Oil Price Risk and Financial Contagion

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ABSTRACT

In this paper we test for the existence of equity market contagion, originating from oil price fluctuations, to regional and domestic stock markets. The data are collected over the period from April 1993 to April 2015. We apply an empirical multifactor asset pricing model with three-factor setting to capture the unexpected return and disentangle simple correlation due to fundamentals and contagion. We investigate four regions: the European Monetary Union (EMU), Asia-Pacific (AP), the Non-European Monetary Union (NEMU) and North America (NA). We define contagion as the excess correlation that is not explained by fundamental factors. Oil price risk is shown to be a factor as important as contagion. In addition, oil price fluctuations amplify contagion in the context of regional markets strongly interlinked with the USA.

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

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1. INTRODUCTION

The link between oil prices and the business cycle, including variables such as real GDP, industrial production, unemployment, inflation and market uncertainty, has often be debated in the macroeconomic literature (Kilian, 2008, 2014; Baumeister and Kilian, 2016). To quantify the impact of oil on the economy, one can distinguish different modeling approaches. First, in the asset pricing literature (Cochrane, 1996, 2009), cross-sectional analysis can typically inform us on how oil average returns affect different stock-, bond- or energy-dominated portfolios (see, e.g. Basher and Sadorsky, 2006 or Arouri and Rault, 2012). Second, the relationship between stock prices and oil prices has been studied using time series methods. For example, Kilian and Park (2009) quantified the relationship between shocks in global oil markets and U.S. stock returns using a structural VAR model. Additionally, in the international macroeconomics literature, capital asset pricing models (Creti and Guesmi, 2015) and multivariate GARCH (Malik and Ewing, 2009) allow identifying spillovers and crisis transmission channels stemming from the oil price, via contagion effects, and spreading off to other sectors in the economy.

Whereas a large body of econometric models à la Fama-French typically accounts for the financial consequences of oil pricing, relatively few academic studies have focused on the concept of "oil price risk" in a broad framework. This might be due to the fact that the notion of oil price 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 bal-

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ance and oil security of supply, as well as the correlation effects between oil and stock markets. The concept of oil price risk has been firstly used 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 paper, few applications have been made, enlarging the sample or the time span (see for instance El-Sharif et al. 2005; Boyer and Filion, 2007; Park and Ratti, 2008), or more recently looking at asymmetric effects of stock markets to increasing or decreasing oil prices (Ramos and Veiga, 2011). In an aggregate perspective, countries exposure has been studied (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. Our paper 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, found to be significant and positive. Our work is also related to Alquist et al. (2013) and Baumeister and Kilian (2014) who used a vector autoregressive model to quantify oil price risks and how they change under different economic scenarios. Finally, the IMF regularly publishes risk assessments for the price of oil derived from options prices.

We also contribute to the 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 ours, 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 price risk premia of NYSE-listed oil and gas firms' returns by applying the Fama-French factor model. Mohanty and Nandha (2011) 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 price risk premia. This approach 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 proposed and 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 has been shown as a factor that can reinforce 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 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. 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). The data sample span form April 1993 to May 2016. The empirical multifactor asset pricing model allows us to take into account trade flows, monetary and financial effects. All in all, we consider the U.S. equity market return, the regional equity market return and the oil price risk. In this context, we define contagion effects as an excess of correlations. Whenever co-movements are explained by the common sources of risk, contagion is the portion of risk not explained by the fundamentals part. The dimension of the correlation fluctuation depends on the factor loadings. As a consequence, contagion is basically explained by the correlation of the residuals part.

Segmentation integration and contagion play a critical role in our tests. If regional stock markets are internationally integrated for of the entire studied period but unexpectedly see their intraregional correlations increase intensely during a regional crisis, our test rejects the null hypothesis of no contagion. If, nonetheless, stock markets are segmented from the global CAPM but rather a regional CAPM, the increased correlations may simply be a consequence of increased factor volatility

Given the different components of our model, and in particular the empirical multifactor asset pricing modeling choice, segmentation versus integration plays an important role. 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 Bekaert et al. (2005), who considers the contagion effect as correlation among the model residuals, allowing for time-varying expected returns risk prices. The authors find no evidence of contagion caused by the Mexican crisis. However, a meaningful increase in residual correlation, especially in Asia, during the Asian crisis, is documented.

