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Dynamic linkages among European carbon markets

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Abstract

A number of studies have tested for cointegration between spot and futures prices in the European carbon markets. These studies tend to focus on the price discovery role of futures versus spot prices. In this paper, we draw the attention to the short- and long-run dynamic linkages among distinct European carbon markets by investigating the interdependence and the transmission efficiency between European Climate Exchange (ECX), Nordic Power Exchange (NordPool) and European Energy Exchange (EEX). To this end, we test for cointegration between European Union carbon allowances (EUAs) futures prices and also, we conduct causality tests to examine spillover dynamics. Our findings indicate that the markets exhibit a reasonable degree of efficiency in both short- and long-run.

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1. Introduction

The European Union (EU) has created the largest Emissions Trading Scheme (ETS) in the world in order to reduce dioxide carbon (CO2) emissions by companies from the energy and other carbon-intensive industries. The EU ETS is being introduced in three phases. The first phase which ran from 2005 to 2007 is considered as a pilot phase; the second phase which ranges from 2008 to 2012, coincides with the period when the EU must meet the 8% decrease in emissions from 1990 levels under the Kyoto Protocol. As proposed recently by the European Commission, the third phase will ran from 2013 to 2020. In order to improve the fluidity of the EU ETS, organized allowance trading has been segmented across trading platforms namely European Climate Exchange (ECX, futures contracts), Nordic Power Exchange (NordPool, spot and futures contracts), Powernext[1] (spot contracts) , European Energy Exchange (EEX, spot and futures contracts), Energy Exchange Austria (EXAA, spot contracts) and Climex (spot contracts).

This study attempts to investigate the price transmission among markets for European Union carbon allowances (EUAs) by providing an important insight as to how price shocks at any market are transmitted to all other market prices, thus reflecting the extent of dynamic market linkages, as well as the extent to which markets function efficiently and considers three questions:

- Is there long-run interdependence in European carbon markets in the sense that the equilibrium for one market depends on the equilibrium for other market?
- Is there short-run interdependence in European carbon markets? In other words, do short-run fluctuations in one market spillover to the other?
- What is the direction of causality between these carbon markets? Can we identify one market as being the « cause » and the other the « effect »?

The paper is motivated by several reasons. First, investigating the price transmission is crucial, since it is an important indicator of market efficiency as the objective of carbon markets is to enable firms to achieve their emissions reductions at minimum cost. Second, European carbon markets represent a large segment of international carbon markets and understanding the linkages between them is important for firms and investors to design effective portfolio diversification strategies[2]. If there is significant comovement across European carbon markets, the benefits of international diversification might not be realized in the long-run. Finally, prior studies primarily focus in assessing efficiency and price discovery from spot and futures carbon markets with less attention paid to transmission efficiency on distinct trading platforms. For examples, Uhrig-Homburg and Wagner (2007) assume that the spot and futures price dynamics for EUAs can be described sufficiently well with the cost-of-carry approach after December 2005, meaning that spot prices plus accrued interest should be equal to futures prices. Their empirical results suggest that after December 2005 the market efficiency increased, and spot and futures prices seem to be linked by the cost-of-carry approach. Daskalikas *et al.* (2007) find that the pricing mechanism of intra-phase and inter-phase derivatives in Nordpool, Powernext and ECX is very different due to the prohibition of banking between the distinct phases of the market. They also find that the substantial stochastic convenience yields in inter-phase futures markets imply additional uncertainty and

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[1] In December 2007, Bluenext has taken over Powernext's spot carbon market. In April 2008, Bluenext has launched EUA futures contracts for delivery at maturities from December 2008 to 2012.

[2] Oberndorfer (2008) and Veith et al. (2009) find that EUA price affects significantly the value of electricity companies. Furthermore, Boutaba (2009) discovers that EUA price moves oil companies' equity values.

