is registered for the European countries. These results on betas and variance ratios give us a first explanation about the behavior of OECD equity markets towards the global, regional and oil risks, and are in line with what we would expect, given the relative idiosyncratic nature of various markets. According to our findings, we, first, remark that the country-specific beta parameter is positive, denoting that higher volatility in the U.S. market, or regional one may affect the market i. We note that the return volatility of market i is positively related to the conditional variances of the USA, the regional markets and oil risks. Potential asymmetric effects in the USA or regional markets seem to induce asymmetry in the conditional return volatility of any equity market.

Table 3: Decomposition of Total variance

Country Group	VR_i^{US} *(%)*	VR_i^{Reg} *(%)*	VR_i^{Oil} *(%)*
	North America (NA)		
US		10,018	11,024
	---	(4,112)	(5,111)
Canada	10,913	9,113	10,213
	(3,743)	(3,243)	(2,143)
	European Monetary Union (EMU)		
Finland	6,439	5,139	4,333
	(2,678)	(3,228)	(1,045)
France	0,668	0,758	0,888
	(0,370)	(0,240)	(0,112)
Germany	10,165	10,122	10,155
	(3,677)	(2,177)	(3,222)
Ireland	8,579	7,439	8,489
	(2,611)	(2,721)	(3,421)
Italy	6,455	7,255	6,111
	(2,684)	(1,784)	(1,112)
Netherland	11,210	12,110	13,145
	(3,509)	(2,109)	(3,333)
Spain	10,239	11,139	12,111

	(3,602)	(4,102)	(4,456)
	Non European Monetary Union (NEMU)		
Denmark	7,844	6,544	6,768
	(3,190)	(2,140)	(2,167)
Norway	7,916	8,916	8,678
	(3,023)	(2,021)	(2,055)
Sweden	8,180	10,280	10,110
	(3,513)	(2,413)	(2,567)
Switzerland	9,699	10,509	11,556
	(3,303)	(3,303)	(3,322)
United Kingdom	11,636	9,136	10,135
	(4,121)	(3,021)	(2,021)
	Asie-Pacific (AP)		
Australia	8,799	9,119	9,001
	(2,980)	(1,456)	(1,322)
Japan	4,019	7,019	7,044
	(1,755)	(1,765)	(1,555)
New-Zealand	5,095	5,156	5,245
	(0,905)	(0,400)	(0,55)

In Table 4 we analyze the correlations for each region with the U.S; market, the regional one and the oil risk. We remark that for North America, the correlation with the U.S. market is positive, significant, and is the highest one. Moreover, in each region, the correlations are all positive, significant and more pronounced with the USA than with the regional factor or even with the oil factor. Our results confirm those of other empirical studies. For example, Siklos and Ng (2001) showed the existence of strong interdependencies between the Asian markets and the U.S. Also, Ratanapakon and Sharma (2002) and Lim et al. (2003) showed that Asian markets are partially integrated regionally.

These cross-patterns described in this subsection capture co-movements between markets during crises as well as normal events. Therefore, although the results in this section document trends in interdependence over time, this does not necessarily capture contagion. Moreover, Forbes and Rigobon (2002) showed that higher volatility in one country's stock market will automatically increase the unconditional correlation in returns with another country. If volatility in one country increases, even if the transmission mechanism between the two countries is constant, a larger share of the return in the second country will be driven by the larger, idiosyncratic shocks in the first country. That is why we try in the next section to disentangle the contagion effects.

Table 4: Correlations

	Corr (i,US)	*Corr (i,Reg)*	*Corr (i,Oil)*
USA	-	0,02025 (0,00393)	0,01932 (0,00405)
Canada	0,02293 (0,00367)	0,02231 (0,00366)	0,02148 (0,00378)
Finland	0,08491 (0,01971)	0,08342 (0,01984)	0,08118 (0,02030)
France	0,03065 (0,00318)	0,02983 (0,00311)	0,02860 (0,00320)

Germany	0,03240 (0,00453)	0,03150 (0,00449)	0,03026 (0,00464)
Ireland	0,04656 (0,01332)	0,04540 (0,01345)	0,04295 (0,01368)
Italy	0,04260 (0,00430)	0,04157 (0,00423)	0,03974 (0,00438)
Netherland	0,03210 (0,00858)	0,03123 (0,00865)	0,03013 (0,00897)
Spain	0,03135 (0,00374)	0,03055 (0,00371)	0,02931 (0,00383)
Denmark	0,03166 (0,00539)	0,03083 (0,00539)	0,02970 (0,00557)
Norway	0,04867 (0,00660)	0,04756 (0,00658)	0,04570 (0,00680)
Sweden	0,04496 (0,00800)	0,04400 (0,00804)	0,04265 (0,00817)
Switzerland	0,02386 (0,00338)	0,02335 (0,00337)	0,02248 (0,00347)
United Kingdom	0,02122 (0,00352)	0,02070 (0,00353)	0,01972 (0,00360)
Australia	0,01699 (0,00321)	0,01652 (0,00321)	0,01574 (0,00325)
Japan	0,03352 (0,00657)	0,03284 (0,00655)	0,03123 (0,00671)
New-Zealand	0,02178 (0,00491)	0,02148 (0,00490)	0,02089 (0,00497)
N.A	0.09311 (0.00213)	0.01034 (0.02699)	0.01900 (0.10522)
E.M.U	0.07245 (0.00322)	0.07253 (0.00947)	0.07237 (0.03724)
N.E.M.U	0.07122 (0.00214)	0.04331 (0.00135)	0.03741 (0.04289)
A.P	0.04123 (0.004149)	0.02021 (0.01803)	0.02695 (0.07109)

Note: Standard errors are given between parentheses.

