ENERGY STUDIES REVIEW
SPECIAL ISSUE
2014

# ON THE LINK BETWEEN OIL AND COMMODITY PRICES: A PANEL VAR APPROACH

**VINCENT BRÉMOND**
University of Paris Ouest, France

**EMMANUEL HACHE**
IFP Energies Nouvelles, France

**MARC JOËTS**
University of Paris Ouest, France

**ABSTRACT**

The aim of this paper is to study the relations between the price of oil and a large dataset of commodity prices, relying on panel data settings. Using second generation panel cointegration tests, our findings show that the WTI and commodity prices are not linked in the long term. Nevertheless, considering our results in causality tests, we show that short-run relations exist, mainly from the price of crude oil to commodity prices. We thus implement a panel VAR estimation with an impulse response function analysis. Two main conclusions emerge: (i) fast co-movements are highlighted, while (ii) market efficiency is emphasised.

**KEYWORDS**
Oil prices, commodities, panel VAR, impulse response, cointegration, causality.

**Corresponding Author**
Marc Joëts: marcjoets@ipag.fr
Ipag Lab, Ipag Business School and Economix-Cnrs University of Paris Ouest, France.

# 1. INTRODUCTION

## CONTEXT

From the beginning of the 21st century, commodities markets have experienced major changes, with a steady and continuous upward trend until mid-2008, when prices collapsed in the wake of the financial crisis, before re-engaging in an upward movement from 2009 to 2011. This represents an apparent break from the pattern observed during the 1980s and the 1990s when prices fell around 1% per year on average. Although the timing and the magnitude have been quite different in the various different segments of the commodities markets (energy, non-ferrous metals, agricultural raw materials and beverages), the price increase that began in late 2001 had spread into all commodities markets by 2004-2005 in the context of steady world economic growth. Indeed, from 2004 to 2008, the world registered an average economic growth higher than 5%, the strongest number observed since the early 1970s. This strong world economic growth has increased demand for industrial and energy commodities. It was particularly fuelled by the Chinese economy, which became the world's leading consumer in the main non-ferrous metals and agro-industrial raw materials markets. Between 2001 and 2006, it accounted for all additional demand in the lead market, 80% of that in the cotton and zinc markets, and some 50% in the copper and aluminium markets. This Chinese boom had a particular impact, as it followed the 1990s, which were marked by a lack of investment and a strong downsizing trend in key industrial activities such as the non-ferrous metals sector.

Commodities prices have tripled, on average, between 2001 and 2012 with a first price boom from 2001 to 2008 (interrupted by the world financial crisis in the third quarter of 2008) and then followed by a price collapse from 2008 to 2009. Since then, commodities markets have experienced a new upward trend from 2009 to 2011 with a huge increase in price volatility.

In the history of commodities markets, three periods of sharp market price increases are generally identified (Radetzki, 2006). The first one followed the Great Depression of the 1930s: it reflected an upward adjustment in the commodities markets due to the sharp decline observed during 1929-1932 especially in food commodities. The second one, which was limited to a few quarters from 1949 to 1952, was the consequence of the Korean War and the need to build emergency storage capacity from the United States in critical raw materials such as rubber which was produced in Asian areas. During this period, the prices of raw materials[1] increased on average by 45%, with a surge in prices of agro-industrial products (+84%) and, to a lesser extent, of metals and minerals (+34%). The third one, during the 1970s, is clearly marked by external factors (first and second oil shocks) and severe weather conditions in certain production areas between 1971 and 1974 which also led to an increase in the prices of agricultural raw materials. Overall, during this period, energy prices quadrupled and the prices of raw materials tripled, on average.

Thus, during the 20th century, price volatility was more driven by strong movements in agricultural products as a result of severe weather conditions or geopolitical uncertainties, or by structural changes in industrial organisation (new pricing system in the light of the OPEC creation, dismantlement of cartels in non-ferrous metals industries, etc.).

[1]. Measured by the International Monetary Fund (IMF) index.

**TABLE 1: INDICES OF COMMODITY PRICES**

| | 2001 | 2003 | 2005 | 2007 | 2008 | 2009 | 2010 | 2011 | 2012 |
|---|---|---|---|---|---|---|---|---|---|
| **All primary commodity** | **58.3** | **65.0** | **100** | **135.1** | **172.3** | **120.7** | **152.2** | **192.2** | **186.2** |
| **Non-Fuel** | **75.8** | **81.8** | **100** | **140.5** | **151.1** | **127.3** | **160.9** | **189.5** | **170.9** |
| **Food** | **80.4** | **88.5** | **100** | **127.2** | **156.9** | **133.9** | **149.2** | **178.6** | **175.5** |
| **Beverages** | **65.6** | **85.5** | **100** | **123.3** | **152.0** | **154.4** | **176.2** | **205.5** | **167.4** |
| **Industrial** | **72.6** | **75.3** | **100** | **154.3** | **145.7** | **118.7** | **169.9** | **197.8** | **167.1** |
| **Agricultural raw materials** | **95.1** | **95.4** | **100** | **114.2** | **113.4** | **93.9** | **125.1** | **153.5** | **134.0** |
| **Metal** | **56.3** | **60.7** | **100** | **183.3** | **169.0** | **136.5** | **202.3** | **229.7** | **191.0** |
| **Energy** | **48** | **55.2** | **100** | **131.9** | **184.7** | **116.8** | **147.1** | **193.8** | **195.2** |

Notes: Source, International Monetary Fund

The 2001-2011 boom and bust appears quite different. Indeed commodities markets experienced major changes with a steady and continuous upward trend mid-2008 when prices collapsed in the wake of the financial crisis, before re-engaging in an upward movement from 2009 to 2011. This represents an apparent break from the one observed during the 1980s and the 1990s when prices fell around 1% per year on average. The price increase that began in 2001 was first largely driven by higher prices for non-ferrous metals and this trend spread to the other commodities markets, illustrating the so-called *co-movement phenomenon*. Moreover, it happened in a new economic and financial context. On the one hand, due to the energy crisis, the relation between non-energy commodity prices and crude oil prices has grown in importance in recent years, especially for agricultural and beverage commodities (maize, soybeans and sugar) which are used for the production of alternatives to crude oil. The food versus fuel debates are nowadays a key issue for policy makers and the relations between agricultural commodities and crude oil are being widely studied in the academic literature. On the other hand, some authors (Coleman and Levin, 2006; Masters, 2008; Masters and White, 2008a, 2008b) claim that the new commodity price dynamic is widely linked to the movement of "financialisation" of the commodities markets with the introduction of new financial tools such as the Exchange Traded Fund (ETF), or can be explained during certain periods of financial instability by speculative trading (Hache and Lantz, 2013). In this new context, the co-movement question seems to be highly relevant.