hedging costs for market participants. They conclude that the EUA market is efficient. Milunovich and Joyeux (2007) find that none of the futures contracts traded on the European Climate Exchange (ECX) follow a cost-of-carry relationship with the spot price and interest rates, suggesting that the existence of arbitrage opportunities in the EUA market. They also find that the spot and futures markets share information efficiently and contribute to price discovery jointly. Using daily EUA spot prices from the European Energy Exchange (EEX), Seifert *et al*. (2008) test the hypothesis of no autocorrelation in EUA returns and conclude that the EUA market seems to be relatively efficient compared to the U.S SO2 permit market and the DAX. Daskalakis and Markellos (2008) assessed the weak form efficiency by analysing spot and futures market data from Powernext, Nordpool and ECX. Their empirical results reveal that EUA returns are serially predictable and that simple trading strategies can be employed in order to exploit these predictabilities and to produce substantial risk-adjusted profits. They explain the inefficiency of the EUA market by the existing restrictions on short-selling and banking EUAs. Benz and Klar (2008) is the only work which considered price discovery between futures prices on distinct markets (ECX and Nord Pool). Their results revealed that trading frictions in forms of transaction costs have decreased over the first trading phase, trading volume has increased and price discovery takes place across trading platforms.

This paper empirically investigates the dynamic linkages among European carbon markets. Using daily EUA futures prices, we have examined EU ETS efficiency by investigating the interdependence between European Climate Exchange (ECX), Nordic Power Exchange (NordPool) and European Energy Exchange (EEX). The analysis was performed with the aid of time series analysis techniques such as unit root tests with and without structural break, cointegration tests, vector error-correction models and Granger causality tests. Results show that the three carbon markets exhibit a reasonable degree of efficiency in both long-and short-run. The remainder of the paper is structured as follows. Section 2 describes the empirical methodology. Section 3 presents the data and the empirical results. Section 4 concludes.

2. Empirical methodology

The assessment of EUA market interdependencies is based on the joint testing for the presence and number of cointegrating vectors as well as on considering the relevant error correction model for causal relationship between these EUA markets. Indeed, the empirical analysis is done in three steps. The first step is to verify the non stationarity of the variables since the cointegration test is valid only if all the variables are integrated of order 1 (I (1)). The second step involves testing for cointegration using the Johansen procedure (Johansen 1988; Johansen and Juselius 1990). Cointegration allows estimation and testing of a long-run equilibrium relationship in the presence of short-run deviations from equilibrium. The Johansen method consists of estimating and testing the number of cointegrating relationships and common stochastic trends among the components of a vector z_t of non-stationary variables. Let z_t be an $(n \times 1)$ vector of I (1) variables. Then, it is possible to specify the following unrestricted vector autoregression (VAR) involving up to k-lags of z_t:

$$z_t = A_1 z_{t-1} + ... + A_k z_{t-k} + \mu_t , \tag{1}$$

where A_i is an $(n \times n)$ matrix of parameters and μ_t are a Gaussian error term. The above equation can be expressed as a vector error-correction form:

$$\Delta z_t = \Pi z_{t-1} + \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \mu_t, \tag{2}$$

where $\Pi = \sum_{i=1}^{k} A_i - I_n$ and $\Gamma_i = -\sum_{i=i+1}^{k} A_i$. The Johansen test focuses on the analysis of the Π matrix and Π can be interpreted as a long-run coefficient matrix, since in equilibrium, all the Δz_{t-i} value will be zero, and setting the error terms, μ_t, to their expected value of zero will leave $\Pi z_{t-1} = 0$. Testing for cointegration is related to the consideration of the rank of Π, that is finding the number of r cointegrating vectors. Two test statistics can be used for determining the number of cointegrating vectors under the Johansen approach. First, the trace test, i.e. the likelihood ratio test statistic for the hypothesis that there are at most r distinct cointegrating vectors against a general alternative, given by:

$$\lambda_{trace}(r) = -2\log(Q) = -T\sum_{i=r+1}^{n} \log(1-\hat{\lambda}_i) \tag{3}$$

where $\hat{\lambda}_i$s are the $(n-r)$ smallest squared canonical correlations of z_{t-1} with respect to Δz_t corrected for lagged differences and T is the sample size used for estimation. Alternatively, the maximum eigenvalue test can be used to compare the null hypothesis of r cointegrating vectors against the alternative of $(r+1)$ cointegrating vectors. The likelihood ratio test statistic for this hypothesis given by:

$$\lambda_{\max}(r, r+1) = -2\log(Q) = -T\log(1-\hat{\lambda}_{r+1}) \tag{4}$$

If the test statistic is greater than the critical values then we reject the null hypothesis that there are r cointegration vectors in favour of the alternative that there are r + 1 (for λ_{trace}) or more than r (for $\lambda_{\max}$).

