An Analysis of the EU Emissions Trading Scheme

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Abstract

The European Union's Emissions Trading Scheme (ETS) is the key policy instrument of the European Commission's Climate Change Program aimed at reducing greenhouse gas emissions to eight percent below 1990 levels by 2012. A critically important element of the EU ETS is the establishment of a market determined price for EU allowances. This article examines the extent to which several theoretically founded factors including, economic growth, energy prices and weather conditions determine the expected prices of the European Union CO_2 allowances during the 2005 through to the 2009 period. The novel aspect of our study is that we examine the heavily traded futures instruments that have an expiry date in Phase 2 of the EU ETS. Our study adopts both static and recursive versions of the Johansen multivariate cointegration likelihood ratio test as well as a variation on this test with a view to controlling for time varying volatility effects. Our results are indicative of a new pricing regime emerging in Phase 2 of the market and point to a maturing market driven by the fundamentals. These results are valuable both for traders of EU allowances and for those policy makers seeking to improve the design of the European Union ETS.

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1 Introduction

In January 2005 the European Union (EU) emissions trading scheme (ETS) was introduced formally. The scheme has been instigated as part of the EU agreement to cut worldwide emissions of carbon dioxide (CO_2) within the Kyoto Protocol. Under the Kyoto agreement, the EU has committed to reduce greenhouse gas (GHG) emissions by eight percent (relative to 1990 levels) by 2008-2012. The scheme issues a restricted amount of emission allowances to firms on an annual basis. At the end of the year firms must hold the required amount of emission permits to meet their emissions of CO_2 over the previous year.[1] The ETS allows firms to trade the amount of emission permits that they hold and as a result has applied a market value to this externality. In the EU ETS context the first phase of trading was 2005-2007 and the second one, which coincides with the first compliance period of the Kyoto Protocol, is 2008-2012. The third European trading phase will commence in 2013. Non-compliance with the commitments will result in a penalty of 40 (100) euros per tonne of CO_2 produced without allowances for the first (second) commitment period. The aim of the ETS is that this cost will encourage firms to reduce their emissions. Paolella and Taschini (2008) highlight that the ultimate aim of this scheme (as well as the US CAAA-Title IV scheme) must be to create an environment where there is scarcity of allowances which will lead to an upward trend in prices. As a result we might expect to see mean reversion around an upward trend. However, there has been a considerable amount of uncertainty associated with the price of CO_2 emissions over its short life to date.

Concomitant to the recent dramatic fall in allowance prices (spot falling from 30 euro in the summer 2008 to just under 10 euro in the spring 2009) has been growing calls for intervention by the European Commission into the market. Those calling for intervention see the low prices as incentivising higher rather than lower carbon based technology.[2] Any intervention is likely to seriously distort the market and may impede investment in low carbon technology in the future. As noted by Lowrey (2006) a centrally important element of the EU ETS is the establishment of a market determined price for EU allowances. In this article, we take account of market uncertainty and examine the extent of the emergence of an equilibrium relationship between the expectation of EU allowance prices and a set of theoretical determinants, including economic growth, energy prices and weather conditions.[3] Unlike the vast majority of previous work in the area, we take account of both structural and time series properties in examining the behaviour of EU allowance prices. Taking account of both structural and time series properties will indicate whether prices, although currently low, are determined by a stable relationship. The contributions of the paper to the empirical literature on modeling carbon emissions is threefold. Our study is the first known study to investigate the long-run relationship between theoretically ac-

[1] A report must be submitted to verify the emissions in any year by the 31^{st} March of the following year.

[2] Mark Lewis, director of global carbon research at Deutsche Bank, proposed (6 February 2009) to establish a reserve price for EU emissions allowances (EUAs) to avoid a price collapse in the third phase of the EU ETS, which starts in 2013.

[3] See Convery (2009) and Springer (2003) for a survey of the literature.

cepted determinants and EU carbon allowances using a battery of cointegration procedures. Cointegration is a powerful econometric approach which can indicate whether a stable relationship exists and whether the behaviour of the variables binding this relationship are consistent with economic theory. Secondly, besides taking account of the potential cointegrating relations, we also address the empirical finding of time varying volatility in the EU ETS and augment time varying volatility into the cointegration tests. Finally, given the relatively small sample of data and the considerable uncertainty, in particular during the pilot phase, we examine the extent of the evolving long-run relationships adopting a recursive cointegration approach.

Given the relative paucity of data available and consistent with the previous literature, our analysis will adopt data at a relatively high frequency, daily data. The full sample of data covers the period 2005 to 2009 and so incorporates both Phase 1 and Phase 2 data. The expiration on our futures contracts is December 2008 and December 2009. Unlike the vast majority of the previous studies, our focus will be on futures rather than spot contracts. The justification for examining futures is due to the greater volumes being traded on these contracts (see Mansanet-Bataller and Pardo, 2008).[4] These instruments were not exposed to the dramatic structural breaks that have been previously highlighted in the literature and so results in an additional advantage of adopting the futures based analysis. We will adopt the cointegration procedure to identify the existence of a long-run relationship. We also adopt a number of identifying restrictions to further refine our model. Finally, we also carry out a number of sensitivity tests which take account of time varying volatility and the structural breaks in the data. Our results are consistent with previous work in that we find considerable evidence of uncertainty for EU allowances and the range of determinants (see, Paolella and Taschini (2008) and Benz and Trück (2009)). Although, there have been calls for intervention in the market, our results indicate that for a Phase 2 sample a stable relationship has formed between EU allowances and other determinants. A range of cointegration test results report theoretically consistent relationships have emerged in Phase 2 of the ETS. There is no evidence of this stable relationship occurring for the Phase 1 sample.[5] Our empirical results are also consistent when we take account of the time varying volatility in the data.[6]

The remainder of this article is structured as follows, section 2 discusses the performance of Phase 1 and the implications for Phase 2, along with a detailed analysis of the theoretical and empirical determinants. Section 3 describes the methodologies being adopted, while section 4 presents the data and empirical results. Finally, concluding remarks are presented in section 5.

[4] Mansanet-Bataller and Pardo (2008) report cumulative volumes traded in the different European Carbon markets since the start of the trade in each market until January 2008. The volumes traded in spot is 4%, futures 76% and over the counter (OTC) 20%.

[5] For the remainder of the paper, Phase 1 refers to the Phase 1 sample and Phase 2 refers to the Phase 2 sample.

[6] We find no evidence of a structural break in the data. This is not particularly surprising given that our analysis covers futures contracts that expire in Phase 2.

2 Phase I Empirical Evidence & the Implications for Phase II

A number of studies have examined the performance of Phase 1 EU ETS, mainly using data of a daily frequency, given the paucity of data. Recent studies include Paolella and Taschini (2008), Daskalakis *et al.* (2009) and Benz and Trück (2009) examine the time series properties of a range of different EU ETS instruments.[7] For example, Benz and Trück (2009) adopt a pure time series approach and take account of the non-normality associated with the EU allowance returns and find evidence of regime switching.[8] Unlike the previous cited studies which adopted a pure time series approach, Redmond and Convery (2006) and Alberola *et al.* (2008) examine the behaviour of the price of carbon in relation to energy commodities, meteorological factors and a number of other variables.[9] Redmond and Convery (2006) include for example dummy variables to take account of policy and regulatory issues. While, Alberola *et al.* (2008) examine the extent of extreme temperature and find evidence that extremely cold temperatures have a statistically significant impact but only at a sub-sample setting.

