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Crude oil shocks and stock markets: A panel threshold cointegration approach

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ABSTRACT

This paper proposes a panel threshold cointegration approach to investigate the relationship between crude oil shocks and stock markets for the OECD and non-OECD panel from January 1995 to December 2009. Nonlinear cointegration is confirmed for the oil–stock nexus in the panel. Because threshold cointegration is found, the threshold vector error correction models can be run to investigate the presence of asymmetric dynamic adjustment. The Granger causality tests demonstrate the existence of bidirectional long-run Granger causality between crude oil shocks and stock markets for these OECD and non-OECD countries. However, the short-run Granger causality between them is bidirectional under positive changes in the deviation and unidirectional under negative ones. Moreover, the speed of adjustment toward equilibrium is faster under negative changes in the deviation than that under positive ones in these OECD and non-OECD countries.

1. Introduction

Hamilton (1983) indicates that crude oil price shocks were a factor in the US recession after World War II. Since then, the identification of connections between crude oil prices and the macroeconomy has been a major concern in theory and practice. A large amount of literature tries to shed light on the effects of crude oil price shocks on economic activities, such as aggregate demand, inflation, employment and real economic growth (Bachmeier, 2008; Cunado and Perez de Garcia, 2005; Hamilton, 2003). The aforementioned studies have yielded mixed results. However, in the empirical literature, only a relatively small number of works have looked into the effects of crude oil prices on the stock markets. Studies by Jones and Kaul (1996) and Sadorsky (1999) report a significant negative impact of crude oil shocks on stock returns, a result that is further supported by Papapetrou (2001). According to the latter paper, an oil price shock has a negative effect on stock returns for the first 4 months. In this line of research, however, Chen et al. (1986) and Huang et al. (1996) do not reach the same conclusions. All of these results show that there is no consensus on the relationship between crude oil shocks and stock markets; therefore, more research may be necessary on this subject.

Much of the literature thus far has focused on the connection between crude oil price changes and stock prices. Specifically, almost all publications stress the influence of crude oil shocks on stock markets combined with other economic determinants. The new literature can be categorized into four types: first, the relationship between crude oil shocks and stock markets seems to be significantly evident and negative. This observation tends to be in line with Jones and Kaul (1996) and Sadorsky (1999) (Ciner, 2001; Kilian and Park, 2007, among others). In particular, Hammoudeh and Li (2005) suggest that on a daily basis, there is a negative bidirectional dynamic relationship between crude oil price growth and the world capital market. Ghouri (2006) also reveals that there is a very strong negative relation between West Texas Intermediate Cushing (WTI) and US monthly stock positions. Miller and Ratti (2009) analyze the long-run relationship between the world price of crude oil and international stock markets, utilizing a cointegration vector error correction model (VECM) with additional regressors. Aloui and Jammazi (2009) and Chen (2010), who use Markov-switching models, obtain similar results. Moreover, Basher and Sadorsky (2006) provide strong evidence of the impact of oil price risk on emerging stock market returns. Hammoudeh and Choi (2007) and Nandha and Hammoudeh (2007) further document that oil plays an important role in emerging stock markets. They find that stock prices increase as the crude oil price decreases and decrease as the crude oil price increases. Second, the relationship between crude oil shocks and stock markets seems to be significantly evident and positive (Arouri and Rault, 2011; Chen et

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al., 1986; El-Sharif et al., 2005; Narayan and Narayan, 2010). El-Sharif et al. (2005) find that a rise in oil prices results in an increase in the returns of oil and gas markets. Narayan and Narayan (2010) provide evidence in favor of cointegration of stock prices, oil prices and nominal exchange rates, and have also found that oil prices have positive and statistically significant impact on Vietnam's stock prices. Arouri and Rault (2011) also indicate that oil price increases influence stock prices in Gulf Cooperation Council (GCC) countries positively except in Saudi Arabia. Third, some studies show that oil price shocks have a statistically significant impact on stock markets, but whether the impact is positive or negative depends on the various conditions (Cong et al., 2008; Park and Ratti, 2008). Park and Ratti (2008) demonstrate that the stock market's response to crude oil shocks partly depends on whether the country is an oil importer or oil exporter. Cong et al. (2008) investigate the interactions between oil price shocks and the Chinese stock market, finding that the relative importance of oil price shocks and interest rates is different in various conditions of the Chinese market. Finally, the fourth category of research concludes that there is no significant relationship between oil shocks and stock markets (Al Janabi et al., 2010; Apergis and Miller, 2009; Henriques and Sadorsky, 2008). Using a four-variable vector autoregression model, Henriques and Sadorsky (2008) show that shocks to oil prices do not have a significant impact on the stock prices of alternative energy companies. Apergis and Miller (2009) investigate the effects of oil market shocks on stock markets in a multi-country context. The results also indicate that international stock market returns only respond to oil market shocks in a minor way. Al Janabi et al. (2010) conduct an empirical study of the impact of oil and gold prices on the GCC stock markets. The empirical findings reveal that neither gold nor oil prices Granger-cause the stock price index in each market.