The third step involves the estimation of the vector error-correction model (VECM) which captures the short-run dynamics of the variables. Statistical tests on the individual equations in the VECM can be used to determine the direction of Granger-causality between pairs of variables. In general, a VECM for three cointegrated variables take the following form:

$$\Delta x_t = \beta_1 + \sum_{i=1}^{m} \alpha_{1i} \Delta x_{t-i} + \sum_{i=1}^{m} \gamma_{1i} \Delta y_{t-i} + \sum_{i=1}^{m} \eta_{1i} \Delta v_{t-i} + \omega_1 ECT_{t-1} + \mu_{1t} \tag{5}$$

$$\Delta y_t = \beta_2 + \sum_{i=1}^{m} \alpha_{2i} \Delta x_{t-i} + \sum_{i=1}^{m} \gamma_{2i} \Delta y_{t-i} + \sum_{i=1}^{m} \eta_{2i} \Delta v_{t-i} + \omega_2 ECT_{t-1} + \mu_{2t} \tag{6}$$

$$\Delta v_t = \beta_3 + \sum_{i=1}^{m} \alpha_{3i} \Delta x_{t-i} + \sum_{i=1}^{m} \gamma_{3i} \Delta y_{t-i} + \sum_{i=1}^{m} \eta_{3i} \Delta v_{t-i} + \omega_3 ECT_{t-1} + \mu_{3t} \tag{7}$$

where x, y and v are the variables, Δ is the difference operator, m is the number of lagged difference terms determined in the cointegrating relationship, μ_{1t}, μ_{2t} and μ_{2t} are uncorrelated disturbance terms with zero means and finite variances and the lagged term (ECT) is the error-correction term, obtained from the long-run cointegrating relationship. The ECT ensures deviations from long-run equilibrium are corrected gradually through a series of partial short-run adjustments. The magnitude of the coefficients ω_1, ω_2 and ω_3 determines the speed of adjustment back to the long-run equilibrium state, once the system is shocked. The dynamic

Granger causality can be captured from the vector error-correction model (VECM) by using three channels of causality:

i) By observing the significance of the lagged values of the differenced variables; this is a measure of short-run (or weak Granger) causality. This can be tested using Wald test or the t-test if the lag order of equations is 1.
ii) By observing the error-correction term as a measure of long-run causality. This can be tested be by the t-test.
iii) By testing the joint significance of the first two channels of causation (the error-correction term and the lagged variables in each VECM). This can be tested through a Wald or F-test. The joint test indicates which variable(s) bear the burden of short run adjustment to re-establish long run equilibrium, following a shock to the system (strong causality).

3. Data and empirical results

The data for this study consists of observations on the daily settlement futures prices on EUA contracts for delivery at maturities from December 2006 to 2009[3]. These futures contracts trade on the European Energy Exchange (EEX), the European Climate Exchange (ECX) and the Nordic Power Exchange (NordPool) and contract specifications as well as trading details are available from their websites (www.eex.com, www.europeanclimateexchange and www.nordpool.com). The use of futures prices instead of spot prices is justified by the high liquidity on EUA futures markets. Sample lengths are October 4, 2005–November 29, 2006 (EUA-DEC 06 contracts), October 4, 2005–November 29, 2007 (EUA-DEC 07 contracts), December 2, 2005–December 1, 2008 (EUA-DEC 08 contracts), and January 2, 2006–December 12, 2008 (EUA-DEC09 contracts). Table 1 presents descriptive statistics for the corresponding EUA futures price series. All series are skewed to the right. The Jarque-Bera statistic rejects the null hypothesis for normality for all EUA futures price series.

3.1 Stationarity tests:

As indicated above, the variables must be tested for stationarity before running the cointegration tests. To this end, we first conducted 3 conventional unit root tests, namely Augmented Dickey–Fuller (1979, 1981) (ADF), Phillips–Perron (1987) (PP) and Kwiatkowski–Phillips–Schmidt–Shin (1992) (KPSS). ADF and PP tests have a null hypothesis stating that the series in question has a unit root against the alternative that it does not. The null of KPSS, on the other hand, states that the variable is stationary. In the literature, KPSS is sometimes used to verify the results of ADF and PP because their probability of rejecting the false hypothesis is low. The unit root test results, shown in Table 2, indicate that there is a unit root in all the level series but not in the first-difference series. Therefore, we conclude that each series follows an I (1) process.