The empirical studies to date have highlighted the difficulties associated with Phase 1 (pilot phase). In particular there was considerable uncertainty and volatility associated with the market price of EUA's. In April 2006, coincident to the unofficial release of the 2005 emissions data by some of the EU member states the price of EUAs collapsed. EU ETS spot prices had reached a high of 30.50 euro prior to April 2006. Following the official release by the EU commission on the 15^{th} May 2006, showing a larger than expected surplus in the market, the spot price fell to 15.63 euro on the 17^{th} May 2006. Given that banking EUA's was prohibited between phases, the price eventually converged to close to zero at the end of Phase 1. As well as the April 2006 break, Alberola *et al.* (2008) also highlight a break in October 2006. This break relates to an announcement by the European Commission (EC) of considerably stricter policy in relation to the allocation of permits for Phase 2.[10] Overall for Phase 1, it would appear that the cap placed on emissions was far too lax and so downward pressure on the spot and futures (those expiring in Phase 1) price continued.

As has been highlighted by a number of authors including, Christiansen *et al.* (2005), Bunn and Fezzi (2007), Redmond and Convery (2006) and Alberola *et al.* (2008), energy prices are a key driver of carbon prices. Large installations, in particular power plants, are likely to switch between various forms of energy depending on the associated cost. In particular, power plants pay close attention to the profits from producing electricity depending on whether the input is coal (profits are referred to as dark spread) or gas

[7] Paolella and Taschini (2008) examine both SO_2 (in the US) and CO_2 (EU) spot price dynamics.

[8] The only study that has addressed the market microstructure issues for this market has been Benz and Hengelbrock (2009) and Bredin, Hyde and Muckley (2009). Both studies find evidence of an increase in market liquidity for Phase 2 expiring futures contracts.

[9] Alberola *et al.* (2009a and 2009b) have also examined the role of market structure and industrial sectors.

[10] On 26 October 2006, the EC announced a stricter policy for national allocation plans (NAP) for Phase 2 of the EU ETS.

(profits are referred to as spark spread). Given the costs of CO_2 emissions, dark and spark spreads are adjusted further to take account of the additional cost and are referred to as clean dark and spark spreads. Along with energy prices, weather conditions are considered a theoretically important variable in determining the price of carbon. Studies that have incorporated weather conditions in explaining movements in Phase 1 EU ETS include the Redmond and Convery (2006), Mansanet-Bataller *et al.* (2007) and Alberola *et al.* (2008). In all cases the authors take account of temperature extremes and the likely effects with some evidence to suggest the importance empirically of these variables.[11]

Clearly a number of difficulties remain. These include the fact that the cap was only aimed at large emitters from the power and heat generation industries and in selected energy intensive industries.[12] As has been highlighted earlier the over allocation of allowances has been problematic. The national allocation plans (NAP) submitted by member states to the European Commission were not reviewed in Phase 1 and these were distributed free of charge by member states to the emitting firms.[13]

2.1 EU Allowance Determinants

Our analysis of the determinants of CO_2 prices draws on the significant number of recent studies examining the empirical relationship between the EU allowance (EUA) prices and its fundamentals. An important theoretical review of the material is included in Springer (2003), while Christiansen *et al.* (2005) identifies the key drivers of EUA prices as economic growth, energy prices and weather conditions. The empirical literature on the estimation of the determinants of EUAs is based on the following long-run relationship where we take account of economic factors, energy factors and climate factors (see Mansanet-Bataller *et al.* (2007) and Alberola *et al.* (2008));

$$EUA_t = \alpha_0 + \alpha_1 y_t + \alpha_2 p_t + \alpha_3 T_t \tag{1}$$

where EUA_t, y_t, p_t, and T_t stand for EUA futures price, income, energy prices and finally temperature. All series are in logs except temperature.

Christiansen *et al.* (2005) has highlighted the role of economic growth as a determinant of EUA prices, with higher economic growth leading to a rise in the EUA price. We include two proxies, industrial production and equity price movements. While industrial production would be a standard measure of economic growth, a potential difficulty here is that it is only available at a monthly frequency. A solution which we adopt here is to interpolate

[11] Redmond and Convery (2006) find no evidence of a statistically significant weather effect, while Alberola *et al.* (2008) do find evidence but only for certain sub-samples of Phase 1.

[12] The European Commission (2005) has estimated that these installations account for 45% of CO_2 emissions. Airlines will be included in the next phase of the EU ETS, from 2013-2020.

[13] Member states were allowed to auction up to 5% of their total allowance allocation in Phase 1 (Convery and Redmond, 2007). To date Denmark, Hungary, Ireland and Lithuania have used auction provisions.

the data using a piecewise cubic spline methodology.[14] A euro zone equity (futures) index is also considered as a measure of economic conditions. The motivation for including this variable is that it offers an up-to-date indicator of expectations on both financial and economic conditions at the required daily frequency. Further, given the financial nature of the underlying asset, we consider including such a proxy informative.

As has been highlighted by a number of theoretical studies energy prices are highly influential factors on CO_2 prices (see Burniaux (2000), Ciorba *et al.* (2001), Sijm *et al.* (2000) and van der Mensbrugghe (1998)). Studies that have included energy variables in a similar context to the current study include Redmond and Convery (2006), Mansanet-Bataller *et al.* (2007) and Alberola *et al.* (2008). Our proxy for energy prices is the brent crude futures oil price and we would expect a positive relationship between oil price movements and the EUA price. In order to take account of the abatement options for large installations and the impact of relative fuel prices and consistent with Alberola *et al.* (2008) we also include two specific spread terms, clean dark and spark spreads. The clean dark (spark) spread represents the difference between the price of electricity at peak hours and the price of coal (gas) used to generate that electricity, corrected for the energy output of the coal (gas) plant. Hence, both coal and gas prices are also implicity included in our analysis. A negative relation between the EUA price and clean spark spreads (CSS) is expected to arise as greater profitability from generating electricity from natural gas, *ceteris paribus*, would result in switching to natural gas fueled electricity generation and hence a short run abatement with respect to CO_2 emissions.[15] EUA prices are likely to decline following the fall in demand. Similarly, the opposite relation is expected to hold between EUAs and clean dark spreads (CDS).

Finally, while Considine (2000) and Davis *et al.* (2002) document the significant impact of temperature on the intensity of carbon emissions in the United States, Ciorba *et al.* (2001) highlight temperature as being one the most influential factors in determining the CO_2 price. Empirical studies have examined a range of temperature proxies find that variables such as mean temperature do not have a statistically significant influence on CO_2 prices, while extreme temperatures (temperature deviations from seasonal averages) consistently have a statistically significant positive influence (see Mansanet-Bataller *et al.* (2007) and Alberola *et al.* (2008)). The motivation here is that energy use and so emissions are higher during extreme weather (hot or cold) than during moderate weather (see Moral-Carcedo and Viciens-Otero, 2005).