4.2 Time-series patterns of the residuals: Contagion Effects

The correlation detected in the previous section itself is not evidence of contagion. We will focus on studying contagion effects and are most interested in the time-series patterns of the residuals. For that, we use a panel regression of the country's idiosyncratic shocks onto a country-specific constant, and both global and regional residuals whose slope coefficients are allowed to change both in uneventful and turbulent periods.

We estimate the model described by Eq. (6), using panel data for each group of countries. We consider four groups: North America, Asia-Pacific, the Non-European Monetary Union, and the European Monetary Union. Then, we test the significance of parameters p and q. Recall that significant increases of correlations between residuals are signs of contagion. We test the existence of contagion during two specific periods: the Subprimes crisis, and the global financial one. In this analysis, we are mostly interested in the time-series patterns of these residuals. In panel A, the q_1 and q_2 coefficients measure respectively, the additional correlation during the subprime and the global financial crises. Regardless of the benchmark or region, those coefficients are positive, suggesting that the idiosyncratic residuals are better correlated during the considered crises. The correlations with respect to the U.S. index residuals are significantly higher for all regions; however, the correlations with the regional residuals are positive but not high for North America in the subprimes crisis and even are negative during the financial crisis. Considering the sum of the country-specific residuals, we find surprisingly that the correlations are less pronounced during the turmoil periods. The joint test made is an overall test of contagion. We accept at the 5% level for all the regions with respect to the U.S. index, with respect to regional return residuals, and for all regions with respect to the ''sum of other residuals'' benchmark.

Table 5: Contagion test

US. Return Residuals ($\hat{e}_{m,t} = \hat{e}_{US,t}$)					
				Wald Test	
Country	P	q_1	q_2	$\{\pi_i = 0\} \forall i$	$p = q_1 = q_2 = 0$
North America	-0.013 (0.0051)	0.790 (0.016)	0.009 (0.105)	5,021 (1,008)	0.007*** (0.003)
European Monetary Union	-0.010 (0.002)	0.897 (0.009)	0.066 (0.037)	3,111 (0,035)	0.854*** (0.007)
Non European Monetary Union	-0.011 (0.002)	0.997 (0.001)	0.027 (0.042)	7,434 (1,130)	0.027*** (0.0042)
Asia-Pacific	-0.011 (0.0041)	0.990 (0.018)	0.066 (0.071)	8,024 (2.211)	0.056*** (0.0065)
Regional. Return Residuals ($\hat{e}_{m,t} = \hat{e}_{reg,t}$)					
				Wald Test	
Country	P	q_1	q_2	$\{\pi_i = 0\} \forall i$	$p = q_1 = q_2 = 0$
North America	-0.007 (0.007)	0.011 (0.033)	-0.034 (0.132)	6,114 (1,089)	-0.011* (0.006)
European Monetary Union	-0.015 (0.001)	0.945 (0.008)	0.067 (0.031)	5,567 (1,022)	0.067** (0.031)
Non European Monetary Union	-0.010 (0.003)	0.945 (0.011)	0.067 (0.044)	6,567 (1,008)	-0.011*** (0.002)
Asia-Pacific	-0.015 (0.0042)	0.989 (0.019)	0.048 (0.072)	7,024 (1,211)	0.066 (0.0071) ***
Oil. Return Residuals (($\hat{e}_{m,t} = \hat{e}_{oil,t}$))					
				Wald Test	
Country	P	q_1	q_2	$\{\pi_i = 0\} \forall i$	$p = q_1 = q_2 = 0$
North America	0.011 (0.0032)	0.730 (0.002)	0.10 (0.111)	4,022 (0.786)	0.227*** (0. 000)
European Monetary Union	0.012 (0.001)	0.392 (0.001)	0.145 (0.032)	4.781 (0.123)	0.854*** (0.000)

Non European Monetary Union	0.024 (0.013)	0.697 (0.012)	0. 210 (0.051)	6.912 (1.230)	0.657*** (0.000)
Asia-Pacific	0.012 (0.004)	0.451 (0.022)	0.077 (0.021)	7.024 (0.112)	0.345*** (0.000)

Notes: *, **, and *** indicate that the coefficients are significant at the 10%, 5% and 1% levels respectively.

5. Conclusion

This paper studies the contagion effect in some developed stock market during the subprime crises. We use the International CAPM framework and we consider that local, regional, currency and global risk explain the co-movements part. Contagion is tested as a significant excess correlation, both in USA and developed stock markets factors, among the model residuals during calm and crisis periods. Our results show provides strong evidence of contagion effects originating in US equity markets to the European equity markets.

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