The main contributions of our study to the literature on contagion effects are as follows: *(i)* we take into account the dynamics of oil price risk, which is crucial in the case of international portfolio choice, *(ii)* we make clearly the difference between simple correlation due to fundamentals, or to contagion in the ICAPM, and *(iii)* we consider asymmetric effects and enable stock markets to vary through time.

The novelties of the paper with respect to the literature on oil price risks are twofold: *(i)* we introduce oil price 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 tested yet, to the best of our knowledge (despite attempts to capture contagion through pure DCC or asymmetric DCC models as in Cappiello et al., 2006). Moreover, 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. Finally, we test the time variation and cross-sectional patterns in intra-regional versus regional correlations.

Our results provide strong evidence of contagion effects originating in the US equity markets toward the European equity markets. This effect becomes evident when considering in particular correlations, in which the comovements with the US are the strongest, and in the contagion results with respect to the US index residuals. The role of oil is found significant in the variance decomposition and in the dynamic correlations, giving evidence of a factor that can amplify financial contagion, whenever it exists.

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.

2. THE MODEL

As in Bekaert et al. (2005), we use a three-factor model with time-varying loadings: the U.S. market return, the oil price and the regional equity portfolio return. Therefore, we take into account in our framework a local risk source, global and regional factors in addition to oil price risk.

The global model is the international version of the empirical multifactor asset pricing model with three-factor setting. Within this model, we specify several steps, each one devoted to a specific mechanism at stake, such as variance decomposition, contagion effects, oil price risks effects.

We assume that the Purchase Power Parity (PPP) holds, and that 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 $e_{i,t}\left|\Omega_{t-1} \sim N(0,\sigma_{i,t}^2)\right.$

$r_{i,t} = E((R_{i,t} / \Omega_{t-1}) - R_{f,t})$ 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.

$e_{us,t}$, $e_{oil,t}$ and $e_{reg,t}$ are, in the order, the unanticipated returns of the global market, oil prices and the regional market; $e_{i,t}$ is the idiosyncratic shock of any market *i*.

$Z_{i,t-1}$ is the set of local information variables available until the date t – 1 and δ_i is the vector of coefficients to be estimated.

$\beta_{i,t-1}^{us}, \beta_{i,t-1}^{reg}$ and $\beta_{i,t-1}^{oil}$ 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 :

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

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

where $\beta_{oil,t-1}^{us}$ is the sensitivity of the oil prices to the U.S. market. $\beta_{reg,t-1}^{oil}$ and $\beta_{reg,t-1}^{us}$ 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 (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.

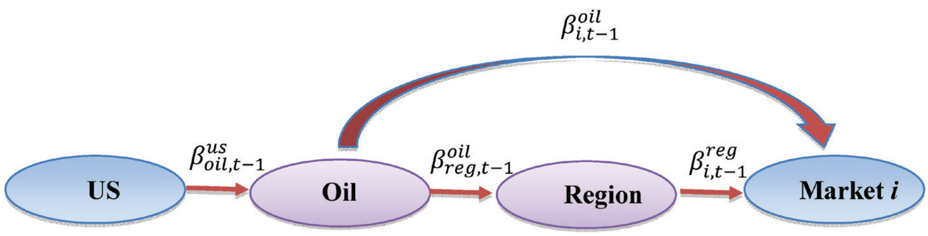
We should notice that the expected excess returns on market *i* proposed by Bekaert et al. (2005) is special case of Eq. (2) with:

$\beta_{i,t-1}^{oil} = 0$ and $\beta_{reg,t-1}^{oil} = 0$

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

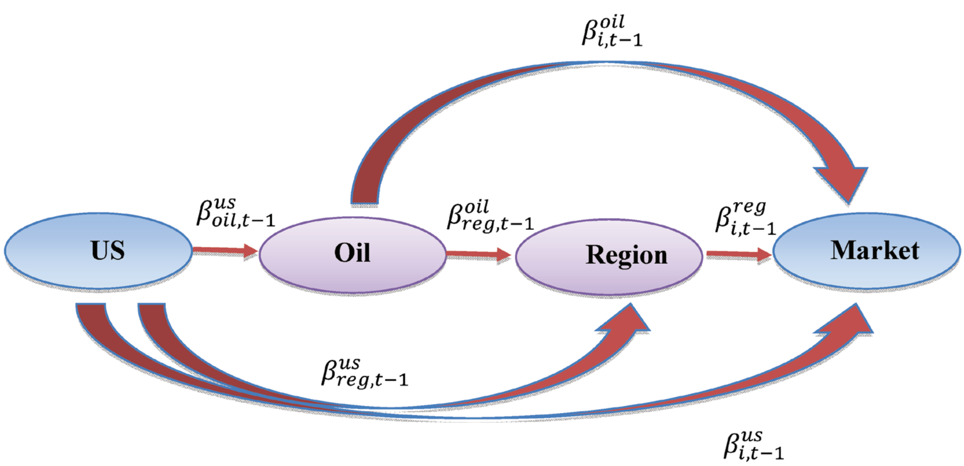
i) a direct impact, as measured by $\beta^{us}_{i,t-1}$, that would translate into :