[14] A second possible issue of note is that EUA prices are likely to be predominantly affected by growth in those sectors covered by the EU ETS. In particular, a rise in industrial production from those sectors covered by the EU ETS is likely to result in a rise in the purchases of EUAs. See, Alberola *et al.* (2009b) for an example of a study that investigates the income effect on EUAs using disaggregated industrial production.

[15] Gas fired energy plants emit considerably lower CO_2 pollutants compared to coal fired energy plants.

3 Methodology

We examine the development of cointegration relations in a system containing the EU allowance futures contracts, alongside the variety of other theoretically founded factors described in Section 2.1. Our statistical testing procedures include conventional versions of the Engle and Granger (1987) and Johansen (1988) cointegration tests alongside a modified Johansen (1988) cointegration test which is an adaptation of the methodology provided by Gannon (1996). The Engle and Granger (1987) and Johansen (1988) methodologies are well known in the literature and so we provide only a brief description of these statistical tests.

The methodology is applied to examine the over all behaviour of the system and its evolution with respect to the criterion of cointegration. We perform static analysis including hypotheses tests focused on the significance of the EU allowance futures contracts in the estimated cointegrating vectors. In addition, to gain an insight into the evolution of the system, we recursively estimate the outlined tests for cointegration. In particular, with regard to our recursive methodology, we perform the tests over the initial 250 observations and subsequently repeat the testing procedure over an extended window of data, where the window is extended by a single observation prior to each incremental estimate of the test statistics.

3.1 Engle-Granger (1987) cointegration test

The Engle-Granger (1987) methodology requires initial testing of the order of integration of the variables concerned. The augmented Dickey Fuller test is adopted, with critical values from Dickey and Fuller (1979). In the instance where the variables are found to be integrated of order 1, $I(1)$, it is adequate to estimate the long-run relationship described in equation (2). If the variables, $x_{i,t}$ where $i = 1...k$, are actually cointegrated, then the ordinary least squares regression yields a super consistent estimator of the cointegrating parameters, $\beta_0...\beta_k$. This possibility is investigated by means of equation (3) to establish if the deviations from long-run equilibrium are stationary. MacKinnon (1991) provides appropriate critical values.

$$y_t = \beta_0 + \beta_1 x_{1,t} + ... + \beta_k x_{k,t} + e_t \tag{2}$$

$$\Delta \hat{e}_t = a_1 \hat{e}_{t-1} + \sum_{i=1}^{n} a_{i+1} \Delta \hat{e}_{t-1} + \epsilon \tag{3}$$

Unfortunately, the Engle-Granger (1987) technique does exhibit several important defects. In particular, it is sensitive, in finite samples, to the choice of variable for normalisation. Even in the simple two variable setting, a potential drawback of the Engle-Granger approach is that there could be a simultaneous equations bias if the causality between the variables runs in both directions. The problem is clearly compounded using three or more variables given that any of the variables can be selected as a left hand side variable. In addition, the methodology precludes the possibility of estimating multiple cointegrating vectors. Finally, another defect of the Engle-Granger (1987) procedure is that it is a two-step procedure, this imparts invidious implications for the efficiency of the testing procedure. Fortunately, the Johansen (1988) procedure is a maximum likelihood estimator which circumvents the requirement for a two-step estimator. The Johansen approach allows all variables to be endogenous and, as a result, can estimate and test for multiple cointegrating relations.

3.2 Johansen (1988) cointegration test

The Johansen (1988) cointegration testing framework hinges on the relationship between the rank of a matrix and its characteristic roots. In the first instance, the maximum likelihood estimation of the vector error correction model (henceforth VECM), as outlined in equations (4), (5) and (6) is performed. Specifically, the test for cointegration involves the performance of likelihood ratio statistical hypotheses tests regarding the rank of the long-run information matrix, π. These statistical tests provide an estimate of the number of characteristic roots that are insignificantly different to unity.

$$\begin{aligned} \Delta x_t &= \pi x_{t-1} + \sum_{i=1}^{k-1} \pi_i \Delta x_{t-1} + \varepsilon_t && (4) \\ \pi &= \sum_{i=1}^{k} \pi_i - I && (5) \\ \pi_i &= -\sum_{i+1}^{k} \pi_j, (i = 1, ..., k-1) && (6) \end{aligned}$$

In order to ascertain the rank of the long-run information matrix, π, a set of so-called trace statistics, $\lambda_{trace}(r)$, is estimated.

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{n} ln(1 - \hat{\lambda}_i) \quad (7)$$

In this formulation, T refers to the number of available observations. The symbol, $\hat{\lambda}_i$, denotes the estimated value of the i^{th} characteristic root, or equivalently an eigen value in the long-run information matrix. The $\lambda_{trace}(r)$ statistic assesses the null hypothesis that the number of cointegrating vectors is less than or equal to r against a general alternative hypothesis. Typically, results are generated for each possible value of r. Osterwald-Lenum (1992) provide critical values.

3.3 Modified Johansen (1988) cointegration test

The literature in the area of cointegration testing, in the context of ARCH effects, is in its infancy. The theoretical literature (see Lee and Tse (1996), Silvapulle and Podivinsky (2000) and Hoglund and Ostermark (2003)) indicates that ARCH effects aggrandise the size of the Johansen (1988) cointegration test. For example, Lee and Tse (1996) indicate that while the Johansen (1988) cointegration test tends to overreject the null hypothesis of no cointegration in favour of finding cointegration, the problem is generally not very serious. Silvapulle and Podivinsky (2000) report similar results. In contrast, Hoglund and Ostermark (2003) conclude that the eigenvalues of the long run information matrix for the Johansen (1988) cointegration test are highly sensitive to conditional heteroskedasticity and that therefore this multivariate statistic is only reliable in the context of homoskedastic processes. This latter finding, regarding the size of the cointegration test, becomes increasingly pronounced the more integrated the ARCH process considered. That said, these contributions pertain to low dimensional systems and, as a result, are of limited empirical relevance. For example, empirical contributions (see Alexakis and Apergis (1996), Gannon (1996) and Pan *et al.* (1999)), across a wider range of system dimensions, tend to indicate that these ARCH effects and their variants exert a significant and deleterious impact on the statistical test's power properties. Specifically, the aforementioned empirical literature identifies significant gains in statistical power once ARCH effects are controlled, when testing for cointegration, using the Johansen (1988) technique.

As a result of these contributions to the literature, a modified Johansen (1988) testing procedure is estimated with a view to mitigating for the deleterious implications of heteroskedasticity effects on the estimation of the rank of the long run information matrix in a specified VECM. Specifically, following Gannon (1996) and Pan *et al.* (1999), we adopt a modified test for common roots in which we account for heteroskedasticity effects in the correlating combinations of residuals. Consider again the m-dimensional VECM outlined in equations (4), (5) and (6).