Perron (1989) states that conventional unit root tests are subject to misspecification bias and size distortion when the series involved undergo structural breaks, which leads to a spurious acceptance of the unit root hypothesis. To capture a possible structural break during the sample periods, the Zivot and Andrews (1992) test is used, which treats the presence of a

[3] EUA futures contracts for delivery at maturities from December 2010 to 2012 are not considered in this study because of their lack of liquidity. We also don't consider EUA futures contracts for December 2005 because EEX hasn't launched contract for delivery at this maturity. Indeed, EEX has started futures trading in EUAs on 4th October of 2005.

structural break in the series under investigation endogenously. Table 3 reports the minimum t-statistics from testing the stationarity assuming a shift in mean for the first differences of the four contracts traded on the three trading platforms. The results suggest that at 5% level of significance none of the estimated variables are stationary around a shift in the mean. The estimated breakpoint for EUA futures contracts for December 2006 and 2007 traded on the three trading platforms is in April 25, 2006. The estimated breakpoint for futures contracts for December 2008 traded on EEX is in June 30, 2006; however, it is in June 24, 2006 for those traded on ECX and Nordpool. The timing of these structural breaks is explained by their proximity to the announcement by some countries of their 2005 emissions data in April and May 2006, before the official deadline of May 15 fixed by the European Commission, indicating a generous attribution of quotas in their national allocation plan. The estimated breakpoint for futures contracts for December 2009 traded on EEX, ECX and Nordpool are in August 28, 2007; April 20, 2007 and April 23, 2007 respectively. The timing of these structural breaks can be related to the release of the 2006 emissions data on April 25, 2007 confirming that the EUA market is long.

3.2 Cointegration analysis

Having verified that all the variables are integrated of order one, the next step is to test for the existence of a cointegration relationship between the variables. As indicated, the basic idea behind cointegration is to test whether a linear combination of three individually non-stationary time series is itself stationary. The Johansen cointegration tests were performed for each type of futures contract traded in the three trading platforms and use an intercept but no trend. A dummy variable taking the value of 1 in the period ranges from April 24, 2006 to May 15, 2006 and zero otherwise is created in order to take into account the crash of EUA prices in spring 2006 which is mainly explained by the release of official carbon emissions report that showed an oversupply of allowances over the 2005 reporting period. We determine the optimum lag length for Johansen cointegration test based on minimum Akaike information criterion (AIC) through unconstrained vector autoregression (VAR) estimation (1 lag interval in first differences for each series in the four equations). The lag length is further validated by tests for normality and absence of serial correlation in the residuals in VAR to make sure that none of them violates the standard assumptions of the model. Table 4 displays the outcome of these tests. The analysis indicates that the futures price series for each contract traded on the three trading platforms have more than one cointegrating relationship (in Table 4, $H_0 : r = 0$ and $r \leq 1$ is rejected at 5% level in the four equations). The evidence of cointegration has several important consequences. First, it eliminates spurious correlations, and suggests at least a unique channel for Granger causality test (either uni-directional or bi-directional). Second, the long-run relationship is incorporated by including the lagged error correction term (ECT) in the relevant VECM model. Third, it shows a high degree of price transmission and therefore a reasonable degree of efficiency, suggesting that future fluctuations of prices in one market can be determined or predicted to some extent using a part of the information set provided by the other market prices. Fourth, the benefits of international portfolio diversification are likely reduced since the prices seem to exhibit the same behavior in the long-run. Finally, cointegrated EUA prices converge towards a common long-run equilibrium path, as environmental policies tend to be coordinated.

3.3 Error-correction models and Granger causality tests

The presence of cointegration Granger causality requires the inclusion of an error-correction term (ECT) in the stationary model in order to capture the short-run deviations of series from their long-run equilibrium path. The Granger causality tests results according to the test method discussed above are reported in Table 5. Before confirming the results, the models are subjected to a battery of diagnostic tests for normality (Jarque–Bera), serial correlation (LM), and parameter instability (CUSUM and CUSUM square). The error-correction terms in all the models are also checked for unit roots. It can be seen that the lagged error-correction terms, in each VECM are significant with a negative sign, suggesting that all variables dynamically interact to return to the long-run equilibrium whenever there is a deviation from the cointegrating relationship and therefore transmission efficiency takes place. The estimated ECT coefficients are small, suggesting that the adjustment process is slow for each VECM.

