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TESTING MARKET EFFICIENCY AND PRICE DISCOVERY IN EUROPEAN CARBON MARKETS

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ABSTRACT

We examine the issues of market efficiency and price discovery in the European Union carbon futures market. Our findings suggest that none of the carbon futures contracts examined here are priced according to the cost-of-carry model, although two of the three futures contracts studied here form a stable long-run relationship with the spot price, and hence act as adequate risk mitigation instruments. We apply a new testing procedure and find weak evidence of convenience yield in the market for carbon allowances. In terms of price discovery, it appears that the spot and futures markets share information efficiently and contribute to price discovery jointly. Similar to the information diffusion pattern found in returns, we report some evidence of bi-directional volatility transfers between the spot and various futures contracts.

JEL Classifications: C32, G13, G14, C32, Q25, Q40

Keywords: Carbon-dioxide allowances, futures, cost-of-carry, price discovery, market efficiency, cointegration, granger causality, volatility spillover, global warming.

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1. Introduction

Climate change is a major challenge faced by the international community. The effects of human-caused greenhouse gases and global warming are becoming increasingly visible from record temperatures to rising sea levels. In economic terms the Stern review (Stern 2006) reports that, in the absence of policy action, climate change is likely to reduce world GDP by between 5% and 20% permanently. In contrast, the cost of acting to reduce greenhouse gas emissions is estimated to be around 1% of world GDP. In an attempt to slow down and stabilise the pace of climate change, most countries (excluding the US and Australia¹) have signed and ratified the Kyoto Protocol (UNFCCC 1998) to the UN Framework Convention on Climate Change (UNFCCC). Based on a “cap and trade” system, the protocol sets targets for the reduction of greenhouse gases (GHG) and facilitates the trading of permits to emit GHGs between countries and individual entities. The existence of a trading mechanism allows most GHG abatement to occur in those sectors of the economy in which it is cheapest—achieving the cap with the lowest possible economic impact. So far several international markets for carbon permits have emerged, with the European Union Emissions Trading Scheme (EU ETS) at the forefront in terms of both the market size and its regulatory organisation.

While the existing literature on carbon markets provides theoretical arguments for and against such schemes (e.g. Cooper 1996; Klepper *et al.* 2003; McKibbin *et al.* 1999; McKibbin and Wilcoxon 2006) the effectiveness of any existing carbon trading scheme will rely on the ability of the market mechanism to produce prices which accurately reflect the true marginal costs of GHG abatement. In this context, the important issue of market efficiency and price discovery in carbon *derivative* markets also arises. Derivative products, such as futures, options and swaps facilitate risk mitigation through reassignment of risk between economic agents of different risk preferences. However, a necessary condition for effective risk management is the existence of a long-run link

¹ These countries have implemented their own state mandatory/voluntary carbon trading systems.

between the spot price and the derivative price (futures price in our case) that is achieved through an equilibrium pricing relationship. If such a link did not exist, the spot and futures prices could diverge by assuming independent stochastic paths and the futures position that was meant to mitigate price risk would instead result in additional risk exposure. Thus, an independent (inefficient) futures market has a potential to undermine the efficacy of any carbon trading scheme.

In this paper we aim to answer three questions related to carbon market efficiency. First, do the EU ETS carbon spot and futures prices form a stable long-run relation? If the answer to this question is yes, we then proceed to test a stronger assumption by asking the following question: Is the long-run link between the spot and futures prices given by a no-arbitrage cost-of-carry pricing model? Within this framework we also test for the existence of a convenience yield obtained by holding a spot position. Lastly, we address the issues of price discovery by analysing information spillovers between the cash and futures prices. Thus, we seek to uncover which market reflects new information first and leads the price discovery process. The three issues addressed here have been studied in the context of other commodities and financial assets but not carbon permits. Therefore we present, to our knowledge, a first attempt at analysing market efficiency and price discovery in the setting of carbon allowance markets.

Our empirical methodology includes a cointegration analysis of spot and futures carbon prices and interest rates, Granger causality tests and multivariate GARCH volatility methods. We also present a new method for measuring and testing for convenience yield within the framework of cointegration. Our dataset consists of the European Union carbon emission allowance (EUA) spot and futures prices and interest rates. In our approach we track three futures contracts with expiry dates in December 2006, December 2007 and December 2008 through time, match the contracts' time to expiry with interest rates of equal maturity and test the cost-of-carry model. The period analysed is determined by the availability of data and covers the time frame June 2005–November 2006.

We find stable long-run links between the carbon cash price and the price of the December 2006 and December 2007 futures contracts. This finding has a

significant implication for risk management and suggests that each of these two futures contracts can be used effectively for hedging purposes. However, our tests suggest that neither of these two contracts' prices is consistent with the cost-of-carry model. Although, a likelihood ratio test does not reject the null of cost-of-carry model for the December 2006 futures price, we argue that this is due to a large standard error associated with the estimated parameter on the interest rate variable according to which the null of the coefficient being equal to zero cannot be rejected either. While the December 2006 contract does not seem to be affected by changes in the interest rates, the relation between the December 2007 contract and interest rates is different from that predicted by the pricing model. The violation of the cost-of-carry relation can be interpreted as an indicator of market inefficiency and points to arbitrage opportunities that may have existed over the sample period.