The residuals, ϵ_t, are assumed independent normally distributed m-dimensional with mean zero and variance, Ω. The parameters $(\pi, \pi_1, ..., \pi_{k-1}, \Omega)$ are unrestricted and are estimated by maximum likelihood estimation. The x_t are vectors of series containing the EU allowance futures prices and the previously described theoretically founded determinants. Now, consider two auxiliary equations:

$$\Delta x_t = \sum_{i=1}^{k-1} \delta_{1i} \Delta x_{t-1} + r_{0t} \quad (8)$$

$$x_{t-1} = \sum_{i=1}^{k-1} \delta_{2i} \Delta x_{t-1} + r_{1t} \quad (9)$$

where δ_1 and δ_2 are estimated by ordinary least squares (see Johansen and Juselius, 1990 and Juselius, 1991). The vectors of series r_{0t} and r_{1t} contain the residuals from the auxiliary regressions. Note that the VECM, equations (4), (5) and (6) can now be reformulated as a two-stage estimation process:

$$r_{0t} = \alpha\beta' r_{1t} + \epsilon_t \tag{10}$$

The null hypothesis, H_0, that the components of x_t are cointegrated may be stated as

$$H_0 : \pi = \alpha\beta' \tag{11}$$

This implies that q = rank (π) < m. The rows of the (m×q) matrix β' are the distinct cointegrating vectors of x_t i.e., $\beta'(x_t)$ are I(0). The elements of α represent the loadings of each of the r cointegrating relations.

The canonical correlations can be estimated from the stacked residuals where the weights, $\omega_{1i}, ..., \omega_{pi}$ and $\kappa_{1i}, ..., \kappa_{pi}$ are canonical weights.

$$\begin{aligned}\hat{\nu_i} &= \hat{\omega_{1i}} r_{01i} + ... + \hat{\omega_{pi}} r_{0pi} && (12)\\ \hat{\eta_i} &= \hat{\kappa_{1i}} r_{11i} + ... + \hat{\kappa_{pi}} r_{1pi} && (13)\end{aligned}$$

Where r refers to the residuals from equations (8) and (9) and the subscript i refers to the i^{th} pair of canonical variables. Therefore, the variables $\hat{\nu_i}$ and $\hat{\eta_i}$ have a zero mean. These variates are constructed using canonical coefficients as weights, as outlined in equations (12) and (13). Finally, estimate, using the GARCH(1,1) equation specification equations for $\hat{\nu_i}$ and $\hat{\eta_i}$ for $i = 1, ..., q$

$$\begin{aligned}\hat{\nu_{it}} &= \rho_i \hat{\eta_{it}} + u_{it} && (14)\\ h_{it} &= Var(\hat{\nu_{it}}/\hat{\eta_{it}}) = \alpha_{i0} + \alpha_{i1} u_{t-1}^2 + \beta_{i1} h_{t-1} && (15)\end{aligned}$$

and compare the t-statistic for ρ with the tabulated values of the statistic given in Mackinnon (1991). Hence, an estimate of each eigenvalue, λ_i, is available as $\rho_i \approx \sqrt{\lambda_i}$. Neglecting heteroskedasticity effects provides inefficient estimates of the λ_i's while allowing for heteroskedasticity effects accounts for simultaneous volatility effects in the system. If there is common volatility across the series entering the system then linear combinations of the deviations from long-run paths will capture these common factors. The concern is that in neglecting to account for common volatility shocks, the test statistics may fail to reveal significant common roots. The test statistics are estimated from the procedure described in equations (12), (13), (14) and (15). We perform the two-stage procedure with and without accounting for GARCH(1,1) effects. When we do not acount for GARCH(1,1) *i.e.* when we do not adopt equation (15) in our estimation of the eigenvalues λ_i we use the Newey West (1987) procedure to control for heteroskedasticity which is critical when testing for the statistical significance of each eigenvalue, λ_i. This procedure provides an estimate, robust to heteroskedasticity effects, of the number of cointegrating vectors.

4 Data and Empirical Results

4.1 Data

Our data includes daily closing prices on EU allowance futures contracts, clean dark and spark energy spreads, Eurex Dow Jones EURO STOXX futures contracts, absolute deviations from mean temperatures and production as well as futures contracts on oil fossil fuel prices. All the examined data is plotted in Figure 1 and 2, with data construction and sources detailed in the appendix.

Besides having similar time series properties to other asset prices in our sample, the EUAs also behave in a similar fashion following the start of the significant economic and stock market downturn in the summer of 2008. Table 1 presents the summary statistics for the full sample, as well as EU ETS Phase 1 (from July 1 2005 to December 31 2007) and Phase 2 (January 1 2008 to March 2 2009). The last two columns report the results from the Phillips-Perron (1988) and Lee-Strazicich (2004) unit root tests.[16] Given the market uncertainty, the Lee-Strazicich (2004) test is considered as a further sensitivity test on the Phillips-Perron (1988) unit root test. The Lee-Strazicich (2004) unit root test indicates that each series contains a unit root in levels and is stationary in first differences, except in the instances of production and temperature. Although there are some minor exceptions in the full sample and Phase 1, the Phillips-Perron (1988) and Lee-Strazicich (2004) unit root test results indicate consistent results in relation to the level of integration.

The summary statistics indicate that the clean dark spreads (CDS) and the clean spark spreads (CSS) exhibit the greatest level of variance by a large order of magnitude. Given the recent findings, in particular Paolella and Taschini (2008) and Benz and Trück (2009), we also test for the level of autoregressive conditional heteroskedasticity (ARCH) using a lagrange multiplier (LM) test. According to a LM test the CDS differences appear homoscedastic, during all sample periods examined. According to the same statistical test the differenced CSS, oil, equity, production, temperatures and the EUAs series exhibit pronounced heteroskedasticity. In addition, the summary statistics indicate that the return distributions for several of the series are characterized by higher peakedness and thick tails relative to a normal distribution, particularly during Phase II of the EU ETS. Taken together, these summary statistics reflect the possibility of cointegration relations governing the system of variables examined. However, as has been discussed earlier in the paper, the heteroskedasticity that is apparent in our data may compromise the capacity of the classic Johansen (1988) cointegration test.

[16] Lee and Strazicich (2004) provide a lagrange multiplier unit root test that endogenously determines a structural break in intercept and trend

4.2 Empirical Results

The Engle and Granger (1987), the Johansen (1988) multivariate likelihood ratio cointegration approach and the Gannon (1996) modified cointegration tests are used to assess whether there are common forces driving the long-run movements of the full set of variables examined. The Engle-Granger approach is adopted purely as a preliminary investigation of the potential long-run relationships. Table 2 presents Engle-Granger (1987) style results. Specifically, it contains linear regression coefficients (all variables are logged, with the exception of temperature) corresponding to the full sample, Phase 1 and Phase 2 of the EU ETS. The Dickey-Fuller test statistic (last column) is not statistically significant in any of the samples and indicates the lack of cointegration in the full sample, Phase 1 and 2. However, it is noteworthy that there has been a marked heightening of the significance of elements in the cointegration equation during Phase 2, relative to the effects in Phase 1. The t-statistics are calculated using Newey-West (1987) adjusted standard errors.