## 1.1 Literature Review

Since the founding paper of Pindyck and Rotemberg (1990), the relations between commodity prices through the co-movement analysis have been often studied. The first studies were dedicated to the excess co-movement phenomenom. Pindyck and Rotemberg brought to light the excess of co-movement for a set of seven commodities (namely, wheat, cotton, copper, gold, crude oil, lumber, and cocoa), defined as when a co-movement between unrelated commodities[2] prices remains "...in excess of anything that can be explained by the common effects...". They explain that raw commodities may have a common trend because of direct effects (an increase of industrial production leading to an increase in industrial commodities' demand for the production process and non-industrial commodities' demand due to the increase of revenues) or indirect effects (through the expectations for commodities, affecting the storage process and then the current prices) resulting from macroeconomic changes. They include in their model macroeconomic variables (nominal interest rate, industrial production, consumer price index, etc.) in order to take into account common effects. They estimate the commodity price equation using Ordinary Least Squares (OLS), where the explicative variables are "common" to all the commodities, and then study the correlation between the errors. If the

[2]. Pindyck and Rotemberg explained that the commodities are unrelated because "none … are substitutes or complements, none are co-produced, and none is used as a major input for the production of another".

correlations between the errors are different from 0, excess co-movement is detected, according to the authors. They conclude by explaining that excess co-movement can be due to: (i) herd behaviour from the traders; (ii) liquidity constraints; (iii) frequency of data. Nevertheless, few studies have supported similar conclusions.

Palaskas and Varangis (1991), considering a dataset of nine commodity prices (cocoa, coffee, wheat, cotton, rubber, copper, lead, crude oil, and silver), conclude that if the co-movement can be detected (via cointegration tests), an excess of co-movement, unlike the findings of Pindyck and Rotemberg, is extremely rare. Leybourne et al. (1994) compare these two studies, and, after splitting the concept of excess co-movement into two sub-definitions (namely "strong" and "weak" excess co-movement), estimate empirically their own model with twelve commodities (cocoa, coffee, copper, cotton, gold, lumber, aluminium, crude oil, soya, sugar, wheat), and finally conclude that excess co-movement does not occur frequently. Deb et al. (1996), using GARCH processes on two sample periods (1960-1985 and 1974-1992), and focusing on four possible definitions of the excess of co-movement, find that no excess co-movements are found in the first sample, and that in the second sample, the excess of co-movement is not found with tests of the "correct size and good power".

Using two models (a macro model and an equilibrium model that takes inventory into account), Ai et al. (2006) draw similar conclusions[3], supporting the results that the excess co-movement hypothesis is not validated[4]. They demonstrate that the price co-movements are explained by demand and supply side factors and also by fundamentals[5].

As the majority of studies have rejected the excess of co-movement hypothesis, recent studies have focused more and more on the co-movement relations. Actually, Cashin et al. (1999), with a concordance statistic[6], conclude that co-movement does not exist between two unrelated commodity prices, while it is common for two related commodity prices[7].

Saadi (2001) applies cointegration and causality tests to mining (aluminium, copper, tin, nickel, lead, zinc) and agricultural (cocoa, coffee, cotton, rubber) commodity prices and finds evidence of co-movement for the majority of pairs of prices over the study period, which runs from 1970:01 to 1998:12.

Yang (2004), using a Hansen and Johansen (1993) procedure, advances the fact that governmental policy matters, too. They find a long-run equilibrium relation between four agricultural US futures commodity prices (corn, oat, soybeans, and wheat) from mid-1996, while the Federal Agricultural Improvement and Reform Act was implemented the same year, leading to a more liberalised agricultural market[8].

Turning to the studies that specifically consider oil, Natanelov et al. (2011) estimate the co-movement relation between oil prices and a set of agricultural commodities (cocoa, coffee, corn, soybeans, soybean oil, wheat, rice, and sugar) and gold futures prices. Using a Johansen cointegration test for the period 2002-2010, the authors find that four commodity prices are cointegrated with crude oil (cocoa, wheat, coffee, and gold) with a causality link running from crude oil prices to cocoa and gold prices, and from wheat and coffee prices to crude oil prices. However, in five cases, the hypothesis of no cointegration is not rejected.

Campiche et al. (2007), using the same methodological tools (i.e., a Johansen cointegration test) and about the same set of data (agricultural commodity prices), find no cointegration evidence for the period 2003-2005, whereas they do find cointegration between crude oil prices and corn and soybean prices in 2006-2007.

Baffes (2007), with an OLS estimation, examines the pass-through relations between crude oil prices and (i) commodity indexes and (ii) individual commodity prices[9]. Regarding the results for the indexes, the highest elasticities with oil prices are obtained when fertiliser and beverage prices are considered (when raw material prices are almost inelastic). Individual elasticities for 35 commodity prices with oil price are also provided.

---

3. The authors consider five agricultural commodities: wheat, barley, corn, oats, and soybeans, from 1957:01 to 2002:09.

4. The equilibrium model gives better results than the macro models.

5. Moreover, the authors conclude that "there is] raising doubts on the role of speculation per se in causing the large price movements".