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 :



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

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 price risks and the region risks given by the following equation:

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

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

Similarly to Bekaert et al. (2005, 2011), we decompose the total variance in five terms:[1]

$$\begin{aligned}
h_{i,t} &= E(\varepsilon_{i,t}^2 \mid \Omega_{t-1}) = (\beta_{i,t-1}^{us})^2 \sigma_{us,t}^2 + (\beta_{i,t-1}^{oil})^2 \sigma_{oil,t}^2 + (\beta_{i,t-1}^{reg})^2 \sigma_{reg,t}^2 + \sigma_{i,t}^2 \\
h_{i,us,t} &= E(\varepsilon_{i,t}\varepsilon_{i,t} / \Omega_{t-1}) = \beta_{i,t-1}^{us}\sigma_{us,t}^2 \\
h_{i,oil,t} &= E(\varepsilon_{i,t}\varepsilon_{oil,t} / \Omega_{t-1}) = \beta_{i,t-1}^{us}\beta_{oil,t-1}^{us}\sigma_{us,t}^2 + \beta_{i,t-1}^{oil}\sigma_{oil,t}^2 \\
h_{i,reg,t} &= E(\varepsilon_{i,t}\varepsilon_{reg,t} / \Omega_{t-1}) = \beta_{i,t-1}^{us}\beta_{reg,t-1}^{us}\sigma_{us,t}^2 + \beta_{i,t-1}^{oil}\beta_{reg,t-1}^{oil}\sigma_{oil,t}^2 + \beta_{i,t-1}^{reg}\sigma_{reg,t}^2 \\
h_{i,j,t} &= E(\varepsilon_{i,t}\varepsilon_{j,t} / \Omega_{t-1}) = \beta_{i,t-1}^{us}\beta_{j,t-1}^{us}\sigma_{us,t}^2 + \beta_{i,t-1}^{oil}\beta_{j,t-1}^{oil}\sigma_{oil,t}^2 + \beta_{i,t-1}^{reg}\beta_{i,t-1}^{reg}\sigma_{reg,t}^2
\end{aligned} \tag{4}$$

Additionally, we analyze the shares of the total variance explained respectively by the global market (VR_i^{us}), the regional one (VR_i^{reg}) and the oil market for the country i, (VR_i^{oil}), calculated as follows:

$$\begin{aligned}
VR_{i,t}^{us} &= \frac{(\beta_{i,t-1}^{us})^2 \sigma_{us,t}^2}{h_{i,t}} \\
VR_{i,t}^{oil} &= \frac{(\beta_{i,t-1}^{oil})^2 \sigma_{oil,t}^2}{h_{i,t}} \\
VR_{i,t}^{reg} &= \frac{(\beta_{i,t-1}^{reg})^2 \sigma_{reg,t}^2}{h_{i,t}}
\end{aligned} \tag{5}$$

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 crisis periods, we will investigate whether the model can generate sudden increases in correlations across markets in the aftermath of a crisis.

1.1 Contagion setting

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:

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

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 "stable" periods from "turmoil" ones, 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 a model to study the contagion phenomenon. Our model tries to uncover the sources of contagion through the various p and q coefficients.

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

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

3. DATA

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 April 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 01/04/2015. All the deseasonalized series are extracted from MSCI Inc. Indexes and Data. 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).