In the case of EUA futures contracts for December 2006 and 2007, we see that, in the short run dynamics, ECX and EEX are significant in Nordpool equation, but none are significant in ECX equation. This suggests that there is only uni-directional Granger causality running from ECX and EEX to Nordpool in the short run. We also note that there is only uni-directional short-run Granger causality running from ECX to EEX. When we consider the F-test statistics for the joint significance of the sum of the lags of the explanatory variable and the error-correction term, we find that Nordpool is affected by both EEX and ECX. Conversely, ECX is not Granger caused neither by EEX nor by Nordpool. This indicates that there exists only uni-directional Granger causality running from ECX and EEX to Nordpool in the long-run. We also observe a uni-directional long-run Granger causality running from ECX to EEX.

In the case of EUA futures contracts for December 2008, we observe bi-directional short-run Granger causality between EEX and Nordpool. We also find a uni-directional causality running from ECX to EEX in the short-run. However, ECX and Nordpool fail to demonstrate at least uni-directional short-run causality. Considering the joint F-test, EEX and Nordpool are confirmed of Granger endogeneity. This implies that there exists bi-directional long-run Granger causality between EEX and Nordpool. We also find that there is only a uni-directional Granger causality running from ECX to EEX in the long-run. However, ECX and Nordpool fail to demonstrate at least uni-directional Granger causality in the long run.

In the case of EUA futures contracts for December 2009, we observe that, in the short run dynamics, EEX and Nordpool are symmetrically significant in respective equations. This suggests the existence of bi-directional causality in the short run. We also find a uni-directional short run Granger causality running from Nordpool and EEX to ECX, but not vice versa. Considering the joint F-test, we find bi-directional long-run Granger causality between Nordpool and EEX. We also observe a uni-directional long run Granger causality running from Nordpool to ECX, but not vice versa. However, EEX and ECX fail to demonstrate at least unidirectional causality in the long run.

Considering the four EUA futures contracts as a whole, the results seem to indicate a dominant role for the ECX in European carbon markets. This can be explained by the high trading volume over the sample periods in this trading platform. However, it is noteworthy that we also detect a feedback between Nordpool and EEX since there is bi-directional Granger causality in both short-and long run for EUA futures contracts for December 2008 and 2009 and a uni-directional short-run Granger causality running from EEX to Nordpool for EUA futures contracts for December 2006 and 2007. This suggests that Nordpool responds to innovations originating in EEX and it can be explained by the reasonable liquidity in Nordpool. Indeed, Nordpool was the first trading platform which started to trade EUA futures contracts. The evidence suggests that European carbon markets exhibited a reasonable degree of efficiency in both short-and long-run.

4. Conclusion

The aim of this paper was to investigate the causal relationship among European carbon markets. To this end, causality tests have been performed using recent techniques in the time series literature and adapted in a framework where both traditional and additional channels of causality could be exposed. In summary, time series properties of the data have been analyzed using unit root with and without structural break and cointegration tests before applying Granger-causality tests and vector error-correction models were performed to test for the direction of Granger-causality. Our empirical results clearly show that the European Climate Exchange (ECX) is more influential in the information transmission process since there is causality from ECX to all others for the four futures contracts. We also detect that European Energy Exchange (EEX) affects ECX. In light of our findings, we suggest that European carbon markets are reasonably efficient.

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Table 1. Summary statistics of EUA futures prices

	EEX	ECX	NORDPOOL
DEC 06			
Mean	19.354	19.308	19.34
Maximum	30.53	30.45	30.5
Minimum	8.08	8.18	7.95
Std. Dev.	5.897	5.887	5.894
Skewness	0.031	0.038	0.037
Kurtosis	1.911	1.919	1.914
Jarque-Bera	14.979 (0.000)	14.788 (0.001)	14.909 (0.000)
Observations	302	302	302
DEC 07			
Mean	11.262	11.236	11.249
Maximum	31.6	31.5	31.6
Minimum	0.05	0.04	0.03
Std. Dev.	10.389	10.363	10.38
Skewness	0.312	0.312	0.313
Kurtosis	1.636	1.637	1.639
Jarque-Bera	52.791 (0.000)	52.714 (0.000)	52.615 (0.000)
Observations	563	563	563
DEC 08			
Mean	20.937	20.925	20.944
Maximum	32.03	32.25	32.6
Minimum	12.22	12.25	12.15
Std. Dev.	3.639	3.637	3.658
Skewness	0.183	0.191	0.194
Kurtosis	2.666	2.669	2.716
Jarque-Bera	7.997 (0.018)	8.324 (0.015)	7.535 (0.023)
Observations	782	782	782
DEC 09			
Mean	21.577	21.555	21.537
Maximum	32.78	32.9	33.1
Minimum	12.72	12.8	12.65
Std. Dev.	3.738	3.737	3.733
Skewness	0.183	0.188	0.149
Kurtosis	2.608	2.588	2.558
Jarque-Bera	9.12 (0.01)	(0.007)	(0.011)
Observations	761	761	761