The point coefficients give a preliminary indication as to the likely empirical relationships between EUA prices and the key variables of determination. While there is evidence of sign switching as we move from Phase 1 to Phase 2, the results do indicate a greater level of cointegration. In particular the Phase 2 coefficients are statistically significant, with theoretically consistent relationships, which are supportive of a developing equilibrium relationship. In Phase 2 the Engle-Granger results indicate that all variables, with the exception of productivity, have signs that are consistent with economic theory.[17] CDS and CSS represent the profitability for electricity generators depending on whether coal or gas is the principle input. While one would expect a negative (positive) relation between EUAs and CSSs (CDSs), this only emerges in the Phase 2. A negative relation between EUAs and CSSs is expected to arise as greater profitability from generating electricity from natural gas, *ceteris paribus*, would result in switching to natural gas fueled electricity generation and hence a short run abatement with respect to CO_2 emissions. EUA prices are likely to decline following the fall in demand. The theoretically consistent positive relationship is also found between CDS and EUAs for Phase 2. Oil prices are statistically significant in both phases with a coefficient close to unity. Finally, the temperature variable capturing unanticipated innovations in temperature (measured in absolute terms) is statistically significant in Phase 2. These latter effects may reflect increased demand for entitlements to emit carbon as a result of heightened demand for heating or air conditioning due to unexpected changes in temperatures.

Table 3 presents the normalised distinct cointegration equations and related hypotheses testing results, with respect to the Johansen (1988) estimation of the vector error correction model specification, corresponding to the full sample and Phases 1 and 2 of the EU ETS. The model specification (deterministic components and lag length) is inferred with respect to the Schwarz information criterion.[18] In Panel A, the normalized cointegrating equations are

[17] In particular the negative sign on production may be adversely affected by the dramatic decline over the 14 months representing Phase 2.

[18] These results are not presented here although they are available from the authors upon request.

presented alongside the t-statistics on the coefficients, while Panel B presents the hypotheses test results. The hypotheses that there are at most r (r = 0...4) distinct cointegrating vectors are examined, with the critical values sourced from Osterwald-Lenum (1992). As can be seen from hypothesis test (i) in Panel B, the Trace test results indicate that a long-run relationship exists over the full sample and for Phase 2.[19] The lack of cointegration for Phase 1, while marginal, is consistent with the pilot nature of the first phase and the considerable uncertainty associated with the market start-up. Consistent with the overall cointegration results, none of the coefficients in the normalized cointegrating vector are significant for Phase 1. However, we do find evidence of cointegration for Phase 2 and the signs on the coefficients for the normalized cointegrating vectors are all consistent with theory. In particular, oil, CSS and CDS are all statistically significant and have theoretically consistent signs. Our results clearly indicate evidence of an evolving long-run relationship which is consistent with theory. The remaining hypothesis test results in Panel B (hypotheses tests (ii), (iii) and (iv)) provide further support of a cointegrating relationship emerging in Phase 2. The hypotheses tests (ii), (iii) and (iv) correspond to null hypotheses of a zero loading coefficient on the disequilibrium error in the EUA equation, a zero coefficient on the EUA in the cointegrating equation and a joint null hypothesis with respect to these latter hypotheses, respectively. The results are robust to alterations of the deterministic components in the vector error correction model.[20]

As a result of the prevalence of ARCH effects in the data, a modified cointegration test with GARCH effects is performed. Table 4 presents the results.[21] The test statistics are estimated from the procedure described by Equations 14 and 15. The $\rho = 1$ test results are based on variates constructed from the weights for the maximum canonical correlation, whereas the second highest canonical correlation is used for $\rho = 2$, and so forth. Our results indicate evidence of an increase in cointegration as we move from Phase 1 to Phase 2 and are generally consistent with the Johansen results reported in table 3. However, a notable distinction of the Gannon results in contrast to those of the Johansen test is the greater number of cointegration vectors identified. The Gannon results identify two cointegrating vectors in Phase 2, while they identify one cointegration vector in the Full Sample period and during Phase 1. Taken in the context of the results concerning stationarity presented in table 1, the decline in the number of common stochastic trends implies that the remaining non-stationarity in the system is determined by fewer shocks with a permanent effect. Equivalently, from the modified cointegration test results, it is apparent that, during Phase

[19]Only the Trace test statistic and associated P-value for the null hypothesis of no cointegration against a general alternative are reported. The set of unpresented Trace test statistics fail to reject their corresponding null hypotheses.

[20]In Phase 2, although the individual hypothesis that the EUA futures contract does not respond to the disequilibrium is not rejected, it is evident that the EUA futures contract plays a significant role in the long run relation that has emerged in the system and that the joint hypothesis of zero loading on the disequilibrium and the cointegration equation is clearly rejected.

[21]Further robustness tests have examined the sensitivity of seasonality (day of the week and monthly) on our results. The seasonally adjusted results are quantitatively consistent with those reported here and are available upon request.

2, there are a larger number of relations binding the system together in the long-run, than during Phase 1.

In light of the likelihood of evolving dynamics within the full system of data examined we turn to the recursive cointegration analyses, in relation to the Engle-Granger approach, the Johansen approach and finally the modified cointegration test accounting for heteroskedasticity. Figure 3 presents the results for Engle-Granger, Johansen and the modified cointegration approach. As can be seen the Engle-Granger (1987) recursive test indicates a lack of cointegration throughout and is clearly consistent with the results from table 2. The Johansen (1988) recursive analysis indicates, notwithstanding a brief period in early 2006, a lack of significant distinct cointegration vectors throughout the sample, until a marked strengthening of this result from 2008. The implication is that our finding of contegration is heavily influenced by long-run relationships emerging in Phase 2 of the EU ETS. Finally, turning to the recursive results provided by the robust cointegration methodology (using a Newey-West adjustment), the results again suggest a cointegration relationship developing over Phase 2 only.

5 Conclusion

In January 2005 the European Union (EU) emissions trading scheme (ETS) was instigated, within the Kyoto Protocol, with a view to reducing European emissions of carbon dioxide (CO_2). The scheme issues a restricted amount of emission allowances to firms on an annual basis and allows firms to trade the amount of emission permits that they hold. Hence, the scheme has applied a market value to this externality. In the EU ETS context the first phase of trading was 2005-2007 and the second one, which coincides with the first compliance period of the Kyoto Protocol, is 2008-2012.

A number of studies have examined the performance of the EU ETS market, however given the infancy of the market the emphasis has been on phase 1. Recent studies that examine the time series properties of a range of different EU ETS instruments include Paolella and Taschini (2008), Daskalakis *et al.* (2009) and Benz and Trück (2009). Unlike the previous cited studies which adopted a pure time series approach, Redmond and Convery (2006) and Alberola *et al.* (2008) examine the behaviour of the price of carbon in relation to energy commodities, meteorological factors and a number of other variables. Our current study represents an extension of the later two studies on a number of levels.

In this article we have taken account of market uncertainty and have examined the extent of the emergence of an equilibrium relationship between the expectation of EU allowance prices and a set of theoretically founded factors, including, economic growth, energy prices and weather conditions. Our analysis covers the period 2005 to 2009, so we examine for both Phase 1 (2005-2007) and the current Phase 2 (2008-2012) of the EU ETS. In addition, unlike the vast majority of the previous studies, our focus has been on futures rather than spot contracts. The justification for examining futures (expiration in December 2008 and 2009) is due to the greater volumes being traded on these contracts (see Mansanet-Bataller and Pardo, 2008). These instruments were also not exposed to the dramatic structural breaks that have been previously highlighted in the literature and so results in an additional advantage of adopting the futures based analysis. Specifically, the contribution of the paper is threefold. Our study is the first known study to investigate the long-run relationship between theoretically accepted determinants and EU carbon allowances. Consistent with previous findings, in particular Paolella and Taschini (2008) and Benz and Trück (2009), we find considerable evidence of autoregressive conditional heteroskedasticity (ARCH) effects. As a result, besides taking account of the potential long-run relations, we also address the empirical finding of time varying volatility when testing for cointegration. Finally, given the relatively small sample of data and the considerable uncertainty, in particular during the pilot phase, we examine the extent of the evolving long-run relationships adopting a recursive cointegration approach.

