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Spot and futures commodity markets and the unbiasedness hypothesis - evidence from a novel panel unit root test

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Abstract

A controversial view of the evolution of commodity markets is that the engagement of speculative capital arguably introduces volatility and price movements unrelated to changes in traditional demand and supply factors. Thus, the efficiency of spot and futures markets is an important topic in this context, as the price of a futures contract in the current period should be an unbiased estimator of next period´s spot price under the joint assumption of risk neutrality and rationality. In this vein, the present study contributes to the literature by applying the novel panel unit root test provided by Demetrescu and Hanck (2012) which simultaneously allows for cross-sectional dependence and unconditional heteroskedasticity. Our findings show that most spot and futures markets for commodities were efficient until the turn of the millennium, but appear to be inefficient thereafter owing to an increase in volatility, which might be attributed to the intense engagement of speculation in commodity markets.

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1. Introduction

The increasing engagement of speculative funds and financial derivatives in commodity markets over the last decade has raised the question of whether spot and futures markets still work efficiently. The potential impact of speculative trading in these markets has created an ongoing discussion in both academic literature and the media in recent years (Masters, 2008; Masters and White, 2008; CFTC, 2008; Chilton, 2008; Piesse and Thirtle, 2009; Chevillon and Rifflart, 2009; Cifarelli and Paladino, 2009; Kaufmann and Ullman, 2009; Irwin et al., 2009; Sanders and Irwin, 2010; Fan and Xu, 2011; Wright, 2011; Lombardi and van Robays, 2011; Irwin and Sanders, 2011, 2012; Bohl and Stephan, 2012). The controversial question is whether futures markets perform a price discovery function or if they offer a stage for investors to create non-fundamental price pressures.

Following Fama (1970), the concept of market efficiency is based on the principle that prices reflect all publicly available information and that, therefore, investors cannot gain arbitrage revenues.[1] Thus, under the joint assumption of risk neutrality and rationality the price of a futures contract in the current period for delivery in the next period is an unbiased estimator of the expected next period's spot price. Following Brenner and Kroner (1995), the adjustment process of spot and forward prices should ensure that investors will be indifferent between (a) the purchase of a particular commodity at the spot market while accepting the storage costs and benefiting from a convenience yield, and (b) the investment in a risk-free bond and the purchase of the currently quoted futures contract for that commodity later at the forward market. If the presence of a cointegrating relation between spot and futures prices reflects this arbitrage condition, the unbiasedness hypothesis should hold in the long-run (Kellard, 2002).

Previous studies have applied several cointegration techniques when testing this hypothesis. However, the results are mixed (MacDonald and Taylor, 1988; Serletis and Banack, 1990; Sephton and Cochrane, 1990, 1991; Schroeder and Goodwin, 1991; Chowdhury, 1991; Lai and Lai, 1991; Crowder and Hamed, 1993; Krehbiel and Adkins, 1993; Schwartz and Szakmary, 1994; Aulton et al., 1997; Chow, 1998; Peroni and McNown, 1998; Gülen, 1998; McKenzie and Holt, 2002; Kellard, 2002; Wang and Ke, 2005; He and Hong, 2011). The potential incapability of cointegration tests to produce evidence for the efficiency of spot and futures markets for commodities has been vindicated by, among other factors, the inefficiency of agents in conveying new information to the market (Kaminsky and Kumar, 1990), the inability of futures prices to reflect all publicly available information (Beck, 1994), and the existence of a risk premium (Kellard et al., 1999; He and Hong, 2011).

[1] Fama (1970) also distinguishes between weak, semi-strong and strong forms of market efficiency.

However, owing to the fact that standard unit root and cointegration tests for individual time series may suffer from low power, panel cointegration techniques may be useful when analyzing the question of market efficiency (Campbell and Perron, 1991).[2] Compared with commodity by commodity studies, panel data analyses have the advantage of increasing the sample size and *ceteris paribus* lead to more precise estimates. Considering that financial data is not usually cross-sectionally independent, it is also important to control for the occurrence of cross-section dependence, since the assumption of independent panel members results in a misleading inference (Banerjee et al., 2004; Beckmann et al., 2012). Cross-section dependence can result either from the occurrence of common stochastic trends or from correlations between errors across panel members (Breitung and Pesaran, 2008).[3] To the best of our knowledge, there is no study which accounts for this issue when analyzing the market efficiency of spot and futures prices for commodities.