Within the debate on the role of crude oil prices in stock markets, the previous analysis is conducted mainly with the assumption that the underlying variables exhibit a linear and symmetrical adjustment process. In reality, however, asymmetry is manifested in most macro variables. In finance variables, for example, asymmetry may be caused by the fact that bullish and bearish markets behave differently; stock markets may respond in various ways to increases and decreases in the price of crude oil. Likewise, financial market frictions and the availability of future contracts, as well as institutional and regulatory constraints in financial markets, are also major factors in stock and oil prices, and may influence the movement toward long-run equilibrium. For example, the existence of adjustment costs may prevent economic agents from adjusting continuously. It is useful for investors and managers to explore the asymmetric behaviors of these economic and finance variables. Ignoring the asymmetric adjustment among variables may lead to biased inference and misleading conclusions. Asymmetric adjustment assumes that the adjustment toward equilibrium varies with internal or external characteristics of the system. Specifically, the convergence toward long-run equilibrium may be faster under positive deviations than under negative ones, or vice versa. Therefore, conventional Granger causality and cointegration approaches have been criticized on the ground that they neglect the potential asymmetric adjustment. Some research has considered the asymmetry in the oil price-macroeconomy relationship (Hamilton, 1996, 2003; Lee et al., 1995; Mork, 1989). However, except for Chiou and Lee (2009), the existing analysis does not recognize the asymmetric adjustment involving threshold cointegration tests.[1] Therefore, this study combines traditional cointegrating and threshold cointegrating analysis to examine the links between crude oil shocks and stock markets.

In the majority of the prior empirical works, the linkage exploration between crude oil and stock markets generally uses time series for single countries. More recently, there have been some studies that deal with the interactions between energy and economic activities based on panel data. To the best of our knowledge, no work has analyzed the oil–stock nexus by applying a panel approach at the international level. This paper aims to provide the first empirical study that tests a panel of 14 OECD and non-OECD countries in terms of long-run equilibrium and the Granger-causal relationship between crude oil shocks and stock markets. Furthermore, using threshold cointegration in panel framework, we are able to better estimate the asymmetric adjustment.

This study makes several unique contributions. First, the estimates adopting a large group of panel data are more robust than those based on time series models because the former account for the long run relationship and Granger-causal relationship between crude oil shocks and stock markets with more reliable identification. Second, we consider a multivariate framework with additional regressors. Except for crude oil prices and stock market prices, we also incorporate interest rates and industrial production. Following Fama's (1981) hypothesis, measures of economic activity and inflation have played a role in the analysis of stock market activity. It is thus important to consider interest rates and industrial production in the context of international economies. The effect of crude oil prices on stock markets is better captured by considering these necessary variables.[2] Thirdly, we take into account the possibility that the long-run relationship between crude oil shocks and stock markets may involve threshold effects. Thus, the panel threshold cointegration tests are conducted to examine the possible asymmetric effects on the oil–stock nexus, which accommodate asymmetric adjustment in the long run.[3]

The rest of this paper proceeds as follows. Section 2 introduces the econometric methodology employed in this study. Section 3 presents the data and preliminary investigation. Section 4 reports the empirical results, and the final section presents our conclusions.

2. Econometric methodology

To evaluate the potential linkages between crude oil prices and stock market prices, we employ panel cointegration techniques. Unlike the existing studies in the single nonstationary time-series literature, recently developed panel methods have produced new strands of panel cointegrating regression analysis. In a panel context, the number of observations available is greatly increased when testing the long-run relationship, and as a result, more informative data can be obtained. Thus, the panel-based tests can gain statistical power substantially and overcome the low power problem of asymmetric adjustment in the tests' univariate counterparts.

In this paper, the test for long-run equilibrium between crude oil prices and stock markets in a panel framework is conducted based on panel cointegration tests. Before proceeding to the panel-based cointegration tests, the panel-based unit root tests are performed. Panel-based cointegration tests are then conducted to examine the long-run relationship between the variables in question. In addition, it is possible that there are threshold effects in a potential cointegrating relationship between crude oil prices and stock markets, which indicates asymmetric adjustment in the oil–stock nexus. Thus, panel threshold cointegration tests are used to detect the asymmetry. Given that the variables are cointegrated, a panel VECM can be adopted to check whether a linear combination of nonstationary variables exists, which then suggests that a long-run equilibrium relation holds between the variables. However, if the variables are threshold

[1] See Balke and Fomby (1997), Enders and Granger (1998), Enders and Siklos (2001), and Pippenger and Goering (1993).

[2] These particular covariates should be included in levels, and we choose the real economic variables but not the nominal ones.

[3] Since our interest in this study is merely discovering whether there exists asymmetric cointegration relationship, the identification of the number of cointegration relationships is beyond the scope of this study.

cointegrated, the alternative panel threshold VECM should be estimated to perform nonlinear Granger-causality tests.