Notes: Jarque-Bera statistic is used to test whether or not the series resemble normal distribution. P-values are in parentheses.

Table 2. Unit root tests

Series	ADF				PP		KPSS	
	Level		First difference		Level	First Difference	Level	First Difference
	Lag	Test statistic	Lag	Test statistic	Test statistic	Test statistic	Test Statistic	Test statistic
	DEC 06							
EEX	3	-1.059 (1)	3	-7.248** (1)	-1.079 (1)	-12.250** (1)	1.415** (2)	0.085 (2)
ECX	2	-1.114 (1)	1	-10.823** (1)	-1.082 (1)	-13.058** (1)	1.409** (2)	0.082 (2)
Nordpool	3	-1.055 (1)	3	-7.030** (1)	-1.083 (1)	-12.031** (1)	1.410** (2)	0.087 (2)
	DEC 07							
EEX	3	-1.497 (1)	3	-9.891** (1)	-1.53 (1)	-16.639** (1)	2.828** (2)	0.062 (2)
ECX	2	-1.580 (1)	2	-10.889** (1)	-1.593 (1)	-17.616** (1)	2.828** (2)	0.061 (2)
Nordpool	3	-1.485 (1)	3	-9.566** (1)	-1.539 (1)	-17.046** (1)	2.828** (2)	0.062 (2)
	DEC 08							
EEX	2	-0.645 (1)	1	-18.011** (1)	-0.565 (1)	-22.451** (1)	0.644** (2)	0.097 (2)
ECX	2	-0.653 (1)	1	-18.110** (1)	-0568 (1)	-24.918 ** (1)	0.646** (2)	0.088 (2)
Nordpool	2	-0.656 (1)	1	-18.095** (1)	-0.606 (1)	-24.091** (1)	0.644** (2)	0.072 (2)
	DEC 09							
EEX	2	-0.589 (1)	1	-17.672** (1)	-0.575 (1)	-22.503** (1)	0.684** (2)	0.094 (2)
ECX	2	-0.610 (1)	1	-17.851** (1)	-0.586 (1)	-24.821** (1)	0.695** (2)	0.084 (2)
Nordpool	2	0.607 (1)	1	-17.714** (1)	-0.612 (1)	-22.746** (1)	0.711** (2)	0.080 (2)

Notes: ADF: Augmented Dickey-Fuller test. PP: Phillips-Perron test. KPSS: Kwiatkowski–Phillips–Schmidt–Shin. (1): Model without constant or deterministic trend. (2): Model with constant, without deterministic trend. The optimal lag structure is determined by the Durbin Watson test. If the regression model includes lagged dependent variables as explanatory variables, we use the Durbin's h test. ADF and PP critical values are taken from MacKinnon (1991). KPSS critical values are sourced from Kwiatkowski *et al*. (1992). All null hypothesis except KPSS are unit root; while, in KPSS null is stationarity. ** denotes rejection of the null hypothesis at the 5% significance level.

Table 3. Zivot-Andrews minimum *t*-statistics

Variables	*t*-statistic	Period
	DEC06	
EEX	-4.969 (3)	25/04/2006
ECX	-5.054 (4)	25/04/2006
Nordpool	-4.925 (4)	25/04/2006
	DEC07	
EEX	-3.913 (3)	25/04/2006
ECX	-3.957 (4)	25/04/2006
Nordpool	-3.871 (4)	25/04/2006
	DEC08	
EEX	-3.146 (1)	30/06/2006
ECX	-3.660 (4)	24/06/2006
Nordpool	-3.458 (1)	24/06/2006
	DEC09	
EEX	-3.067 (1)	28/08/2007
ECX	-2.916 (1)	20/04/2007
Nordpool	-3.135 (1)	23/04/2007

Notes: All *t*-statistics estimated from a break in intercept model. Values in parentheses are lag length used in the test for each series. Critical values are those reported in Zivot and Andrews (1992).