Hence, we contribute to the existing literature by testing market efficiency in spot and futures markets for several commodities following a residual-based cointegration approach. The novel panel unit root test applied in this study has been developed by Demetrescu and Hanck (2012) and allows simultaneously for cross-sectional dependence between different commodities and for unconditional heteroskedasticity.

As to the remainder of the paper, the following section presents our methodology while the empirical results of our study are given in Section 3. Section 4 concludes.

2. Methodology

As stated above, under the joint assumption of risk neutrality and rationality, market efficiency implies that the price of a futures contract in the current period for delivery in the next period is an unbiased estimator of the expected next period's spot price (Gülen, 1998; Kellard, 2002; Switzer and El-Khoury, 2007; Huang et al., 2009; Lin and Liang, 2010). Thus, the unbiasedness hypothesis is given below:

$$E_t(s_{i,t+k}) = f_{i,t,k}, \quad i = 1, \dots, N, \quad t = 1, \dots, T, \tag{1}$$

where $s_{i,t+k}$ denotes the logarithm of the spot price of commodity i at time $t+k$, $f_{i,t,k}$ denominates the logarithm of the price of a futures contract for commodity i observed at time t for delivery at time $t+k$, and $E_t(.)$ gives the expectations operator conditional on information available at time t. Equation (1) can be transformed as

[2] The problem of low power is of particular importance if a short span of data is analyzed. However, even for the dataset under investigation in this study, further insights may be gained due to this argument.

[3] The latter does not necessarily imply cointegration across panel members (Breitung and Pesaran, 2008).

follows:

$$s_{i,t+k} = f_{i,t,k} + u_{i,t+k}, \tag{2}$$

where $u_{i,t+k}$ indicates an uncorrelated random error term with zero mean and constant variance. Considering that spot and futures prices for commodities are usually found to be integrated of order one, e.g. $I(1)$,[4] equation (2) yields the following cointegrating regression:

$$s_{i,t+k} = \beta_{i,0} + \beta_{i,1} f_{i,t,k} + u_{i,t+k}. \tag{3}$$

Market efficiency requires that spot and futures prices be cointegrated with $\beta_{i,0} = 0$ and $\beta_{i,1} = 1$. Thus, the efficiency hypothesis corresponds to testing whether a proportional long-run relation between the futures and the spot price,

$$b_{i,t+k} = f_{i,t,k} - s_{i,t+k}, \tag{4}$$

is stationary, e.g. $I(0)$, for each commodity. Hence, we rely on a residual-based cointegration framework in the spirit of Engle and Granger (1987) and check whether the proportional relation between the futures and the spot price for commodity i contains a unit root. In order to test for a unit root in the series $b_{i,t+k}$ for each commodity included in our panel dataset we apply the robust Hartung (1999) panel unit root test provided by Demetrescu and Hanck (2012). As previously mentioned, first generation panel unit root tests suffer from dramatic size distortions in the presence of cross-section dependence (Banerjee et al., 2004; Lyhagen, 2008).

To account for this, we start by estimating the following ADF-type test regression

$$\Delta x_{i,t} = \phi_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta x_{i,t-j} + \varepsilon_{i,t}, \tag{5}$$

where $x_{i,t} = b_{i,t+k} + d_i$ and d_i is considered to be either a constant, or a linear trend plus constant. Under the unit root null, $\phi_i = 0 \quad \forall \quad i$. This instrumental variable (IV) Cauchy estimator uses the sign of the first lag $x_{i,t-1}$ as an instrumental variable for $x_{i,t-1}$ and allows for unconditional heteroskedasticity as well as cross-sectional dependence between the i's.[5] Demetrescu and Hanck (2012) have shown that the IV t-type statistic follows a standard normal limiting distribution and that the new version of the test has more power relative to second-generation panel unit root tests, such as those proposed by Im et al. (2003). The latter kinds of test are also robust

[4] See among others Shen and Wang (1990).

[5] The lagged terms of the endogenous variable $\Delta x_{i,t-j}$, $j = 1, \ldots, p_i$, are used as instruments for themselves. See Demetrescu and Hanck (2011, 2012) for details.

to cross-sectional dependence. However, they are not able to handle unconditional heteroskedasticity. Demetrescu and Hanck (2012) suggest that N should be smaller than T, a condition which is considerably fulfilled for the dataset under investigation which is described in the next section.