2.1. Panel unit root tests

There are a number of panel unit root tests that aim to solve the severely undersized and low power problems in single time series unit root tests. Important developments in the panel unit root tests have been made by Bai and Ng (2004), Breitung (2000), Choi (2001), Im et al. (2003), Levin et al. (2002), Maddala and Wu (1999), Moon and Perron (2004), and Pesaran (2007), among others. Consider the following basic autoregressive specification:

$$y_{it} = a_i + \rho_i y_{i\,t-1} + \mu_{it}, \tag{1}$$

where $i=1, \ldots, N$ denotes the country in the panel and $t=1, \ldots, T$ represents the time period, a_i is the effect of country i, ρ_i is the autoregressive coefficient and μ_{it} is the stationary error term. If $\rho_i<1$, y_{it} is weakly trend stationary, otherwise, if $\rho_i=1$, y_{it} contains a unit root. Based on the assumptions of cross-sectional dependence, these panel unit root tests can be grouped into two categories. Under the assumption of cross-sectional independence, Choi (2001) and Maddala and Wu (1999) propose the Fisher-type tests for heterogeneous panels, and Im et al. (2003) further take heterogeneity into account. We thus call the latter test the IPS test. In the second class of tests, Bai and Ng (2004), Moon and Perron (2004) and Pesaran (2007) allow for cross-sectional dependence.

Choi (2001) and Maddala and Wu (1999) propose comparable unit root tests to be conducted by using the non-parametric Fisher statistics. The panel test statistics of Choi (2001) and Maddala and Wu (1999, hereafter MW) are

$$P^{MW} = -2\sum_{i=1}^{N} \log(\mathrm{pv}_i), \tag{2}$$

$$Z^{C} = N^{-1/2}\sum_{i=1}^{N} \Phi^{-1}(\mathrm{pv}_i), \tag{3}$$

where $\Phi^{-1}(\cdot)$ is the inverse of the standard normal cumulative distribution function, and pv_i is the p-value. To avoid potential biases caused by improper specifications, the IPS test is then utilized to consider heterogeneous autoregressive coefficients. Moreover, the IPS test suggests an average of the ADF unit root tests when μ_{it} is serially correlated with different orders, i.e., the model given in $\mu_{it}=\sum_{j=1}^{p_i}\theta_{ij}\mu_{i\,t-j}+e_{it}$. Therefore, the IPS test specifies a separate ADF regression for cross-sectional units as expressed in Eq. (4):

$$y_{it} = a_i + \rho_i y_{it-1} + \sum_{j=1}^{p_i}\theta_{ij}\,\mu_{it-j} + e_{it}, \tag{4}$$

where p_i is the lag order in the ADF regression and is allowed to vary across cross-sections as well as ρ_i. The null hypothesis is that each series in the panel contains a unit root, that is, $H_0:\rho_i=1$ for all i, and the alternative hypothesis allows for at least one of the individual series in the panel to be stationary, i.e., $H_1:\rho_i<1$ for a specific i. The IPS t-bar statistic is specified by averaging the individual ADF statistics:

$$\bar{t} = N^{-1}\sum_{i=1}^{N} t_{\rho_i}, \tag{5}$$

where t_{ρ_i} is the individual t-statistic for testing $H_0:\rho_i=1$ for all i in Eq. (4). In the case where p_i may be nonzero for some cross-sections, the asymptotic distribution of a properly standardized version of the $\bar{t}$ statistic converges at a standard normal distribution $N(0, 1)$ by the limit theory.

Considering potential cross-sectional correlation, the Bai and Ng (2004) and Pesaran (2007) tests have been adopted in our empirical analysis for a combined check to obtain sensible integration results.

2.2. Panel threshold cointegration tests

If the variables to be integrated are first-order, they may form a stationary linear combination. More specifically, let SP_{it}, R_{it}, IP_{it} and OIL_{it} represent the natural logarithm of the real stock price series, short-term interest rates, industrial production and the real crude oil prices in country i and time t, respectively. We consider the following long-run model, assuming that all variables are unit root processes:

$$SP_{it} = \beta_0 + \beta_1 R_{it} + \beta_2 IP_{it} + \beta_3 OIL_{it} + u_{it}, \tag{6}$$

where $i=1, \ldots, N$ and $t=1, \ldots, T$, and u_{it} is the stochastic disturbance term. To test the null hypothesis of no cointegration, the unit root test is conducted on the long-run residuals $\hat{u}_{it}$ of Eq. (6). However, in the presence of asymmetric adjustment to the long-run equilibrium, that is, if the adjustment depends on the sign of the shocks, the test for traditional cointegration is misspecified (see Balke and Fomby, 1997; Enders and Siklos, 2001). Threshold cointegration proposed by Balke and Fomby (1997) and Enders and Granger (1998), seeks to address this problem. These authors replace the standard ADF auxiliary regression with the threshold autoregressive (TAR) process. Therefore, threshold cointegration can characterize the asymmetric dynamic adjustment process where the movement toward equilibrium under positive deviations may be different from that under negative deviations.

In this paper, the residuals for the unit root are examined by using the Caner and Hansen (2001) test to investigate the possible threshold cointegration relationships between crude oil shocks and stock markets. By aggregating the individual results with Fisher-type tests of Choi (2001) and Maddala and Wu (1999), the univariate TAR unit root methodology of Caner and Hansen (2001) is extended to a panel data context. Unlike former studies, Caner and Hansen (2001) develop several Wald tests to discriminate nonstationarity from nonlinearity. The model is given by a two-regime TAR model of the form:

$$\Delta y_t = \varphi_1' x_{t-1} 1_{\{Z_{t-1}<\lambda\}} + \varphi_2' x_{t-1} 1_{\{Z_{t-1}\geq\lambda\}} + v_t, \tag{7}$$