6. Measuring the time share with two series in the same state (i.e., both in increase or decrease).

7. Cashin et al. consider successively the datasets of Pindyck and Rotemberg (1990) for unrelated commodity prices and Deb et al. (1996) for related commodity prices.

8. Unlike the 1981-1996 period, when cointegration is not detected.

9. In three different periods.

Joëts and Mignon (2011) estimate long-run relations between energy forward prices (i.e. oil, gas, coal, and electricity) at distinct maturities and find that, using a non-linear panel cointegration methodology (i.e. PSTR models), oil prices are positively linked to gas and coal.

Nazlioglu and Soytas (2012) contribute to the debate on the relation between oil prices and agricultural commodity prices[10] by implementing a panel data cointegration test, as well as a Granger causality test. Considering 24 different agricultural prices, the authors validate the hypothesis of cointegration with the tests developed by Pedroni (2004). Their estimation of the long-run parameters is realised through the Fully Modified Ordinary Least Squares (FMOLS) and Dynamic Ordinary Least Squares (DOLS) procedure developed by Kao and Chiang (2000) and Mark and Sul (2003). Their estimations reveal that, except for cotton, coffee, and sugar, where the impact is null, an increase in oil prices results in an increase in agricultural commodity prices.

As far as we know, with the exception of Nazlioglu and Soytas(2012), there is no study of the co-movement between commodities with the use of panel data tools. The majority of studies have been made with bivariate time-series procedures. Moreover, while Nazlioglu and Soytas (2012) focused on agricultural commodities with first generation panel unit root test which have the restrictive assumption to not account for cross-section dependence, we rely on several kinds of commodities using second generation unit root and cointegration panel tests[11]. Our approach considers potential cross-section dependence between commodity prices. In this article, we will (i) use panel data methodology, reinforcing the robustness of our results, and (ii) focus on a very large dataset, including more than the agricultural commodity prices generally used. The rest of this article is organised as follows. Section 2 describes the dataset and econometric methodologies used. The empirical results are then provided in Section 3. Finally, Section 4 concludes the article.

## 2. DATA, UNIT ROOT, AND COINTEGRATION TESTS

We consider daily data over the period from July 11, 2000 to July 15, 2011. We rely on thirty commodity series: crude oil West Texas Intermediate (WTI), and 29 other commodity prices[12]. We split the commodity data into eight panel groups according to their characteristics: namely, Energy, Precious Metals, Non-Ferrous Metals, Agro-Industrials, Food, Oleaginous, Exotic, and Livestock. Using such a large sample of commodity prices allows us to take into account possible heterogeneity in the relations between the markets. All series are expressed in Log form and depicted in the Appendix. To control for the economic and financial environment that might impact all commodity price series, we rely on the Standard & Poor's 500 (SP500) equity index - which has the advantage of being available at a daily frequency. This variable also allows treating oil and commodity prices as financial assets and controls for the recent financial turmoil.

In order to see whether oil prices can impact the dynamic of commodity prices, we first investigate the unit root properties of our panel groups. We employ second-generation tests that are based on the assumption of cross-sectional dependence between the panel members[13]. Indeed, the application of the CD test developed by Pesaran (2004)-based on the average of the pairwise correlation coefficients of the OLS residuals from the individual regressions-shows that such cross-sectional correlations exist in our sample (see Table 5 in the Appendix). The Choi (2002) unit root test is used to study the stationarity of our series. It relies on an error-components panel model and removes the cross-sectional dependence by eliminating (i) individual effects using the Elliott, Rothenberg and Stock (1996) methodology (ERS), and (ii) the time trend effect by centering on the individual mean. As shown in Table 2 the Choi test concludes in favor of the unit root hypothesis: *i.e.*, that all price series, except the Livestock panel, are I (1) (*i.e.* stationary in first log-difference)[14].

---

10. The exchange rate of the USD is also considered.

11. Our approach considers potential cross-section dependence between commodity prices.

12. The data were extracted from DataStream.

13. According to Pesaran (2004), cross-sectional dependence can arise for several reasons, such as spatial spillovers, financial contagion, socioeconomic interactions, and common factors.

14. For WTI and SP500, the augmented Dickey-Fuller test has been used to check the unit root properties of each time series. The results reveal that both series are I (1).

### TABLE 2: SECOND-GENERATION PANEL UNIT ROOT TESTS

| | CHOI | | |
|---|---|---|---|
| | $P_m$ | $Z$ | $L*$ |
| **Energy** | **0.480 (0.310)** | **-0.150 (0.430)** | **-0.160 (0.430)** |
| **Precious metals** | **-0.210 (0.580)** | **0.590 (0.720)** | **0.790 (0.780)** |
| **Non-Ferrous metals** | **-0.470 (0.680)** | **-0.200 (0.410)** | **-0.180 (0.420)** |
| **Agro-Industrials** | **1.060 (0.140)** | **-1.240 (0.100)** | **-1.190 (0.110)** |
| **Food** | **0.700 (0.230)** | **1.160 (0.120)** | **-1.070 (0.140)** |
| **Oleaginous** | **0.910 (0.180)** | **-1.220 (0.100)** | **-1.110 (0.130)** |
| **Exotic** | **0.650 (0.250)** | **-.0750 (0.220)** | **-0720 ( 0.230)** |
| **Livestock** | **4.100 (0.000*)** | **-2.960 (0.000*)** | **-3.100 (0.000*)** |

Notes: Between parentheses: *p*-values. For Choi's test, the optimal lag orders in the individual ERS statistics (Elliott, Rothenberg, and Stock., 1996) for each series are determined by $\boldsymbol{p_{max}}$ 012; all tests are computed with individual effects and time trends specifications; under the unit root hypothesis, the Choi statistics are standard normal when $T$ and $N$ converge jointly to infinity.