As one can see 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 (based on the bootstrap procedure proposed by Kilian and Demiroglu,

Table1: Return Series Descriptive Statistics

Countries	Mean	Std. Dev	Skewness	Kurtosis	J.B	LM ARCH (5)
USA	0.0097	0.0739	–0,2877	4,3388	28,5829 [0.000]	11.58 [0.000]
Canada	0.0056	0.0447	–0,8371	5,5574	121,566 [0.000]	3.201 [0.014]
Germany	0.0052	0.0626	–0,5426	4,7260	54,2765 [0.000]	4.605 [0.010]
Australia	0.0045	0.0383	–0,6820	3,6543	29,1435 [0.000]	5.092 [0.000]
Denmark	0.0091	0.0535	–0,3084	5,1603	72,5353 [0.000]	2.309 [0.013]
Finland	0.0074	0.0762	0,1354	4,7171	39,5637 [0.000]	6.424 [0.000]
Spain	0.0058	0.0582	–0,2275	3,6790	10,0065 [0.000]	8.991 [0.000]
France	0.0035	0.0543	–0,3747	3,4249	8,5449 [0.000]	8.996 [0.000]
United Kingdom	0.0034	0.0400	–0,5735	3,5423	19,9354 [0.000]	4.802 [0.000]
Italy	0.0033	0.0614	0,2093	3,8902	14,4626 [0.000]	5.298 [0.000]
Sweden	0.0083	0.0567	–0,2928	3,9402	16,5228 [0.000]	4.685 [0.010]
Switzerland	0.0051	0.0480	–0,8730	5,2897	106,575 [0.000]	5.703 [0.000]
New Zealand	0.0027	0.0389	–0,3872	3,8883	18,8863 [0.000]	9.519 [0.000]
Norway	0.0061	0.0645	–0,8414	5,1574	92,4377 [0.000]	9.783 [0.000]
Netherlands	0.0046	0.0574	–0,7150	4,6328	57,0661 [0.000]	14.14 [0.000]
Japan	0.0035	0.0592	–0,9465	5,6072	26,3056 [0.000]	5.692 [0.000]
Ireland	0.0067	0.0529	–0,5365	3,9073	25,4799 [0.000]	5.599 [0.004]
Brent Crude Oil	0.0046	0.0873	–0,3975	3,6461	13,9003 [0.000]	14.798 [0.000]

Note: We report the sample mean, standard deviation (sd), skewness and kurtosis of each of the series during the estimation period from April 1993 to April 2015 through the measures of skewness and kurtosis proposed by Kilian and Demiroglu (2000). We also display the normality test statistics and corresponding p-values based on the bootstrap procedure proposed by Kilian and Demiroglu (2000) and denoted by J.B. The normality is always rejected individually at 10%. The associated probabilities of the LM ARCH test are reported in brackets.

2000). 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.

4. EMPIRICAL RESULTS

4.1 Cross-patterns and 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.

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 Europe 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.

While the betas associated to oil are all positive, with some reaching high values, the regional impact of oil displays different dynamics (Figure 1 to Figure 3 below). In the NA markets, there are two distinct trends, relatively weak correlations before 2004, and wider movements afterwards, with a positive peak in the aftermath of the 2009 financial crisis, in particular for the US market, which reaches a coefficient above 0.2. In the EMU, before 2004 the betas are generally zero or negative (as perhaps in France, Ireland or Italy), whereas the values become generally positive as from 2005, with high values in Spain, the Netherlands and Germany. A similar and even more pronounced trend applies to Non-EMU countries. It can be argued that the main factor driving this effect is independent from the currency and more linked to the economic impact of oil on the real economy.

Table 2: Estimation Results of the loadings

	β_i^{us}		β_i^{reg}		β_i^{oil}	
Country Group	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
USA	—	—	0.9858	0.0246	0.0196	0.0074
Canada	0.7934	0.2440	0.8482	0.2448	0.0547	0.0154
Finland	1.3005	0.3776	1.1730	0.3155	0.0317	0.0073
France	0.1076	0.0528	0.9412	0.1647	–0.0232	0.0054
Germany	0.9599	0.2127	0.9033	0.1460	0.0819	0.0059
Ireland	1.0241	0.2370	0.7214	0.2103	–0.0222	0.0073
Italy	0.9548	0.2686	1.0577	0.2006	0.0654	0.0060
Netherland	0.9107	0.2232	0.7964	0.2103	0.0769	0.0064
Spain	0.9147	0.2422	0.9116	0.1576	0.0711	0.0055
Denmark	0.6952	0.1717	0.8818	0.09398	0.0445	0.0071
Norway	1.0160	0.3349	1.0943	0.2088	0.1029	0.0117
Sweden	0.9934	0.2374	1.2536	0.1613	0.0325	0.0092
Switzerland	0.5530	0.0680	0.7520	0.0600	–0.0180	0.0037
United Kingdom	0.7438	0.1801	0.7926	0.1373	0.0527	0.0066
Australia	0.6427	0.1833	0.3598	0.1058	0.0452	0.0044
Japan	0.6126	0.1682	1.0036	0.0827	0.0857	0.0038
New-Zealand	0.5631	0.2384	0.3293	0,1283	0.0095	0.0046