where $t=1, \ldots, T$, $x_{t-1}=(y_{t-1}\quad c_t{'}\Delta y_{t-1}\cdots\Delta y_{t-k})'$, $1_{\{\cdot\}}$ is the indicator function, v_t is an iid error, c_t is a vector of deterministic components including an intercept and a possible linear time trend, Z_{t-1} is the threshold variable, $Z_{t-1}=y_{t-1}$ for TAR specification and $Z_{t-1}=\Delta y_{t-1}$ for a modified version of TAR — "momentum" TAR (or MTAR) model. In the interval $\lambda\in[\lambda_1, \lambda_2]$, the threshold parameter λ is estimated by

$$\hat{\lambda} = \mathit{arg\ min}\ \hat{\sigma}^2(\lambda), \tag{8}$$

where $\hat{\sigma}^2(\lambda)$ is the residual variance, $\hat{\sigma}^2(\lambda)=T^{-1}\sum_{i=1}^{N}\hat{v}_t^2(\lambda)$, $P(Z_{t-1}\leq\lambda_1)>0$ and $P(Z_{t-1}\leq\lambda_2)<1$. To test for a threshold effect, the Wald statistic $W_T = T\left(\hat{\sigma}_0^2/\hat{\sigma}^2-1\right)$ is employed to test the hypothesis $H_0:\varphi_1=\varphi_2$, where $\hat{\sigma}_0^2$ is defined as the residual variance under the null. Because the distribution of the Wald statistic is non-standard, the asymptotic distribution and model-based bootstrap approach is conducted. The corresponding bootstrap critical values and p-values are computed through 10,000 bootstrap simulations. Caner and Hansen (2001) use the one-sided Wald statistic $R_{1T} = t_1^2 1_{\{\hat{\varphi}_1<0\}} + t_2^2 1_{\{\hat{\varphi}_2<0\}}$ and two-sided Wald statistic $R_{2T}=t_1^2+t_2^2$ to test for complete unit roots (i.e. $H_0:\varphi_1=\varphi_2=0$) against the first alternative stationary case (i.e. H_1:"$\varphi_1<0$ and $\varphi_2<0$") and the second alternative partial unit roots, namely, H_2:"$\varphi_1<0$ and $\varphi_2=0$" or "$\varphi_1=0$ and $\varphi_2<0$", where ϕ_i

denotes the first element of φ_i. Further, to distinguish the stationary case H_1 from the partial unit root case H_2, the negatives of individual t statistics t_1 and t_2 are examined. Exact p-values for these tests can be obtained using the bootstrap method. If we let pv_i be the exact p-value of a given Wald test for ith country, the panel data version of the Caner and Hansen (2001) tests can be constructed by considering the MW and Choi (2001) Fisher-type statistics as:

$$P^{CHMW} = -2\sum_{i=1}^{N} \log(pv_i), \tag{9}$$

$$Z^{CHC} = \left(1/\sqrt{N}\right)\sum_{i=1}^{N} \Phi^{-1}(pv_i). \tag{10}$$

As proposed by Fisher (1932) and Maddala and Wu (1999), if the statistics are continuous, then the corresponding significance level p-values are uniform (0,1) variables. Consequently, under the assumption of cross-sectional independence, pv_i are independent uniform (0,1) variables, and the Fisher statistic in Eq. (9) has a chi-squared distribution with 2N degrees of freedom when T tends to infinity and N is fixed; the statistic in Eq. (10) converges at standard normal distribution. To allow for the possible cross-sectional correlation in the panel data, the bootstrapping procedure is adopted to approximate the bootstrap distribution of the statistics.[4] Here, the null hypothesis suggests that the series is nonstationary in all countries, while the alternative implies that the series is stationary for some of the countries.

2.3. Panel Granger causality

Panel cointegration methods test the existence or absence of a long-run relationship between crude oil and stock markets. Given that the variables are cointegrated, a panel VECM is used to account for the Granger causality tests. Following the Engle and Granger (1987) two-step procedure, we first estimate the coefficients of the long-run model for Eq. (6) using the OLS procedure to obtain the estimated residuals $\hat{u}_{it}$. Defining the residuals from Eq. (6) as the error correction term, we then incorporate the lagged residuals into a panel VECM. The final dynamic error correction model can be written in the following form:

$$\Delta SP_{it} = \alpha_{1j} + \sum_{k=1}^{q}\gamma_{11ik}\Delta SP_{it-k} + \sum_{k=1}^{q}\gamma_{12ik}\Delta R_{it-k} + \sum_{k=1}^{q}\gamma_{13ik}\Delta IP_{it-k} + \sum_{k=1}^{q}\gamma_{14ik}\Delta OIL_{it-k} + \xi_{1i}\hat{u}_{it-1} + \varepsilon_{1it} \tag{11.a}$$

$$\Delta R_{it} = \alpha_{2j} + \sum_{k=1}^{q}\gamma_{21ik}\Delta SP_{it-k} + \sum_{k=1}^{q}\gamma_{22ik}\Delta R_{it-k} + \sum_{k=1}^{q}\gamma_{23ik}\Delta IP_{it-k} + \sum_{k=1}^{q}\gamma_{24ik}\Delta OIL_{it-k} + \xi_{2i}\hat{u}_{it-1} + \varepsilon_{2it} \tag{11.b}$$