Turning now to the cointegration analysis, we first rely on first-generation tests, namely the seven tests proposed by Pedroni (1999, 2004) which are all based on the null hypothesis of no cointegration. Among the seven Pedroni tests, four are based on the within dimension (panel cointegration tests) and three on the between dimension (group-mean panel cointegration tests). Group-mean panel cointegration statistics are more general in the sense that they allow for heterogeneous coefficients under the alternative hypothesis. We estimate, for each panel, two long-run relations: (i) a relation between all the commodity prices of the considered panel, and (ii) the same relation adding the SP500 as a proxy for the financial and economic environement[15].The results are not clear-cut, but the hypothesis of no cointegration seems to be privileged for almost all series (as expected for the Food and Non-Ferrous Metals panels), and so we do not have co-movements between WTI and commodity prices in the long-run. As with the unit root test, we also apply second-generation cointegration tests allowing for cross-sectional dependence. The four panel error correction based tests proposed by Westerlund (2007) rely on structural dynamics and are a panel extension of the Banerjee et al. (1998) tests developed in the time series context. Among Westerlund's four tests, two consider a homogeneous cointegrating relation under the alternative, while the two others allow for a heterogeneous long-term relation. The results shown in Table 3 show that all the panel groups are not cointegrated with the WTI, which confirms the results of the first generation tests. Considering the fact that our conclusions support the no cointegration hypothesis, the long term relation between WTI crude oil and the commodities panels is not validated. However, short run co-movements between prices may exist, resulting from investors' hedging strategies. To investigate this phenomenon, we test for cross-causality between markets using the new panel Granger causality test proposed by Dumitrescu and Hurlin (2012).

### TABLE 3: SECOND-GENERATION PANEL COINTEGRATION TESTS

| | WESTERLUND | | | |
|---|---|---|---|---|
| | GROUP-MEAN STATISTICS | | PANEL STATISTICS | |
| | $G_\tau$ | $G_\alpha$ | $P_\tau$ | $P_\alpha$ |
| **EnergyG** | 0.736 (0.730) | -0.536 (0.200) | -1.066 (0.380) | -1.507 (0.370) |
| **MetalpG** | 1.698 (0.080) | -2.361 (0.070) | -0.245 (0.600) | 0.052 (0.720) |
| **MetalnfG** | -0.732 (0.260) | -1.612 (0.110) | -0.080 (0.600) | -0.928 (0.320) |
| **AgroG** | -0.047 (0.493) | 0.064 (0.440) | -0.284 (0.403) | -0.437 (0.333) |
| **AlimG** | 0.095 (0.580) | 0.327 (0.593) | -0.409 (0.423) | -0.546 (0.360) |
| **OleaG** | 1.083 (0.830) | 0.869 (0.750) | 0.526 (0.750) | 0.182 (0.610) |
| **ExoG** | -1.514 (0.080) | -1.939 (0.057) | -1.463 (0.123) | -2.121 (0.077) |

[15]. The results are available upon request to the authors.

Notes: (a) Between parentheses: *p*-values with cross-sectional dependence based on bootstrapped distribution (300 bootsrtrap replications). (b) Tests are computed with individual effects and time trends. (c) The Bartlett kernel is used for the semi parametric corrections. (d) The lleads and lags in the error correction test are chosen using the Akaike criterion. (e) The number of common factors is determined by the $\boldsymbol{IC_1}$ criterion (See Bai and Ng, 2004) with five as the maximum number of factors.

Following the seminal work of Granger (1969) on time series causality, we say that a variable $x$ *causes* a variable $y$ if we are able to better predict $y$ using all available information than in the case where the information set used does not include $x$. In order to allow for the property of heterogeneity in the panel data framework[16], Dumitrescu and Hurlin (2011) propose extending the Granger causality approach by adding cross-sectional units to the time series dimension. Hence, they develop a test of no causality which accounts for Homogenous Non-Causality, i.e., no causal relation for any of the units of the panel) under the null. Under the alternative, they specify the heterogenous hypothesis defined as the cross-sectional average of the Wald statistics associated with the individual Granger causality tests. Two subgroups of cross-sectional units are therefore defined: one characterised by causal relations from $x$ to $y$[17] and another subgroup for which there is no causal relation. From the Panel Granger-causality test, Table 4 reveals that in the majority of the cases, the WTI crude oil price Granger-causes the commodity prices. These causal relations appear to be different according to the panel group considered. Indeed, almost all series, except Exotic, seem to be influenced by WTI price movements. The relation between oil and energy prices is not surprising since crude oil is often viewed as a key factor upon which energy markets are set. Just as with the crude oil market, Agro-Industrial and Non-Ferrous Metals are often considered to be good proxies for, respectively, global activity, the industrial and real estate sectors, then causal relations could be the consequence of international economic fluctuations. Turning to the Food and Oleaginous panel groups, causality is the consequence of input-output relations. Reverse causality from commodity groups to crude oil is not validated according to the $\tilde{Z}_{Hnc}$statistic. Because it is relevant that short run dynamics exist from WTI to the commodity markets, we propose to estimate the Panel VAR model and analyse the impulse response functions.

## TABLE 4: PANEL GRANGER CAUSALITY TEST

| | $\boldsymbol{WTI \rightarrow Commodities}$ | | | $\boldsymbol{Commodities \rightarrow WTI}$ | | |
|---|---|---|---|---|---|---|
| | $\boldsymbol{W_{Hnc}}$ | $\boldsymbol{Z_{Hnc}}$ | $\boldsymbol{\tilde{Z}_{Hnc}}$ | $\boldsymbol{W_{Hnc}}$ | $\boldsymbol{Z_{Hnc}}$ | $\boldsymbol{\tilde{Z}_{Hnc}}$ |
| **Energy** | **12.29** | **16.09** | **5.35** | **1.07** | **-3.34** | **-1.11** |
| **Precious metals** | **60.92** | **141.89** | **47.22** | **3.93** | **2.28** | **0.75** |
| **Non-Ferrous metals** | **4.74** | **5.22** | **1.73** | **2.41** | **-1.76** | **-0.58** |
| **Agro-Industrials** | **12.13** | **19.37** | **6.44** | **3.17** | **0.37** | **0.12** |
| **Food** | **5.73** | **6.70** | **2.22** | **1.74** | **-3.08** | **-1.02** |
| **Oleaginous** | **27.61** | **42.63** | **14.18** | **2.77** | **-0.38** | **-0.12** |
| **Exotic** | **3.10** | **0.21** | **0.07** | **1.79** | **-2.55** | **-0.08** |

Notes: Statistics are reported. The test statistic converges under the null to a standard normal distribution when *T* and *N* tend sequentially to infinity.