Alongside Phase 2 of the EU ETS it appears that a new pricing regime is emerging in the market. The new regime is indicative of an increasingly active market, following the increased volumes of emissions trading in Phase 2. In particular, it appears that theoretically established relations between the expectations on EU allowance prices and energy spreads and energy prices are now evident. This is not surprising in light of the heightened activity in the EU allowance market during the course of its development. It provides further evidence of the rising level of efficiency in the EU allowance market and is expected to be of interest both to traders, policy makers and those seeking to improve the design of the European Union's Emissions Trading Scheme.

6 Appendix: Data Description

Series	Description
Energy Spreads	Clean dark and spark energy spreads, denominated in Euro per MWh, comprise the discrepancies between the price of electricity at peak hours and the price of coal and the price of natural gas, respectively, required to generate that electricity. These spreads are adjusted for the energy output of the coal / natural-gas fueled plants. They are calculated by Caisse des Depots Climate Task Force for Tendances Carbone, and are observed at a daily frequency from July 1, 2005 through to March 2, 2009. *Source:* http://www.caissedesdepots.fr.
EUAs	European Union Allowance daily futures contract prices, denominated in Euro, observed from July 1, 2005 through to March 2, 2009 with expiration in December 2008 and December 2009. The expiration is switched to December 2009 in the third week of December 2008. The underlying entitlement is the right to emit one tonne of carbon. *Source:* European Climate Exchange.
Equity	The Dow Jones EURO STOXX 50 is denominated in Euro. It's a stock index of futures contracts on 50 Eurozone stocks designed by STOXX Ltd. The data are observed during the period July 1 2005 through to March 2 2009, at a daily frequency. The contracts switch on the first day of each expiry month to the subsequent expiry month futures contract. *Source:* Thomson-Reuters, Datastream
Oil	ICE (Intercontinental Futures Exchange) brent crude oil futures contracts, denominated in Euro are United Kingdom daily contract prices observed from July 1, 2005 through to March 2, 2009 with expiration December 2005, December 2006, December 2007, December 2008 and December 2009. The expiration is altered in the third week of December annually *Source:* Thomson Reuters.
Production	The Eurostat industrial production index has a of base 100 in 2000 and is seasonally adjusted. Observations are recorded between July 1, 2005 and March 2, 2009. Daily observations are estimated via interpolation by adopting a piecewise cubic spline methodology, provided by Matlab. *Source:* http://ec.europa.eu/eurostat.
Temperature	Temperature deviations (absolute) from monthly average temperatures (13-year average) for the Tendances Carbone European temperature index. The data are observed during the period July 1, 2005 through to March 2, 2009. The Tendances Carbone European temperature index is equal to the average of national temperature indices sourced with Powernext. These national temperature indices are computed using weights determined by intra-country regional populations. The European index is weighted by the share of NAP in the constituent countries: France, Germany, Spain and the United Kingdom. *Source:* Tendances Carbone

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Figure 1

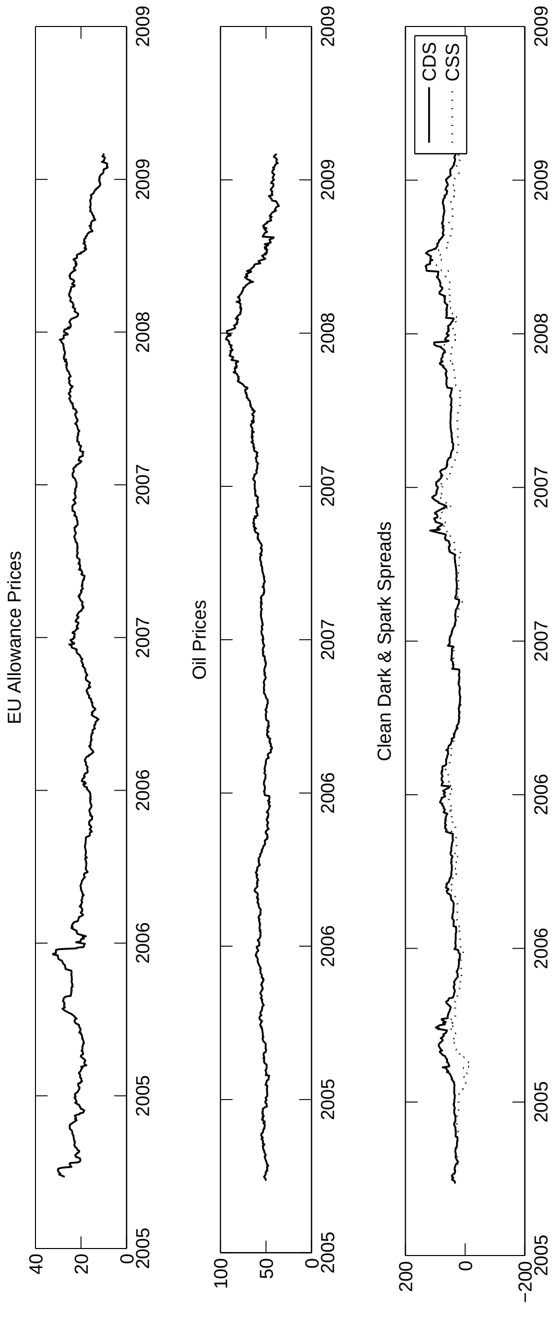
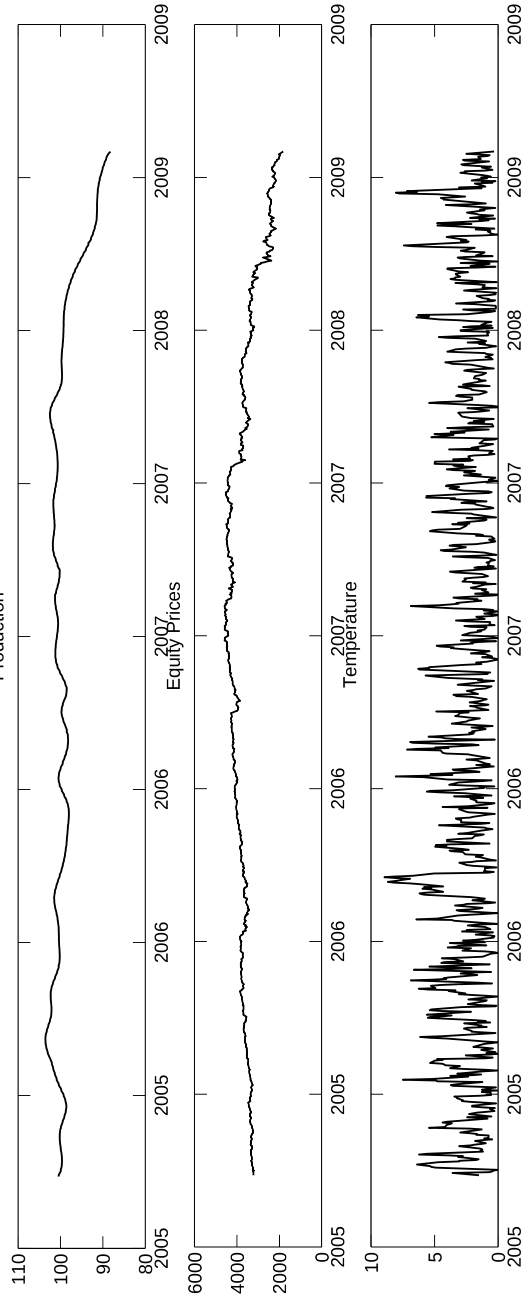