$$\Delta IP_{it} = \alpha_{3j} + \sum_{k=1}^{q}\gamma_{31ik}\Delta SP_{it-k} + \sum_{k=1}^{q}\gamma_{32ik}\Delta R_{it-k} + \sum_{k=1}^{q}\gamma_{33ik}\Delta IP_{it-k} + \sum_{k=1}^{q}\gamma_{34ik}\Delta OIL_{it-k} + \xi_{3i}\hat{u}_{it-1} + \varepsilon_{3it} \tag{11.c}$$

$$\Delta OIL_{it} = \alpha_{4j} + \sum_{k=1}^{q}\gamma_{41ik}\Delta SP_{it-k} + \sum_{k=1}^{q}\gamma_{42ik}\Delta R_{it-k} + \sum_{k=1}^{q}\gamma_{43ik}\Delta IP_{it-k} + \sum_{k=1}^{q}\gamma_{44ik}\Delta OIL_{it-k} + \xi_{4i}\hat{u}_{it-1} + \varepsilon_{4it} \tag{11.d}$$

where γ are the short-run adjustment coefficients, ε are disturbance terms, assumed to be serially uncorrelated, and the optimal lag length q can be determined by the Akaike Information Criterion (AIC) or Schwarz Bayesian Information Criterion (SBC). In the above equations, we test $H_0: \gamma = 0$ for Granger causality tests. For example, the short-run Granger causality from interest rates, industrial production and crude oil prices to stock market prices is tested on the basis of $H_0: \gamma_{1sik} = 0\ \forall\ ik$, for $s = 2, \ldots, 4$, respectively. The short-run Granger causality indicates that the dependent variable responds only to short-term shocks to the stochastic environment. Specifically, shocks in interest rates and industrial production allow us to control demand changes that influence stock prices but are not captured by short-run crude oil price changes. Meanwhile, the significance exploration of the adjustment speed ξ examines the existence of long-run Granger causality. The long-run Granger causality determines the long-run equilibrium in cointegrating regression. It would thus be reasonable to expect that movements along this path could be permanent.

However, if the variables are asymmetric threshold cointegrated, a threshold vector error correction model (TVECM) should be employed to investigate the asymmetric adjustment toward the long-run equilibrium. With other conditions being equal, the two-regime threshold error correction model can be obtained by replacing $\hat{u}_{it-1}$ in VECM with $\xi_{1si}\hat{u}_{it-1}1_{\{\hat{u}_{it-1}\geq\tau\}} + \xi_{2si}\hat{u}_{it-1}1_{\{\hat{u}_{it-1}<\tau\}}$ or $\xi_{1si}\hat{u}_{it-1}1_{\{\Delta\hat{u}_{it-1}\geq\tau\}} + \xi_{2si}\hat{u}_{it-1}1_{\{\Delta\hat{u}_{it-1}<\tau\}}$, $\forall i, t$ and $s = 1, \ldots, 4$. This indicates the TAR and MTAR vector error correction representation, respectively. Here, the threshold τ could be set endogenously using the Chan (1993) methodology, and in many cases, we usually set it to zero. More specifically, the MTAR vector error correction model allows the error correction term to exhibit more momentum in one direction than in the other. Thus, the movement toward long-run equilibrium under positive deviations can differ from that under negative ones.

3. Data

In this study, we collect real stock prices of 14 OECD and non-OECD countries, along with the closing prices of the US price of WTI. The monthly data covers the period from January 1995 through December 2009. The real worldwide crude oil price is the nominal oil price deflated by the US PPI, whereas the real stock price is the stock price deflated by the respective CPI. Data on CPI and crude oil prices come from the International Financial Statistics (IFS) database and the Energy Information Administration (EIA), and stock prices come from OECD database. Data from additional macroeconomic regressors, such as industrial production and short-term interest rate, are obtained from the OECD database for each country. Because the interest rates are expressed as percentages, we define the natural logarithm of the interest rate as $\log(1 + r/100)$. The collected OECD and non-OECD countries include the United States, United Kingdom, Mexico,

Table 1
Panel unit root tests.

Variable	BN Z-statistic		BN P-statistic		Pesaran CIPS-statistic	
	Levels	Differences	Levels	Differences	Levels	Difference
The whole panel						
IP	−2.9739	3.2686***	5.7455	52.46***	−1.7196	−7.1646***
R	0.0927	4.1633***	28.694	59.155***	−2.6271	−6.1384***
SP	−1.7201	5.9786***	15.129	72.74***	−1.986	−9.7949***
WTI	0.0726	7.6804***	28.543	85.475***	−2.4413	−10.027***
The whole panel except Norway						
IP	−2.4379	3.203***	8.4198	49.097***	−1.6485	−6.9818***
R	−0.8015	7.942***	20.221	83.271***	−2.5088	−6.2979***
SP	−0.3255	5.7662***	23.653	67.581***	−1.882	−9.6024***
WTI	0.3918	8.6021***	28.825	88.031***	−2.1132	−10.295***

*** Significant at 1%.

[4] We are grateful to the reviewer who pointed out that the Fisher-type statistic mentioned above is not robust to cross-sectional correlation. Correspondingly, instead of implementing the proposed Fisher-type statistic directly, we employ the bootstrap procedure to allow for cross-sectional correlation.