[16]. The authors distinguish between the heterogeneity of the regression model and that of the causal relation from $x$ to $y$.
[17]. The regression model is not constrained to be the same across units.

## 3. PANEL VECTOR AUTOREGRESSION AND IMPULSE RESPONSES

In order to assess the link between WTI crude oil prices, the stock market, and commodity prices, we consider the following panel VAR model[18].

$Y_{i,t} = \alpha_i + A(L)Y_{i,t} + \varepsilon_{i,t}$

where *i* indicates the type of commodity, *t* runs from 1 to *T*, $Y_{i,t}$ is the vector of endogeneous variables, $\varepsilon_{i,t}$ is the vector of error terms, $\alpha_i$ is the commodity-group specific intercept matrix, and $A(L)$ is the matrix polynomial in the lag operator.

The vector of endogenous variables is given in turn by

$$(Y_{i,t} = C_{i,t}, WTI_{i,t}, SP500_{i,t})'$$

where *i* denotes the individual dimension composed by commodity prices, and *t*=1,...$T$ the time. $C_{i,t}$, $WTI_{i,t}$, and $SP500_{i,t}$ denote the commodity prices[19], the crude oil prices, and the stock market index, respectively.

The usefulness of the panel VAR approach is that it combines a multivariate framework, which takes into account endogenous variables, with panel data dimensions, which takes into account unobserved individual heterogeneity. Furthermore, to overcome the well-known problem of correlation between the regressors and the fixed effects in dynamic panel specifications, we use the Generalised Method of Moments (GMM) method as an estimation procedure, using the codes of Love and Zicchino (2006).

The impulse response functions are given in the Appendix and indicate the impact of a WTI price shock on commodity markets. First, the panels Energy, Precious Metals, and Oleaginous exhibit very similar reactions (*i.e.*, a positive response at period *t+1*). Non-Ferrous Metals have a slower reaction, in terms of both time and magnitude, indicating a less close relation. Concerning the magnitude of the results, it is worth noting that the response at one day for the energy panel group is about 0.002, a little less than for the precious metals. The response of Oleaginous is about 0.0014. As for the agro-industrial panel group, no significant response appears. For the food panels, the response is negative but quite limited in terms of magnitude. This result is not a surprise for the energy group, given that energy prices are often contractually linked to the oil price. Considering the responses of precious metals, oleaginous, and non-ferrous metals, the influence of exchange traded funds can be an explanation, as well as hedging behaviour on the part of traders. Two conclusions emerge from our analysis: (i) considering the rapidity of the return to equilibrium, the hypothesis of market efficiency receives strengthened support, (ii) the fact that various panels have the same reaction emphasises the importance of exchange traded funds. Regarding the responses to a shock from the proxy for global economic activity, all group panels, except for the energy panel group, have the same kind of reaction: *i.e.*, a positive response at one day, followed by a fast return to equilibrium. One can interpret this result as evidence for the fact that these commodity prices remain driven by factors outside the sphere of influence of the price of oil. The magnitudes of the responses for precious and non-ferrous metals are quite similar and high, with a value of 0.0025. This highlights the large influence of global activity upon these variables. For the three other groups, the magnitude is much weaker: between 0.001 and 0.0015, reflecting nonetheless a mixed influence of global activity because of the more specific composition of these panels.

## 4. CONCLUSION

In this article, we focussed on the relation between oil and commodity prices, relying on second-generation panel unit root and cointegration tests. Using a collection of data with a wide scope (energy, precious metals, non-ferrous metals, agro-industrials, food, oleaginous, exotic, and livestock), we first

---

[18]. As mentioned in the description of the data, the SP500 is included in our panel VAR regressions to control for the economic and financial environment, something which might affect all commodity prices.

[19]. The variable $C_{i,t}$ considers one commodity panel group at a time.

conclude that the hypothesis of long-run relations between oil and the panel data groups is not validated. We then do not find evidence of co-movement in commodity prices in the long run. However, considering that short-run relations might exist due to hedging strategies such as cross hedging between correlated markets or new portfolio allocations, we then implemented a short-term analysis, using a panel Granger causality test and a panel VAR methodology. From the panel Granger causality test, we observed that the WTI crude oil price Granger-causes the commodity prices (except for the exotic commodities), while a reverse causality is not validated. From the Panel VAR methodology we find that three panel groups, namely energy, precious metals, and oleaginous, have a similar reaction at a lag of one day to a shock from the price of oil, while non-ferrous metals have a slower reaction. Our results support two main conclusions: (i) the speed of the return to long-term equilibrium for each commodity group emphasises the efficiency of the markets and the importance of the existing fundamental market forces, and (ii): fast co-movements are at play in the short run for these groups of commodities due to hedging behaviour and the influence of ETFs.

## REFERENCES

Ai, C., Chatrath, A., and Song, F.M., (2006), "On the Comovement of Commodity Prices", American Journal of Agricultural Economics, Vol. 88, No. 3, pp. 574-588.

Baffes, J., (2007), "Oil Spills on Other Commodities", The World Bank, Policy Research Working Paper, WPS4333.

Calabre, S., (1997), Filières Nationales et Marchés Mondiaux de Matières Premières, Paris, Economica.

Calabre, S., (2003), "La dynamique des prix et des marchés de matières premières : analyses univariées versus faits stylisés analytiques", Mondes en développement, Vol. 22, pp. 21-35.