Table 4. Results of the Johansen cointegration analysis

Equation	Rank *r*	Eigen value	Max. Eigen Value statistic	Trace test statistic	EEX	ECX	Nordpool	Constant
DEC06	0	0.452	180. 600**	320.228**	1	-1.120***	0.118*	0.007
	≤ 1	0.370	138.475**	139.628**		[-17.658]	[1.858]	[0.174]
	≤ 2	0.004	1.153	1.153				
DEC07	0	0.432	317.238**	549.226**	1	-1.407***	0.404**	0.018
	≤ 1	0.337	231.158**	231.987**		[-23.017]	[6.621]	[1.313]
	≤ 2	0.001	0.829	0.829				
DEC08	0	0.371	362.716**	550.968**	1	-1.471***	0.468***	0.056
	≤ 1	0.209	183.006**	188.252**		[-27.688]	[8.851]	[0.902]
	≤ 2	0.008	5.246	5.246				
DEC09	0	0.367	347.352**	490.520**	1	-1.359***	0.359***	-0.013
	≤ 1	0.166	137.493**	143.169**		[-32.455]	[8.565]	[-0.211]
	≤ 2	0.007	5.676	5.676				

Notes: Rank *r* expresses the number of cointegrating equation according to each tested hypothesis. No restriction is imposed in the cointegration test. The lag length has been chosen based on minimum Akaike information criterion (1 lag interval in first differences for each series). Critical values were taken from McKinnon *et al.* (1999). The right-hand of the table shows the normalised coefficients of the cointegrating equations, with t-statistics in brackets. *, ** and *** denote significance at 10%, 5% and 1% level respectively.

Table 5. Granger causality tests

Variables	Short-run effects			ECT effect	Joint short-and long-run effects (F-statistics)		
	Δ EEX	Δ ECX	Δ Nordpool		ECT& Δ EEX	ECT& Δ ECX	ECT& Δ Nordpool
DEC 06							
Δ EEX	-	-0.382** [-2.336]	0.013 [0.080]	-1.770*** [-9.640]	-	5.548** (0.019)	0.006 (0.936)
Δ ECX	0.366 [1.314]	-	0.026 [0.135]	-1.017*** [-4.585]	1.727 (0.189)	-	0.018 (0.893)
Δ Nordpool	0.838*** [3.828]	-0.323** [-2.100]	-	-1.814*** [-10.398]	14.658*** (0.000)	4.410** (0.036)	-
DEC 07							
Δ EEX	-	-0.219** [-1.981]	0.003 [0.03]	-1.193*** [-11.837]	-	3.924** (0.048)	0.001 (0.974)
Δ ECX	0.220 [1.207]	-	-0.045 [-0.328]	-0.608*** [-4.938]	1.457 (0.227)	-	0.107 (0.743)
Δ Nordpool	0.732*** [5.036]	-0.198* [-1.837]	-	-1.255*** [-12.792]	25.360*** (0.000)	3.373* (0.066)	-
DEC 08							
Δ EEX	-	-0.152* [-1.172]	0.156** [2.119]	-0.743*** [-9.399]	-	2.953* (0.086)	4.492** (0.034)
Δ ECX	0.115 [0.939]	-	0.095 [1.118]	-0.149* [-1.648]	0.881 (0.348)	-	1.251 (0.263)
Δ Nordpool	0.448*** [4.127]	-0.108 [-1.198]	-	-0.916*** [-11.766]	17.035*** (0.000)	1.435 (0.231)	-
DEC 09							
Δ EEX	-	-0.126 [-1.399]	0.257*** [4.008]	-0.795*** [-9.182]	-	1.956 (0.162)	16.064*** (0.000)
Δ ECX	0.008* [0.068]	-	0.193*** [2.600]	-0.190* [-1.893]	0.005 (0.946)	-	6.761** (0.009)
Δ Nordpool	0.286*** [2.939]	-0.074 [-0.839]	-	-0.943*** [-11.183]	8.636** (0.003)	0.704 (0.402)	-

Notes: Δ is the difference operator. *, **and *** indicate significance level at 10%, 5% and 1% respectively. Numbers in parentheses denote p-values. Numbers in brackets indicate t-statistics. The lag length for each model is determined as 1 according to AIC.