Figure 1: North America (NA) Equity markets' sensitivities to Oil prices

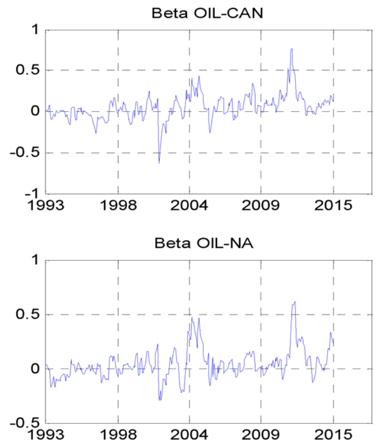


Figure 2: European Market Union (EMU) Equity markets' sensitivities to Oil prices

Conditional betas and correlations during the sub-periods of crisis are the cornerstone of our tests of contagion versus integration process. Our model allows us the capacity to decompose the increased correlation of returns into two components: the part the asset pricing model explains and the part the model does not explain. The explained part provides potential insights about market

Figure 3: Non-European Market Union (NEMU) Equity markets' sensitivities to Oil prices

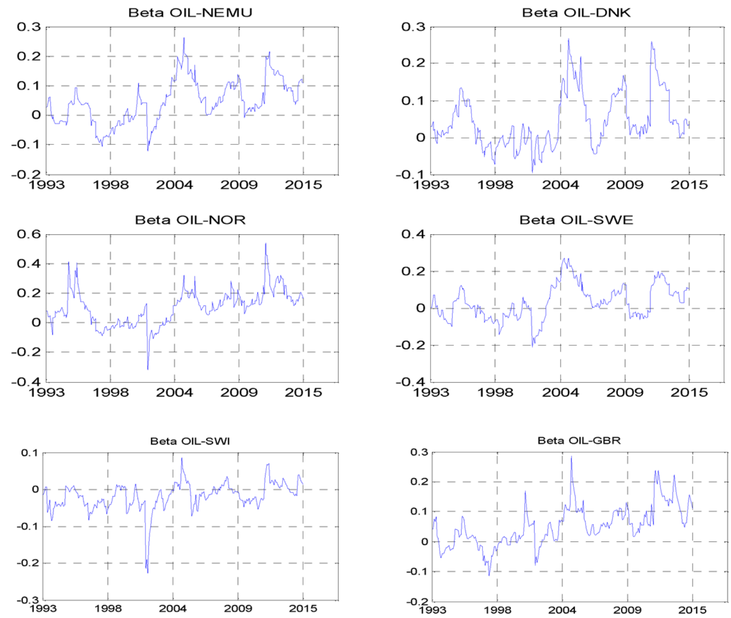
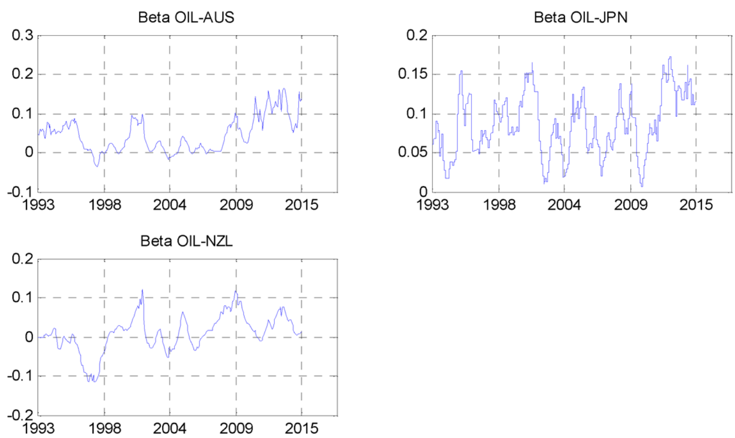


Figure 4: Asia-Pacific (AP) Equity markets' sensitivities to Oil prices



integration through the movements when betas are positives and negatives. We define contagion as the correlation of the unexplained portion.