Table 2
Panel threshold cointegration tests for the panel from 1995/01-2009/12.

	TAR model		MTAR model	
	MW panel test	Choi panel test	MW panel test	Choi panel test
Test for threshold effect				
W-statistic (under $H_0$of a unit root)	35.4604(0.1590)	−4.0419(0.0000)***	66.0662(0.0000)***	−9.1962(0.0000)***
W-statistic (under $H_1$of stationary)	36.2611(0.1376)	−4.2077(0.0000)***	68.8172(0.0000)***	−9.5773(0.0000)***
Test for cointegration				
R_{1T} (unidentified threshold)	30.0893(0.3608)	−2.8561(0.0021)***	76.7491(0.0000)***	−10.6262(0.0000)***
R_{1T} (identified threshold)	28.5658(0.4351)	−2.4995(0.0062)***	82.0435(0.0000)***	−11.2898(0.0000)***
t_1-statistic	82.5679(0.0000)***	−11.3541(0.0000)***	95.0579(0.0000)***	−12.8177(0.0009)***
t_2-statistic	33.80842(0.2075)	−3.6918(0.0001)***	44.7643(0.0244)**	−5.8377(0.0000)***

Note: The values in parentheses are the bootstrapped p-values.
*** Significant at 1%.
** Significant at 5%.

Norway, Sweden, Poland and Turkey, Brazil, India, Chile, China, Israel, Slovenia and South Africa.

4. Empirical results

Table 1 provides the results derived from panel unit roots in level and first difference for the order of panel integration. For each variable in these OECD and non-OECD countries, the results fail to reject the null hypothesis of unit roots in the levels, but they reject the null hypothesis in first differences at the 1% significant level. Thus, we conclude that crude oil prices, real stock prices, interest rates and industrial production of the sample are $I(1)$. We next consider the panel cointegration tests.

To determine the existence or absence of an asymmetric long-run relationship between real stock prices and crude oil prices, we apply panel threshold cointegration tests for all considered variables. Testing for stationarity of the error correction term (the estimated residual of Eq. (6)), with the consideration of possible thresholds by simultaneity, is the central focus in our analysis. We implement the TAR model and the MTAR model to examine potential asymmetry. Given that the null hypothesis of no threshold effects could not be rejected, it may be expected that our analysis above could be applicable and the error correction term is also stationary. Table 2 presents the results from panel threshold cointegration for the whole panel. Wald tests for the significance of threshold effects are shown in the first two rows and the significance tests of stationarity are listed in the last four rows. Further, as the only oil exporting country in the whole panel, Norway might affect the final results. We therefore test for threshold cointegration for these OECD and non-OECD countries excluding Norway, as shown in Table 3. The results of Table 3 also indicate the existence of threshold cointegration, which suggests that the inclusion of Norway does not affect the asymmetric cointegration results. Meanwhile, based on the assumption of asymmetric cointegration, we compare the analysis results from the TAR threshold cointegration model to the MTAR model, and find that the analysis based on the MTAR model is more robust across time[5] than the one based on the TAR model. Therefore, the MTAR specification is adopted to explain the asymmetric adjustment mechanism. The MTAR model for the panel rejects the null hypothesis of no threshold effect and no cointegration at the 1% level. Therefore, a threshold cointegrating relationship exists between crude oil prices and stock markets for these countries. The results indicate that the adjustment toward equilibrium is asymmetric. Furthermore, in Table 2, the Wald R_{1T} statistics in rows 3–4 reject the null hypothesis of a unit root, and the t statistics in the last two rows are both statistically significant at the 1% level, indicating that the error correction term is stationary. Therefore, crude oil prices and stock prices in these OECD and non-OECD countries are cointegrated when allowing for asymmetric adjustment. Given these findings, we can determine that asymmetric threshold cointegrated relations exist for the oil-stock nexus in these countries. The presence of threshold cointegration reveals that there is asymmetry in the transmission of oil price changes to stock markets and vice versa. Thus, asymmetry does exist in the adjustments, whereas positive or negative deviations from the long-run equilibrium will not be corrected in the same manner. The results also indicate, from the perspective of investments, that crude oil and these OECD and non-OECD stock markets can be considered integrated rather than segmented. This may imply that if an investor diversifies his or her portfolio by holding assets in both the oil and stock markets, the resulting portfolio does not improve significantly with respect to the original in terms of reducing market risk or increasing long-term benefits. In other words, investing in oil and these OECD and non-OECD stock markets would not generate any long-run gain in portfolio diversification.

The existence of cointegration between stock prices, crude oil prices, industrial production and interest rates suggests the presence of a long-run relationship. Table 4 displays the results for the whole panel and Table 5 reports the results for the panel except Norway. All coefficients are statistically significant at the 1% level and the coefficients can be interpreted as elasticities. The results indicate that the inclusion of Norway has almost no effect on the long-run coefficients. We find that increased crude oil prices have a positive impact on stock prices in the long run and vice versa, which confirms the view of Narayan and Narayan (2010). Industrial production also has a statistically significant positive effect on crude oil prices, whereas interest rates have a negative effect on crude oil prices. More precisely, a 1% increase in oil prices results in around a 0.43% rise in stock prices, a 1% increase in stock prices results in around a 0.54% increase in oil prices, a 1% increase in industrial production causes oil prices to fall by around 0.61%, and a 1% increase in interest rates leads to a decrease of around 0.5% in oil prices. To test the robustness of the results, we also explore the long-run relationship between stock prices and crude oil prices from 1995/01 to 2003/12 and from 1998/01 to 2007/12. These results suggest that the long-run relationship between stock prices and crude oil prices is robust for the choice of the sample period, although the magnitudes of the effects suffer some variations from time to time.[6]

[5] To obtain robust results, we also test for threshold cointegration from 1995/01 to 2003/12 and from 1998/01 to 2007/12. The asymmetric cointegration results are found to be robust to the choice of the sample period. The detailed results are not reported in the paper due to space limitations, but they are available from the authors upon request.