Cashin, P., McDermott, C. J., and Scott, S., (1999), "The Myth of Comoving Commodity Prices", IMF Working Paper, WP/99/169.

Choi, I., (2002), "Combination unit root tests for cross-sectionally correlated panels", Mimeo, Hong Kong University of Science and Technology.

Chicago Mercantile Exchange Group website: http://www.cmegroup.com/.

Coleman, N., and Levin, C., 2006. The role of market speculation in rising oil and gas prices: A need to put the cop back on the beat, Committee on Homeland Security and Governmental Affairs. Permanent Sub-commitee on Investigations, Washington, DC.

Deb, P., Trivedi, P.K., and Varangis, P., (1996), "The excess Co-movement of Commodity Prices Reconsidered", Journal of Applied Econometrics, Vol. 11, pp. 275-291.

Dickey, D.A., and Fuller, W.A., (1981), "Likehood Ratio Statistics for Autoregressive Time Series With a Unit Root", Econometrica, Vol. 49, pp. 1057-1072.

Dumitrescu, E.-I., and Hurlin, C., (2012), "Testing for Granger Non-causality in Heterogeneous Panels", Economic Modelling, forthcoming.

Hache, E., and Lantz, F., (2013), "Speculative trading and oil price dynamic: A study of the WTI market", Energy Economics, Vol. 36, pp. 334-340.

Hadri, K. (2000), "Testing for unit roots in heterogeneous panel data", Econometrics Journal, Vol. 3, pp. 148-161.

Houillon, C.-A., (2005), Guide pratique des marchés de matières premières et de l'Energie, SEFI éditions.

Hurlin, C., and Mignon, V., (2005), "Une synthèse des Tests de Racine Unitaire sur Données de Panel", Economie et prévision, Vol. 169-171, pp. 251-295.

Hurlin, C., and Mignon, V., (2007), "Une synthèse des Tests de Cointégration sur Données de Panel", Economie et prévision, Vol. 180-181, pp. 241-265.

Im, K.S., Pesaran, M.H., and Shin, Y., (2003), "Testing for unit roots in heterogeneous panel", Journal of Econometrics, Vol. 115, pp. 53-74.

Joëts, M., and Mignon, V., (2011), "On the link between forward energy prices: A nonlinear panel cointegration approach", Energy Economics, Vol. 34, No. 4, pp. 1170-1175.

Levin, A., Lin, C.F., and Chu, C.S.J., (2002), "Unit Root Test in Panel Data: Asymptotic and Finite Sample

Properties", Journal of Econometrics, Vol. 108, pp. 1-24.
Leybourne, S.J., Lloyd, T.A., and Reed, G.V., (1994), "The excess Comovement of Commodity Prices Revisited", World Development, Vol. 22, No. 11, pp. 1747-1758.
London Metal Exchange website: http://www.lme.com/.
Love, I., and Zicchino, L., (2006), "Financial development and dynamic investment behavior: Evidence from a panel VAR", The Quarterly Review of Economics and Finance, Vol. 46, pp. 190-210.
Masters, M.W., (2008). Testimony before the Committee on Homeland Security and Governmental Affairs. United States Senate. May 20.
Masters, M.W., and White, A.K. 2008(a). The accidental Hunt brothers-Act 2: Index speculators have been a major cause of the recent drop in oil prices. Special Report September 10.
Masters, M.W., White, A.K. 2008(b). The accidental Hunt brothers: How institutional investors are driving up food and energy prices. Special Report July 31.
Natanelov, V., Alam, M.J., McKenzie, A.M., and Van Huykenbroeck, G., (2011), "Is there co-movement of agricultural commodity futures prices and crude oil", Energy Policy, Vol. 39, pp. 4971-4984.
Nazlioglu, S., and Soytas, U., (2012), "Oil Price, agricultural commodity prices, and the Dollar: A panel cointegration and causality analysis", Energy Economics, Vol. 34 No. 4, pp. 1098-1104.
Palaskas, T.B., and Varangis, P.N., (1991), "Is there Excess Co-movement of Primary Commodity Prices: A Cointegration Test?", The World Bank, Working Papers, WPS0758.
Pedroni, P. (1999), "Critical values for cointegrations tests in the heterogeneous panels with multiple regressors", Oxford Bulletin of Economics and Statistics, S1, Vol. 61, pp. 653-670.
Pedroni, P. (2004), "Panel cointegration. Asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis", Econometric Theory, Vol. 20, pp. 597-625.
Pindyck, R.S., and Rotemberg, J.T., (1990), "The excess Co-Movement of Commodity Prices", The Economic Journal, Vol. 100, No. 403, pp. 1173-1189.
Radetzki M., (2006), "The anatomy of three commodity booms", Resources Policy No. 31, pp. 56-64.
Saadi, H. (2001), "Le phénomène des mouvements joints des prix internationaux de matières premières", Tiers monde, Vol. 42, No. 168, pp. 865-883.
Westerlund, J. (2007), "Testing for error correction in panel data", Oxford Bulletin of Economics and Statistics, Vol. 69, pp. 709-748.

## APPENDIX

**FIGURE 1. ENERGY AND PRECIOUS METALS MARKETS GROUPS**

**TABLE 5: CROSS-CORRELATION OF THE ERRORS AND CD TEST**

| | $\widehat{\rho}$ | *CD* |
|---|---|---|
| **Energy** | **0.375** | **20.10 (0.00)** |
| **Precious metals** | **0.585** | **76.85 (0.00)** |
| **Non-ferrous metals** | **0.875** | **181.63 (0.00)** |
| **Agro-industrials** | **0.200** | **18.62 (0.00)** |
| **Food** | **0.681** | **89.44 (0.00)** |
| **Oleaginous** | **0.974** | **52.22 (0.00)** |
| **Exotic** | **0.715** | **66.38 (0.00)** |
| **Livestock** | **0.636** | **107.77 (0.00)** |

Notes: Between parentheses: *p*-values. The CD statistics are standard Normal under the null hypothesis of cross-section independence.