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 sig-

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)
		Asia-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)

nificant. 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 provide us with a first explanation about the behavior of OECD equity markets towards the global, regional and oil price risks, and are in line with what we would expect, given the relative idiosyncratic nature of various markets. According to our findings, we notice 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*. The return volatility of market *i* is positively related to the conditional variances of the USA, the regional markets and oil price risks. Potential asymmetric effects in the USA or regional markets seem to induce asymmetry in the conditional return volatility of any equity market. The share of variation due to oil is quite high. The variance coefficients are of the same magnitude or even bigger than the effect of the U .S. or the regionalization. The USA, Canada, Germany, Sweden, Switzerland, the Netherland and Spain attain record values above 10.

We analyze the correlations (see table of correlations in appendix) for each region with the U.S. market, the regional one and the oil price risk. We remark that for North America, the correlation with the U.S. market is positive, significant, and 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. Correlations with the oil factor are all positive and significant. For the EMU, the three correlations are of the same magnitude, whereas for NA and AP the correlation with the U.S. is the highest; therefore, local factors as well as oil have weaker links in co-movements. 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 the reason why we try in the next section to disentangle contagion effects.

4.2 Time-series patterns of the residuals: Contagion Effects

The correlation detected in the previous section is contagion *per se*. To detect contagion, we focus on time-series patterns of the residuals. We then 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. (9), 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 that the correlations are less pronounced during the turmoil periods as perhaps some diversification process was at stake. The joint test made is an overall test of contagion. We accept

Table 4: 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	0.790	0.009	5,021	0.007***
	(0.0051)	(0.016)	(0.105)	(1,008)	(0.003)
European Monetary Union	–0.010	0.897	0.066	3,111	0.854***
	(0.002)	(0.009)	(0.037)	(0,035)	(0.007)
Non-European Monetary Union	–0.011	0.997	0.027	7,434	0.027***
	(0.002)	(0.001)	(0.042)	(1,130)	(0.0042)
Asia-Pacific	–0.011	0.990	0.066	8,024	0.056***
	(0.0041)	(0.018)	(0.071)	(2.211)	(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	0.011	–0.034	6,114	–0.011*
	(0.007)	(0.033)	(0.132)	(1,089)	(0.006)
European Monetary Union	–0.015	0.945	0.067	5,567	0.067**
	(0.001)	(0.008)	(0.031)	(1,022)	(0.031)
Non-European Monetary Union	–0.010	0.945	0.067	6,567	–0.011***
	(0.003)	(0.011)	(0.044)	(1,008)	(0.002)
Asia-Pacific	–0.015	0.989	0.048	7,024	0.066
	(0.0042)	(0.019)	(0.072)	(1,211)	(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	0.730	0.10	4,022	0.227***
	(0.0032)	(0.002)	(0.111)	(0.786)	(0. 000)
European Monetary Union	0.012	0.392	0.145	4.781	0.854***
	(0.001)	(0.001)	(0.032)	(0.123)	(0.000)
Non-European Monetary Union	0.024	0.697	0. 210	6.912	0.657***
	(0.013)	(0.012)	(0.051)	(1.230)	(0.000)
Asia-Pacific	0.012	0.451	0.077	7.024	0.345***
	(0.004)	(0.022)	(0.021)	(0.112)	(0.000)

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

at the 5% 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.

5. CONCLUSION

This paper considers the oil price as a risk factor on its own in the finance and energy literatures. Contagion effects are sought for in developed stock markets, especially during the 2008 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. The global picture that emerges is that the four regions analyzed presents similar characteristics in terms of their role of regional markets. In all of them, the oil price risk is a macroeconomic factor that strengthens the link with the USA, which is itself the source of a contagion effect. Oil therefore cannot be considered as a factor that allows diversification, especially after

2005 in all the region considered. This link should not be overlooked, especially in periods where oil volatility is very high.

Even if we acknowledge that one useful extension of this methodology could be to investigate contagion in currency markets and link equity to currency contagion, we focus on contagion transmission channels other than monetary factors. The framework to which our paper belongs is very different from the typical empirical strategy used in the international economics literature, where crisis indicators in one country (e.g., the probability of a speculative attack or the magnitude of a crisis indicator) are directly linked to indicators in other countries like currencies and exchange rates (see De Gregario and Valde´s 2001 or Salgado et al 2000). As Rigobon (1999) also underlines, this approach may not be robust when common unobservable shocks and increased variances during crisis periods affect the estimations.