[6] We thank the reviewer who provided us with this comment. The detailed results are not reported in the paper due to space limitations, but they are available from the authors upon request.

Table 3
Panel threshold cointegration tests for the panel, with the exception of Norway, 1995/01-2009/12.

	TAR model		MTAR model	
	MW panel test	Choi panel test	MW panel test	Choi panel test
Test for threshold effect				
W-statistic (under H_0 of a unit root)	26.5745(0.4325)	−2.3328(0.0098)***	52.8840(0.0014)***	−7.2967(0.0000)***
W-statistic (under H_1 of stationary)	28.6864(0.3260)	−2.8278(0.0023)***	50.9588(0.0025)***	−6.9969(0.0000)***
Test for cointegration				
R_{1T} (unidentified threshold)	29.3374(0.2961)	−2.9754(0.0015)***	63.1105(0.0000)***	−8.7846(0.0000)***
R_{1T} (identified threshold)	27.9032(0.3643)	−2.6471(0.0041)***	61.6055(0.0001)***	−8.5755(0.0000)***
t_1-statistic	65.6182(0.0000)***	−9.1261(0.0000)***	85.9089(0.0000)***	−11.6348(0.0000)***
t_2-statistic	33.5239(0.1476)	−3.8769(0.0000)***	31.5874(0.2116)	−3.4697(0.0003)***

Note: The values in parentheses are the bootstrapped p-values.
*** Significant at 1%.

Because asymmetric threshold cointegration is evident, the potential nonlinear Granger-causality for the whole panel is further explored by two-regime TVECM. Here, the parameters ξ represent the speeds of long-run adjustment. The asymmetric Granger-causality relations are reported in Table 6. The results for the OECD and non-OECD countries, excluding Norway, which are similar to the results in Table 6, are represented in Table 7. This indicates that the inclusion of Norway does not affect the results. Specifically, the results in Table 6 reflect that there is indeed a long-run relationship running from oil price shocks, interest rates, and industrial production to stock markets and from stock markets, interest rates, and industrial production to oil price shocks in regimes 1 and 2. There is also a long-run Granger causality relation between oil price shocks, interest rates, stock markets and industrial production in both regimes. However, the speeds of adjustment toward equilibrium differ from regime 1 to regime 2, indicating the presence of asymmetric adjustment. For the stock and crude oil prices, the speed of adjustment toward equilibrium is faster under negative changes of the deviation than under positive ones. That is, the speed of adjustment for firms below their targets is found to be faster than that from above, demonstrating the long-run asymmetry of the expected sign. In other words, these results show strong evidence of asymmetry in the transmission of oil price shocks to stock prices and vice versa, possibly suggesting that the speed of the stock–oil transmission becomes faster when the deviation is narrowing. With respect to industrial production, the error correction terms are statistically significant at the 1% level, also suggesting the existence of a long-run relationship between industrial production and oil prices. However, the speed of adjustment to equilibrium is relatively slow when the change of the deviation is negative. In particular, there is no statistically significant long-run relationship between interest rates and oil prices. In addition, it appears that in the short run, there is statistically significant Granger-causality from crude oil prices to stock prices and vice versa when the change of the deviation is positive. However, there is only statistically significant short-run Granger causality from stock prices to crude oil prices but not from crude oil prices to stock prices when the change of the deviation is negative. Thus, the short-run relationship between industrial production and oil prices is present under a negative change of the deviation but not under a positive change.

Table 4
Panel long-run estimates for the panel from 1995/01-2009/12.

Dependent variable	Variable	Coefficient	t-statistic
SP	R	−0.3955***	−4.5502
	IP	0.4494***	12.5444
	WTI	0.4326***	26.8315
WTI	R	−0.5025***	−5.1581
	IP	0.6148***	15.5466
	SP	0.5449***	26.8298

*** Significant at 1%.

Table 5
Panel long-run estimates for the panel, with the exception of Norway, 1995/01-2009/12.

Dependent variable	Variable	Coefficient	t-statistic
SP	R	−0.4082***	−4.7487
	IP	0.5270***	14.6111
	WTI	0.3757***	22.3114
WTI	R	−0.5219***	−5.2716
	IP	0.6637***	16.1093
	SP	0.4996***	22.3123

*** Significant at 1%.

5. Concluding remarks

There is a growing body of literature about the Granger-causal relationship between crude oil and stock markets. The bulk of this literature has adopted time series data. The aim of this paper is to investigate the presence of a Granger-causality relationship between crude oil shocks and stock markets, as well as the existence of asymmetric adjustment toward the long-run equilibrium, in a panel context.