EU Allowance Prices
Oil Prices
Clean Dark & Spark Spreads
CDS
CSS
2005
2006
2007
2008
2009
40
20
0
100
50
200
−200

Figure 2



Production
Equity Prices
Temperature
110
100
90
80
6000
4000
2000
0
10
5
2005
2006
2007
2008
2009

Figure 3: Recursive Cointegration Tests

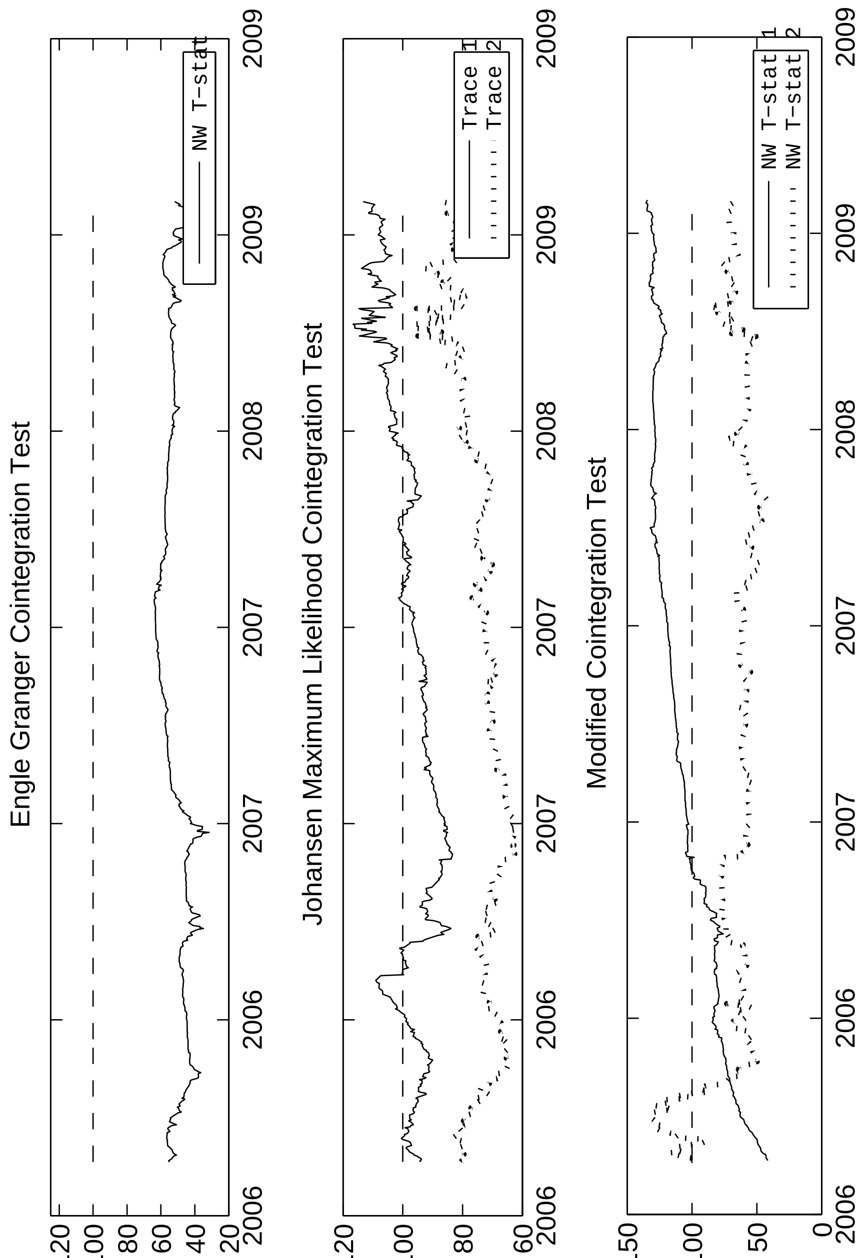


Table 1: **Summary Statistics**

Panel A: *Full Sample*

Series	Mean	Variance	Skewness	Exc. Kurtosis	ARCH	PP Unit Root	LS Unit Root
EUA	-0.11	8.89	-1.08*	12.83*	35.5*	-2.26[-26.99*]	-4.05[-12.43*]
CDS	-0.01	60.71	2.25*	22.34*	0.35	-2.30*[-28.90*]	-3.99[-18.09*]
CSS	-0.02	74.83	-3.47	63.47*	16.28*	-2.99*[-28.91*]	-3.34[-14.92*]
Equity	-0.06	2.62	-0.06	10.06	212.56*	0.73[-31.95*]	-4.58[-16.26*]
Oil	-0.03	5.15	1.58*	20.45*	11.56*	-1.17[-34.25*]	-2.39[-13.80*]
Prod.	-0.01	0.00	-0.31	3.76*	164.45*	4.43[-4.18*]	-6.08*[-5.95*]
Temp.	2.34	2.93	0.98*	0.72*	475.02*	-13.82*[-48.28*]	-10.45*[-21.05*]

Panel B: *Phase I*

Series	Mean	Variance	Skewness	Exc. Kurtosis	ARCH	PP Unit Root	LS Unit Root
EUA	-0.03	9.38	-1.48*	16.36*	20.89*	-3.07*[-22.74*]	-4.14[-10.39*]
CDS	0.17	66.28	2.59*	21.57*	1.40	-2.02[-24.02*]	-2.90[-12.33*]
CSS	0.09	90.23	-4.01*	62.73*	0.91	-2.33[-28.25*]	-3.28[-15.47*]
Equity	0.05	0.82	-0.39*	.99*	49.11*	-1.46[-26.63*]	-3.71[-12.16*]
Oil	0.04	2.53	0.39	2.88	3.46	-1.93[-27.76*]	-3.26[-11.93*]
Prod.	0.00	0.00	0.29	0.20	42.72*	-1.29[-5.33*]	-6.40*[-3.69*]
Temp.	2.46	3.16	0.89*	0.43*	329.32*	-11.13*[-40.17*]	-8.50*[-17.05*]

Panel C: *Phase II*

Series	Mean	Variance	Skewness	Exc. Kurtosis	ARCH	PP Unit Root	LS Unit Root
EUA	-0.28	7.85	0.02*	2.33*	61.71*	.29[-14.52*]	-3.35[-7.90*]
CDS	-0.34	66.28	0.97*	23.41*	1.40	-1.87[-15.98*]	-3.11[-9.26*]
CSS	-0.23	48.60	1.05*	14.29*	0.91	-2.52[-18.63*]	-2.62[-8.95*]
Equity	-0.30	6.43	0.23	3.83*	49.11*	-0.62[-18.22*]	-3.28[-9.85*]
Oil	-0.17	10.77	1.66*	13.70*	1.67	-0.13[-19.53*]	-2.55[-10.24*]
Prod.	-0.05	0.00	-1.36*	9.55*	42.72*	3.43*[-3.92*]	-5.60*[-2.26]
Temp.	2.09	2.37	1.13*	1.53*	138.13*	-8.44*[-26.44*]	-6.20*[-12.4*]