With regard to international shocks transmission channels, our multi-factor CAPM analysis unveils an original oil price risk factor of its own. This new oil-related risk factor can either accelerate contagion effects (through the US region), or dampen diversification benefits. In terms of concrete implications, policy makers and portfolio managers should be wary that—amidst local, regional, currency and global risks—there is unfortunately “no place to hide” for countries seeking to hedge risks from an exposure to oil-oriented investment decisions, or for companies involved in operations specifically related to oil production and development (exploration).

One avenue for future research could be also to nest into our model a distinction between oil importers and oil exporters countries, as this aspect of international oil trade may also affect the magnitude of the oil price risk (see Creti et al. 2014). Kilian et al. (2009) have stressed the different responses of oil exporters and oil importers to shocks in oil markets, focusing on the external balance. Finally, differentiating between oil demand and oil supply shocks, as in Kilian and Park (2009), could deepen our understanding of oil price risks and contagion. This extension is also left for future research.

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APPENDIX A: *ESTIMATION METHOD*

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

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

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

$$\beta_{t-1} = \begin{pmatrix} 1 & 0 & 0 & 0 & \cdots & 0 \\ \beta^{us}_{oil,t-1} & 1 & 0 & 0 & \cdots & 0 \\ \beta^{us}_{reg,t-1} & \beta^{oil}_{reg,t-1} & 1 & 0 & \cdots & 0 \\ \beta^{us}_{1,t-1} & \beta^{oil}_{1,t-1} & \beta^{reg}_{1,t-1} & 1 & \cdots & 0 \\ \vdots & \vdots & \vdots & \vdots & \ddots & \vdots \\ \beta^{US}_{N,t-1} & \beta^{oil}_{N,t-1} & \beta^{reg}_{N,t-1} & 0 & \cdots & 1 \end{pmatrix}$$

$$r_t = \left[r_{us,t}, r_{oil,t}, r_{reg,t}, r_{1,t}, \ldots, r_{N,t}\right]'$$
$$\Psi_{t-1} = \left[\delta'_{us}Z_{us,t-1}, 0, \delta'_{reg}Z_{reg,t-1}, \delta'_1 Z_{1,t-1}, \ldots, \delta'_N Z_{N,t-1}\right]'$$
$$e_t = \left[e_{us,t}, e_{oil,t}, e_{reg,t}, e_{1,t}, \ldots, e_{N,t}\right]' \;;\; e_t \,|\, \Omega_{t-1} \sim N\left(0, \Phi_t\right)$$

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

We adopt a four-stage procedure to estimate the pricing system (7) since the simultaneous estimation of the full model is not feasible given a large number of unknown parameters. In the first stage, the conditional volatility of the return series, ${\sigma_{i,t}}^2$, is modeled by a GARCH model with

$$r_{us,t} = \delta'_{us}Z_{us,t-1} + e_{us,t} \;;\; e_{us,t} \,/\, \Omega_{t-1} \sim N(0, \sigma^2_{us,t}) \tag{8}$$

Based on the estimation results, we estimate in second step the model for the oil price as follows:

$$r_{oil,t} = \beta^{us}_{oil,t-1}\hat{\Re}_{us,t-1} + \beta^{us}_{oil,t-1}\hat{e}_{us,t} + e_{oil,t} \;;\; e_{oil,t} \,/\, \Omega_{t-1} \sim N(0, \sigma^2_{oil,t}) \tag{9}$$

where $\hat{\Re}_{us,t-1}$ and $\hat{e}_{us,t}$ are the conditional expected excess return and residual of the U.S. market.[2] In the third step we estimate the model for the regional market portfolio:

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

with $r_{reg,t} = (r_{reg_{EMU},t}, r_{reg_{NEMU},t}, r_{reg_{AP},t})'$ is the (3, 1) vector of excess returns of the three market regions (European Monetary Union "EMU", Non-European Monetary Union "NEMU" and Asia-Pacific "AP") which are assumed to be normally distributed, where $\gamma_{reg,t-1} = (\delta'_{reg_{EMU}} Z_{reg_{EMU},t-1}, \delta'_{reg_{NEMU}} Z_{reg_{NEMU},t-1}, \delta'_{reg_{AP}} Z_{reg_{AP},t-1})'$; $\beta_{reg,t-1} = (\beta^{us}_{reg,t-1}, \beta^{oil}_{reg,t-1})$ $\beta^{j}_{reg,t-1} = (\beta^{j}_{reg_{EMU},t}, \beta^{j}_{reg_{NEMU},t}, \beta^{j}_{reg_{AP},t})'$ with $j = us, oil$, $\hat{\Re}_{t-1} = (\hat{\Re}_{us,t-1}, \hat{\Re}_{oil,t-1})'$ and $\hat{e}_t = (\hat{e}_{us,t}, \hat{e}_{oil,t})'$ with $\hat{\Re}_{oil,t-1}$ and $\hat{e}_{oil,t}$ are the conditional expected excess return and residual of the U.S. market. For clarity, the regional index used is equal to the weighted average of all regional markets $r_{reg_i,t} = \sum_{l}^{n} \alpha_l r_{l,t} / \sum_{l}^{n} \alpha_l$ where n is the number of markets in the region reg_i with $i = EMU, NEMU, AP$ and α is the market capitalization. Finally, we estimate the model in Eq. (1) for all markets using DCC-GJR-GARCH model, conditioning on the U.S. and regional markets model estimates obtained from the previous steps.

Previous empirical works of 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:

$$\beta^{us}_{i,t-1} = \frac{h_{i,us,t-1}}{h_{us,us,t-1}}, \quad \beta^{reg}_{i,t-1} = \frac{h_{i,reg,t-1}}{h_{reg,reg,t-1}}, \beta^{oil}_{i,t-1} = \frac{h_{i,oil,t-1}}{h_{oil,oil,t-1}} \tag{11}$$

Although our procedure to estimate model is based on four-stage and could not be conducted in a single step due to the large number of parameters. However recently Caporin and McAleer (2013) showed that potential users of the Dynamic Conditional Correlation (DCC) representation for estimating and forecasting time-varying conditional correlations have some limits. The authors present ten properties of the Dynamic Conditional Correlation representation. Among them: DCC represents the dynamic conditional covariances of the standardized residuals, and hence does not yield dynamic conditional correlations; DCC does not have testable regularity conditions; DCC yields inconsistent two step estimators; DCC has no asymptotic properties; DCC is not a special case of Generalized Autoregressive Conditional Correlation (GARCC), which has testable regularity conditions and standard asymptotic properties etc.

2. 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).

Table of Correlations

	Corr (*i,us*)	*Corr* (*i,reg*)	*Corr* (*i,oil*)
USA	—	0,0202	0,0193
		(0,0039)	(0,0040)
Canada	0,0229	0,0223	0,0214
	(0,0036)	(0,0036)	(0,0037)
Finland	0,0849	0,0834	0,0811
	(0,0197)	(0,0198)	(0,0203)
France	0,0306	0,0298	0,0286
	(0,0031)	(0,0031)	(0,0032)
Germany	0,0324	0,0315	0,0302
	(0,0045)	(0,0044)	(0,0046)
Ireland	0,0465	0,0454	0,0429
	(0,0133)	(0,0134)	(0,0136)
Italy	0,0426	0,0415	0,0397
	(0,0043)	(0,0042)	(0,0043)
Netherland	0,0321	0,0312	0,0301
	(0,0085)	(0,0086)	(0,0089)
Spain	0,0313	0,0305	0,0293
	(0,0037)	(0,0037)	(0,0038)
Denmark	0,0316	0,0308	0,0297
	(0,0053)	(0,0053)	(0,0055)
Norway	0,0486	0,0475	0,0457
	(0,0066)	(0,0065)	(0,0068)
Sweden	0,0449	0,0440	0,0426
	(0,0080)	(0,0080)	(0,0081)
Switzerland	0,0238	0,0233	0,0224
	(0,0033)	(0,0033)	(0,0034)
United Kingdom	0,0212	0,0207	0,0197
	(0,0035)	(0,0035)	(0,0036)
Australia	0,0169	0,0165	0,0157
	(0,0032)	(0,0032)	(0,0032)
Japan	0,0335	0,0328	0,0312
	(0,0065)	(0,0065)	(0,0067)
New-Zealand	0,0217	0,0214	0,0208
	(0,0049)	(0,0049)	(0,0049)
N.A	0.0931	0.0103	0.0190
	(0.0021)	(0.0269)	(0.1052)
E.M.U	0.0724	0.0725	0.0723
	(0.0032)	(0.0094)	(0.0372)
N.E.M.U	0.0712	0.0433	0.0374
	(0.0021)	(0.0013)	(0.0428)
A.P	0.0412	0.0202	0.0269
	(0.0041)	(0.0180)	(0.0710)

Note: Standard errors are given between parentheses.