3. Data and empirical results

Our analysis is based on a panel dataset from the Dow Jones UBS Commodity Index (DJ-UBSCI), which is composed of commodities predominantly traded on U.S. exchanges and provided by Dow Jones Indexes (`http://www.djindexes.com/commodity/`). The DJ-UBSCI is weighted by the relative amount of trading activity of a particular commodity and was known as the Dow Jones AIG Commodity Index until 2009. Alternatively, the S&P Goldman Sachs Commodity Index (GSCI) might be used instead of the DJ-UBSCI, but according to Tang and Xiong (2010), the correlation between the GS and the DJ-UBS commodity indices is over 0.9. Thus, applying the GSCI would most likely not change our findings.[6] More precisely, we apply the subindex of prices for spot as well as for the three month futures contracts for 19 commodities on a daily basis, namely coffee, corn, cotton, lean hogs, live cattle, soybeans, soybean oil, sugar, wheat, crude oil, heating oil, natural gas, unleaded gasoline, aluminum, copper, gold, nickel, silver, and zinc. The series are shown in Figure 1, below. We also divide the whole panel into three sub-panels which are classified as agricultural products, energy commodities, and metals. Our sample period covers every working day from January 2, 1991 to October 19, 2011 and thus exhibits the largest available sample size up to the start of our study, which contains the low volatility period up to the year 2000 as well as the high volatility period thereafter, as displayed in Figure 1.[7] Therefore, we split the whole sample period into two subsamples roughly reflecting the low and high volatility period. As mentioned in Section 2, each series is taken as log.

Figure 1 about here

In order to test for a proportional long-run relationship between the futures and the spot price given in equation (4) for each commodity, we have to ensure that each of the spot and futures prices is $I(1)$. Thus, we use the augmented Dickey-Fuller (ADF) test and the more powerful Ng-Perron MZ_α test, either with a constant

[6]See Tang and Xiong (2010) and Gilbert (2010) for details regarding the DJ-UBSCI and its subindices. Among others, Irwin and Sanders (2012) have considered the DJ-UBSCI dataset in a similar case.

[7]Fan and Xu (2011) divided the price fluctuations in the oil market after 2000 into three stages in their study: the 'relatively calm market' period (from January 7, 2000, to March 12, 2004), the 'bubble accumulation' period (from March 19, 2004, to June 6, 2008), and the 'global economic crisis' period (from June 13, 2008, to September 11, 2009).

or with a linear trend plus constant to check the null of a unit root for the levels and first differences of each series (Dickey and Fuller, 1979; Ng and Perron, 2001). The results clearly indicate that the null cannot be rejected for the levels but for the differences of each series. Moreover, to account for the possibility of structural breaks, which in general reduce the power of conventional unit root tests to reject the null, we also applied the same tests for both subsample periods and conducted the Perron (1989) test. In addition, we applied the panel unit root test described above to check the unit root null for the panel of spot prices as well as the panel of futures prices. In both cases the findings remain the same. Thus, as expected, each spot and futures price can be regarded as $I(1)$. To save space, our results for the unit root tests are not reported, but they are available upon request.

In the next step, we apply the panel unit root test to check market efficiency for our different samples, as described above, and present our results in Table 1.

Table 1 about here

For the whole sample period the unit root null cannot be rejected for each of the four different panel sets. Thus, it seems that spot and futures markets for commodities have not been efficient for the whole of the period under investigation. However, as mentioned before, prices for commodities were relatively stable before the year 2000 and have shown a considerable increase in variance thereafter. Thus, it seems reasonable for us to split the sample period after 2000. As expected, the null of a unit root in each series of the whole panel can be rejected at least at a significance level of 10% with regard to the first subsample period. In particular, the results for the sub-panels containing agricultural commodities and metals support that finding. The results also suggest that the spot and futures markets for energy commodities such as crude oil, heating oil, natural gas, and unleaded gasoline were inefficient during that period. Unsurprisingly, the outcomes for the second subsample period clearly provide evidence of market inefficiency. In a nutshell, one can conclude that commodity markets were efficient up to the turn of the millennium while they have become inefficient thereafter. Energy commodity markets are the only exceptions, as they have been inefficient for the whole sample period.

4. Conclusion

Our study has focused on the efficiency of spot and futures prices of the most essential commodities by trading activity. Firstly, we have demonstrated the usefulness of the applied cointegration framework based on a novel panel unit root test, which accounts for both cross-sectional dependence as well as changing variances. In particular, the latter feature seems to be appropriate in our study, since commodity prices turn out to have been highly volatile in recent years. Secondly, our findings indicate that commodity markets were efficient up to 2000 and have become inefficient thereafter. Energy commodity markets, which were inefficient for the whole of the period under review, are the only exceptions. Hence, our findings overall support the argument that the increasing engagement of investors who regard commodities such as agricultures and metals as an asset class may have destabilized those markets. However, further research is necessary to validate this argument, in particular for specific groups of commodities, considering that understanding and explaining the behavior of commodity prices is a notoriously difficult task.