Panels A, B and C correspond to the sample periods examined in this study. Panel A spans the full sample period. Panel B spans part of *Phase I* of the European Union Emissions Trading System (July 1 2005 to December 31 2007) and panel C spans part of *Phase II* (January 2 2008 to March 2 2009) of that system. A constant of 30 is added to the Clean Spark Spread (CDS) observations to facilitate logarithmic calculations. The Clean Spark Spread is denoted CSS. Prod. and Temp. correspond to production and temperature respectively. In column 5 the Lagrange Multiplier (LM) test results are reported for fifth order ARCH effects. In columns 6 and 7, the Phillips-Perron (PP) unit root test statistics and Lee-Strazicich (LS) unit root test statistics are reported. The test statistics for each of the series in logarithmic differences are reported in square brackets, while the test statistics with respect to levels are adjacent. A * indicates statistical significance at the 5% level.

Table 2: **Preliminary Investigation of Cointegration Equations**

Panel A: *Full Sample*

	CDS	CSS	Equity	Oil	Production	Temperature	DF T-Stat
European Union Allowances	-0.001	-0.04	0.16	0.81*	-0.29	0.01	2.44
	(0.21)	(1.07)	(1.24)	(11.89)	(1.12)	(1.12)	

Panel B: *Phase I*

	CDS	CSS	Equity	Oil	Production	Temperature	DF T-Stat
European Union Allowances	-0.04	0.07	-0.66*	1.04*	0.92*	-0.00	2.97
	(1.41)	(1.77)	(4.68)	(5.44)	(3.42)	(0.96)	

Panel C: *Phase II*

	CDS	CSS	Equity	Oil	Production	Temperature	DF T-Stat
European Union Allowances	0.38*	-0.22	0.63*	0.74*	-1.27*	0.01*	3.00
	(4.14)	(1.85)	(6.11)	(11.40)	(6.24)	(2.36)	

Panels A, B and C correspond to the sample periods examined in this study. Panel A spans the full sample period. Panel B spans part of *Phase I* of the European Union Emissions Trading System (July 1 2005 to December 31 2008) and panel C spans part of *Phase II* (January 2 2008 to March 2 2009) of that system. The results are Newey-West Linear regression coefficients and coefficient t-statistics in brackets. The Newey- West t-Statistics (far right hand side column) are reported regarding the null hypothesis of a unit root in the residual from the hypothesised Cointegration Equations. The critical values are sourced in MacKinnon (1991). A * indicate statistical significance at the 5% level.

Table 3: **Maximum Likelihood Cointegration Test and Hypothesis Testing**

Panel A: *Normalized Cointegration Vectors*

Full Sample						
	Oil	CSS	CDS	Production	Equity	Constant
Coeff.	0.36*	0.59*	-0.23*	10.22*	-1.40*	-35.52*
	(2.77)	(5.36)	(3.29)	(7.74)	(6.36)	(7.64)
Phase I						
	Oil	CSS	CDS	Production	Equity	Constant
Coeff.	-0.28	0.09	-0.06	-9.99	-0.34	50.65
	(0.23)	(0.21)	(.24)	(1.22)	(0.39)	(1.44)
Phase II						
	Oil	CSS	CDS	Production	Equity	Constant
Coeff.	0.64*	-1.35*	1.30*	2.12	-0.02	-9.42
	(3.76)	(3.38)	(4.06)	(1.03)	(0.05)	(1.42)

Panel B: *Hypotheses Tests*

Full Sample Period		*Phase I*		*Phase II*	
(i)		(i)		(i)	
Trace Test	113.11	Trace Test	104.22	Trace Test	468.72
P value	0.01	P value	0.05	P value	0.00
(ii)		(ii)		(ii)	
Test Statistic	6.94	Test Statistic	0.01	Test Statistic	0.08
P value	0.01	P value	0.91	P-Value	0.78
(iii)		(iii)		(iii)	
Test Statistic	16.30	Test Statistic	2.17	Test Statistic	7.45
P value	0.00	P value	0.14	P value	0.01
(iv)		(iv)		(iv)	
Test Statistic	17.29	Test Statistic	2.17	Test Statistic	7.47
P value	0.00	P value	0.34	P value	0.02

Panel A presents a distinct normalised cointegration equation, with associated t-statistics in brackets, for each sample period examined in this study. The full sample period extends from July 1 2005 through to March 2 2009. Part of *Phase I* of the European Union Emissions Trading System (July 1 2005 to December 31 2007) is examined and part of *Phase II* (January 2 2008 to March 2 2009) of that system is also examined. The model specifications (deterministic components and lag length) are inferred with respect to the multivariate version of the Schwarz Bayesian information criterion. These results are not presented here. Panel B presents related hypotheses tests. The likelihood ratio Trace test statistic (i) indicates that there is at least a single cointegrating equation (CE) in each of the sample periods examined. The remaining hypotheses tests (ii), (iii) and (iv) assess the null hypotheses of a zero loading coefficient on the disequilibrium error in the EUA equation, a zero EUA coefficient in the distinct cointegration equation and a joint hypothesis test to assess these latter two null hypotheses, respectively. A * indicates statistical significance at the 5% level.

Table 4: **Modified Multivariate Test for Cointegration**

Panel A: *Full Sample*						
				Critical Values		
	OLS Coeff.	GARCH Coeff.	t-statistic	10%	5%	1%
$\rho = 1$	0.22	0.13	6.64*	3.81	4.10	4.65
$\rho = 2$	0.18	0.14	3.57	3.45	3.75	4.29
Panel B: *Phase I*						
				Critical Values		
	OLS Coeff.	GARCH Coeff.	t-statistic	10%	5%	1%
$\rho = 1$	0.18	0.16	6.26*	3.81	4.10	4.65
$\rho = 2$	0.18	0.11	2.99	3.45	3.75	4.29
Panel C: *Phase II*						
				Critical Values		
	OLS Coeff.	GARCH Coeff.	t-statistic	10%	5%	1%
$\rho = 1$	0.26	0.15	4.58*	3.81	4.10	4.65
$\rho = 2$	0.26	0.22	5.50*	3.45	3.75	4.29

Panels A, B and C correspond to the sample periods examined in this study. The *Full Sample* period extends from July 1 2005 to March 2 2009. This study examines part of *Phase I* of the European Union Emissions Trading System (July 1 2005 to December 31 2007) and part of *Phase II* (January 2, 2008 to March 2, 2009) of that system. The Panels report coefficients for $\rho = 1, 2$ which are the estimated square roots of the eigenvalues, while accounting for t-distributed GARCH effects, of the Johansen long-run information matrix. These coefficients are also estimated using OLS in conjunction with the Newey-West standard error estimator. The coefficients are estimated by equation (14). See McKinnon (1991) for critical values. A * indicates statistical significance at the 5% level.