When a group of OECD and non-OECD countries is considered, possible differences in the Granger-causality direction are detected. Panel threshold cointegration tests confirm that there is some evidence for asymmetric threshold effects between crude oil prices, interest rates, industrial production and stock markets of these countries.

There is clear evidence for the existence of a bidirectional long run Granger-causal relation between crude oil and stock markets for the 14 OECD and non-OECD countries from January 1995 to December 2009. Meanwhile, the results of the analysis also indicate that increased crude oil prices have a positive impact on stock prices and that increased stock prices influence crude oil prices positively in the long run. It is not surprising that increased stock prices act as a stimulus to higher crude oil prices. However, the fact that crude oil prices influence stock prices positively is inconsistent with theoretical expectation. This may reflect the fact that oil stocks are contained in the stock prices. More specifically, this positive relationship may be caused by the rise in leveraged investment in stock and may also suggest that the impact of industrial production or other internal and domestic factors on these stock markets were more dominant than the increase in crude oil prices. Furthermore, there is asymmetric

Table 6
Panel TVECM Granger-causality test for the panel, 1995/01-2009/12.

Dependent variable	Regime 1					Regime 2				
	Short run				Long run	Short run				Long run
	ΔSP	*ΔR*	*ΔIP*	*ΔOIL*	*ECM*	*ΔSP*	*ΔR*	*ΔIP*	*ΔOIL*	*ECM*
ΔSP	–	0.1287	−0.2304***	−0.1003**	−0.5421***	–	−0.2491	−0.1674	−0.0488	−0.6364***
		(0.1811)	(0.0737)	(0.0429)	(0.0561)		(0.2151)	(0.1098)	(0.0584)	(0.0638)
ΔR	−0.0735***	–	0.0819***	0.0372***	−0.0167	0.0133	–	−0.0199	−0.0408**	0.0005
	(0.0270)		(0.0255)	(0.0137)	(0.0198)	(0.0257)		(0.0346)	(0.0178)	(0.0201)
ΔIP	−0.1028*	−0.1679	–	0.0293	0.1239***	−0.0979**	−0.2810**	–	0.0712**	0.0875**
	(0.0619)	(0.1446)		(0.0316)	(0.0454)	(0.0484)	(0.1267)		(0.0333)	(0.0377)
ΔOIL	−0.4201***	−0.4739**	0.0574	–	0.4528***	−0.2653***	0.1304	0.2546**	–	0.5489***
	(0.1058)	(0.2410)	(0.0967)		(0.0744)	(0.1012)	(0.2516)	(0.1282)		(0.0750)

Note: The values in parentheses are standard errors.
*** Significant at 1%.
** Significant at 5%.
* Significant at 10%.

adjustment toward long-run equilibrium. The speeds of adjustment to the long-run equilibrium are faster under negative changes of the deviation than that under positive changes. This behavior may suggest that the speed of the stock–oil transmission becomes faster when the deviation is narrowing. Moreover, there is also bidirectional long-run Granger causality between industrial production and crude oil prices. In the short run, the Granger-causal relationship between crude oil and stock markets is bidirectional under positive changes of the deviation, but it is unidirectional under negative changes. However, the short-run Granger causality between industrial production and crude oil prices is bidirectional only when the change of the deviation from the long-run equilibrium is negative. From these panel results, there is long-run Granger causality between crude oil shocks and stock markets for the OECD and non-OECD panel, and it is concluded that the asymmetric dynamic adjustment behavior is indeed evident for these countries. Further research should be directed to investigate the country-specific mechanisms by which crude oil prices and disaggregated energy prices affect economic growth and stock prices.

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Table 7
Panel TVECM Granger-causality test for the panel, with the exception of Norway, 1995/01-2009/12.

Dependent variable	Regime 1					Regime 2				
	Short run				Long run	Short run				Long run
	ΔSP	*ΔR*	*ΔIP*	*ΔOIL*	*ECM*	*ΔSP*	*ΔR*	*ΔIP*	*ΔOIL*	*ECM*
ΔSP	–	0.1430	−0.3072***	−0.0903**	−0.6077***	–	−0.2228	−0.1533	−0.0328	−0.6686***
		(0.1889)	(0.0859)	(0.0437)	(0.0613)		(0.2212)	(0.1166)	(0.0615)	(0.0672)
ΔR	−0.0838***	–	0.0413	0.0340**	−0.0187	0.0140	–	−0.0217	−0.0428**	0.0063
	(0.0288)		(0.0288)	(0.0139)	(0.0209)	(0.0279)		(0.0375)	(0.0191)	(0.0216)
ΔIP	−0.1371**	0.1504	–	0.0373	0.1212**	−0.0867	−0.2511*	–	0.0621*	0.0774*
	(0.0640)	(0.1420)		(0.0309)	(0.0476)	(0.0543)	(0.1376)		(0.0371)	(0.0419)
ΔOIL	−0.4408***	−0.4244*	0.0158	–	0.4219***	−0.2435**	0.1548	0.2661*	–	0.5104***
	(0.1130)	(0.2476)	(0.1097)		(0.0798)	(0.1087)	(0.2621)	(0.1376)		(0.0792)

Note: The values in parentheses are standard errors.
*** Significant at 1%.
** Significant at 5%.
* Significant at 10%.

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