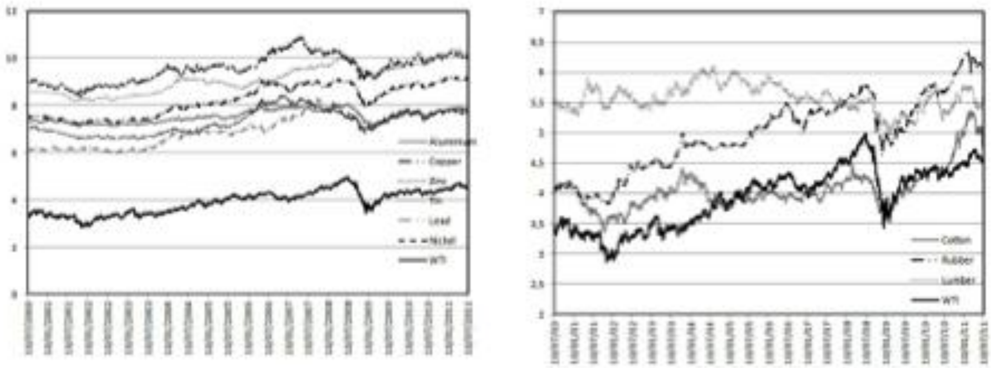

**FIGURE 2. NON-FERROUS METALS AND AGRO-INDUSTRIALS PANEL GROUPS**

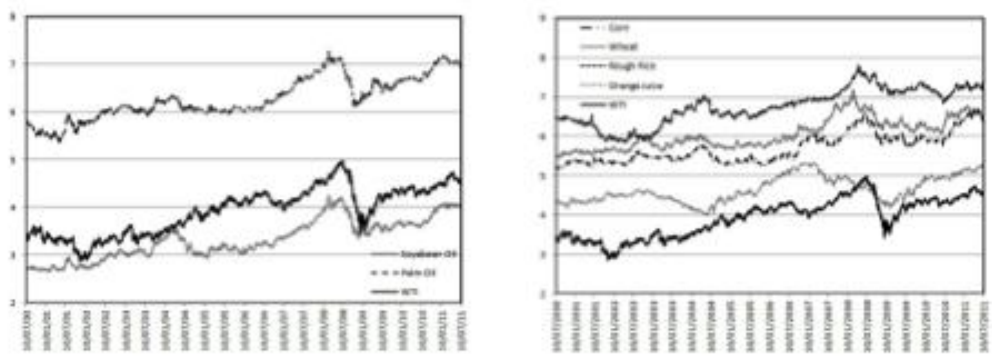

**FIGURE 3. OLEAGINOUS AND FOOD PANEL GROUPS**

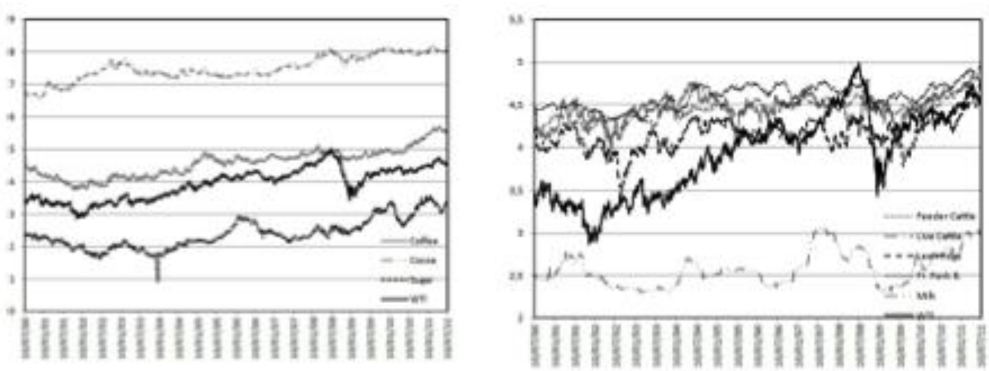

**FIGURE 4. EXOTIC AND LIVESTOCK PANEL GROUPS**

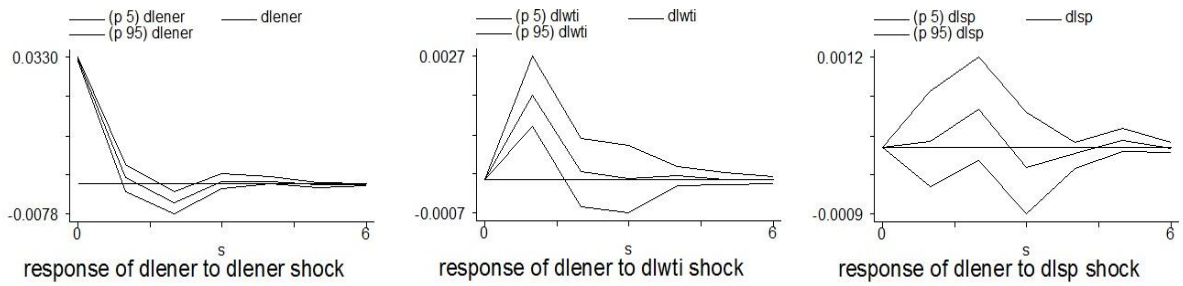


**FIGURE 5. IMPULSE RESPONSE FUNCTIONS OF ENERGY PANEL GROUP**

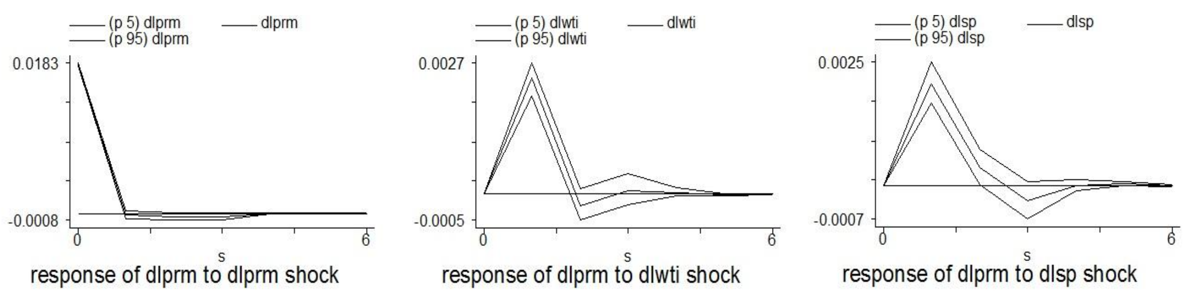