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A. Tables

Table 1: Panel unit root tests

Sample period	Panel	Test statistic		
		$m = 0$	$m = 1$	$m = 2$
Full sample	All commodities ($N = 19$)	-0.458	-0.923	-0.950
(1/1/1991 - 10/19/2011)	Agriculture ($N = 9$)	-0.605	-1.054	-1.073
$T = 5426$	Energy ($N = 4$)	-0.146	-0.745	-0.837
	Metals ($N = 6$)	-0.497	-0.828	-0.828
Subsample I	All commodities ($N = 19$)	-1.307*	-2.012**	-2.012**
(1/1/1991 - 12/31/2000)	Agriculture ($N = 9$)	-1.805**	-1.759**	-1.759**
$T = 2608$	Energy ($N = 4$)	-0.420	-0.916	-0.916
	Metals ($N = 6$)	-1.398*	-3.093***	-3.093***
Subsample II	All commodities ($N = 19$)	-0.281	-0.676	-0.667
(1/1/2001 - 10/19/2011)	Agriculture ($N = 9$)	-0.224	-0.756	-0.749
$T = 2817$	Energy ($N = 4$)	-0.369	-0.678	-0.651
	Metals ($N = 6$)	-0.305	-0.560	-0.560

Note: * Statistical significance at the 10% level, ** at the 5% level, *** at the 1% level. Critical values: 1% -2.326, 5% -1.645, 10% -1.281. T denotes the number of observations for each series, N represents the panel width and m denominates the instrument-generating function. The resulting test statistics are obtained by using an IV-ADF test regression without any lagged regressors of the endogenous variable. To account for potential autocorrelation in the residuals of the latter we also considered several lag lengths. However, we obtained qualitatively the same outcomes. Thus, we rely on the given findings.

B. Figures

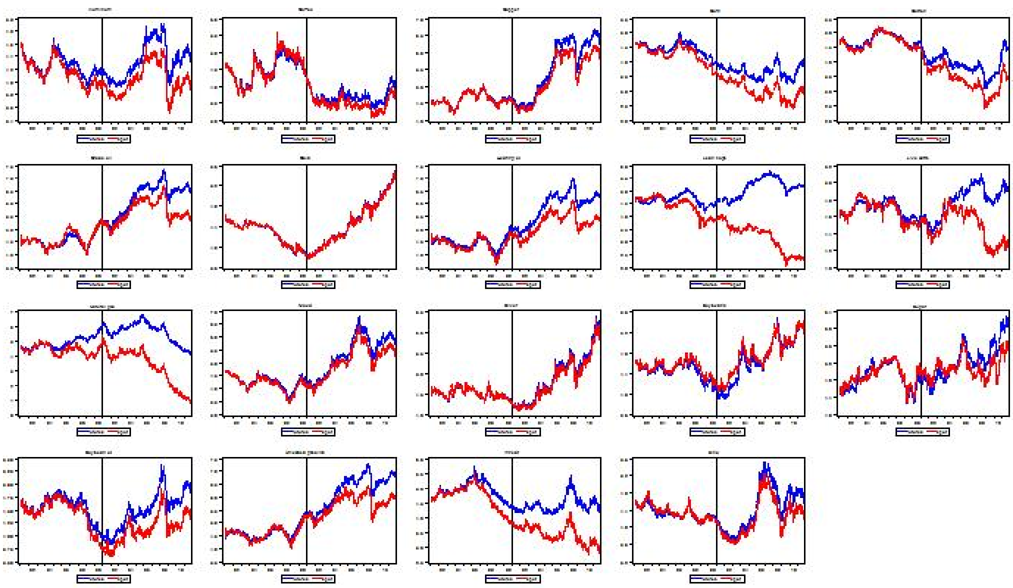

Figure 1: Logs of the time series of the futures and spot prices for the 19 commodities

Note: futures (blue), spot (red); aluminum, coffee, copper, corn, cotton (first row), crude oil, gold, heating oil, lean hogs, live cattle (second row), natural gas, nickel, silver, soybeans, sugar (third row), soybean oil, unleaded gasoline, wheat, zinc (fourth row); vertical line: 12/31/2000.