**FIGURE 6. IMPULSE RESPONSE FUNCTIONS OF PRECIOUS METALS PANEL GROUP**

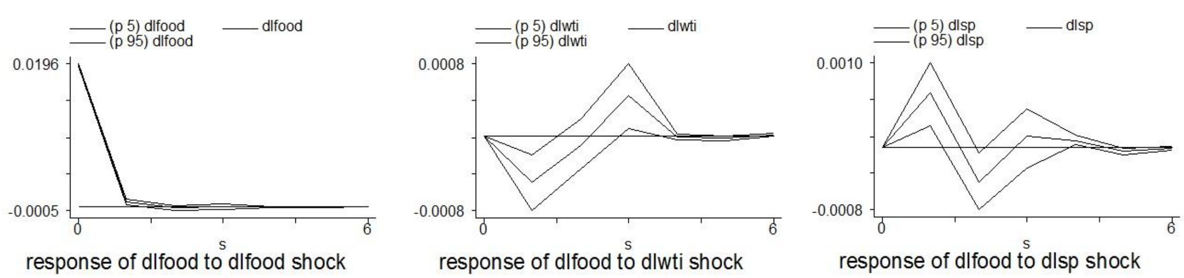


**FIGURE 7. IMPULSE RESPONSE FUNCTIONS OF FOOD PANEL GROUP**

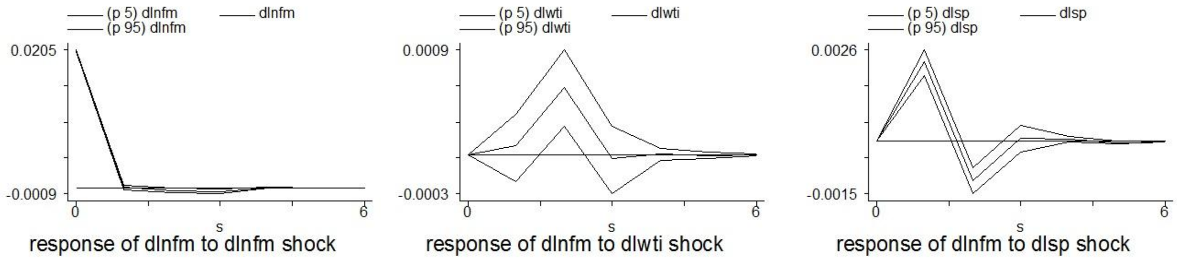


**FIGURE 8. IMPULSE RESPONSE FUNCTIONS OF NON-FERROUS METALS PANEL GROUP**

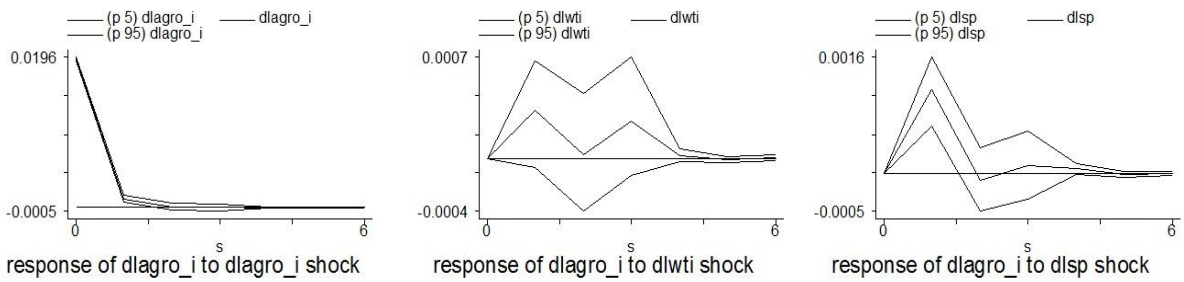


**FIGURE 9. IMPULSE RESPONSE FUNCTIONS OF AGRO-INDUSTRIAL PANEL GROUP**

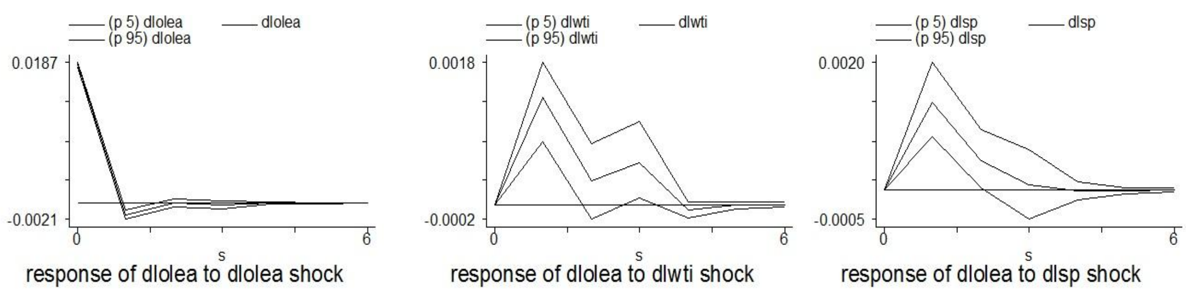


**FIGURE 10. IMPULSE RESPONSE FUNCTIONS OF OLEAGINOUS PANEL GROUP**