The futures contract that expires in December 2008 does not exhibit a stable long-run relationship with the spot price. Even though this finding may, at a first glance, cast doubt on the value of the December 2008 contract, we argue that there is a rational explanation behind it and that the lack of a long-run relationship in this case is due to the unavailability of a relevant carbon spot price for the Phase II of the EU Emissions Trading Scheme. Briefly, the EU ETS is implemented in two stages: Phase I (2005-2007) and Phase II (2008-2012). Unused carbon allowances from Phase I are not expected to be valid in Phase II and will expire at the end of 2007.² Therefore, the December 2008 contract is not expected to be cointegrated with the current spot price but in fact acts as a vehicle of price discovery for the future Phase II spot price.

We find evidence of a convenience yield in the December 2007 futures contract pricing equation, no convenience yield in the December 2006 equation, and we are unable to conduct the test on the December 2008 contract because of the lack of a cointegrating relation between this contract and the cash contract. Although our evidence suggests that there may exist a convenience yield in

² See Section 4 for more detail.

certain carbon market segments, this finding should be interpreted with care because our tests are based on the assumption of the cost-of-carry model.

Mosconi and Giannini (1992) and Toda and Yamamoto (1995) Granger causality tests indicate that in two out of the three cases examined here, there is bi-directional Granger causality between the spot and futures prices. This can be interpreted as evidence of a price discovery process that occurs in both the spot and futures market, a finding in contrast with evidence from other commodity markets in which the futures price serves as a vehicle of price discovery. We also find that the spot and future carbon returns exhibit autoregressive conditional heteroscedasticity (ARCH). Testing for volatility spillovers using a BEKK multivariate GARCH specification (Engle and Kroner 1995), we find no volatility spillovers between the December 2006 futures contract and the spot price, evidence of bi-directional spillovers between the December 2007 futures and the spot contracts, and no evidence of volatility spillovers between the December 2008 contract and the spot price.

Overall it seems that the EU ETS futures market fulfils its roles of providing a means for effective risk management as well as a contribution to the price discovery process in the spot market. However, it also appears that this relatively new market exhibits a number of idiosyncrasies (e.g. the lack of market efficiency) relative to developed markets that participants should take into account when transacting in carbon futures.

The rest of this paper is organised as follows. We present a brief literature survey in Section 2, describe the data and outline econometric method in Section 3, and provide our empirical findings in Section 4. We conclude in Section 5.

2. Theoretical and empirical links between spot and futures prices

In theory, and for assets that allow arbitrage between spot and futures markets, futures contracts can be priced according to a no-arbitrage cost-of-carry model. If $F_{t,T}$ is the current price of a futures contract that expires in $(T-t)$ years and S_t is

the current spot price, the cost-of-carry relation links the spot and futures prices in the following way:

$$F_{t,T} = S_t e^{(r_t - \delta)(T-t)} \quad (1)$$

where r_t is a risk-free interest rate and δ can be thought of as either a dividend yield in the case of a dividend paying stock, or a convenience yield³ in the case of a commodity. The cost-of-carry model thus posits that in an efficient market the futures price should equal the spot price, adjusted for the opportunity cost of holding a spot position, i.e. the interest foregone, less a dividend/convenience yield. Because the cost-of-carry model is derived from a no-arbitrage condition that operates on a risk-free portfolio (see, for example, Hull 1997, Ch. 3) the relevant discount rate (r_t) is the risk-free rate.

The convenience yield (δ) is typically interpreted as the value of privilege of holding a unit of inventory, e.g. to be able to meet unexpected demand or to keep a production process running. It should be measured and tested within a framework of a pricing model, which introduces difficulty in insulating any test of the convenience yield from a test of the pricing model itself. Although the literature has addressed some of these issues (e.g. Dusak 1973; Carter, Rausser and Schmitz 1983; Fortenbery and Hauser 1990) there is no clear consensus as to whether convenience yields exist or not (Zapata and Fortenbery 1996).

In this paper we provide a new method for testing for the existence of convenience yield that is based on a cost-of-carry relationship. Further, we put forward and test a new hypothesis relating to EU ETS carbon contracts which proposes that a convenience yield may exist in the pricing formula for longer maturity futures contracts but should be equal to zero in the last year of any EU ETS carbon contract life. The rationale behind our hypothesis is a market regulation under which EU ETS affected installations settle their carbon obligations. Specifically, the affected firms are required to surrender all their carbon permits for emissions produced during any calendar year at the same time:

³ Convenience yield refers to the value of the flow of services associated with holding an inventory of a commodity.

April of the following calendar year. Thus, a spot permit differs from a futures permit with time to expiry of less than one year only by the time value of money because these contracts will be converted into spot prior to the settlement month of April and hence fulfil the same role as the spot contracts in terms of offsetting carbon obligations. On the other hand, the industry may obtain convenience from holding a cash position relative to a futures contract that expires at a date following the settlement month of April because such futures contracts cannot be used to offset current period carbon obligations.

In a perfectly efficient and frictionless market the pricing relationship expressed in Eq. (1) should hold at every instant over a futures contract life (Stoll and Whaley 1990). However transaction costs and other market imperfections create a band of no-arbitrage price interval within which we can expect the futures prices to fluctuate. MacKinlay and Ramaswamy (1988) list a number of factors that can drive the futures price away from its theoretical value. For example, stochastic interest rates result in arbitrage positions becoming risky because of unanticipated interest earnings/costs associated with marking to market. Further, the longer the time to maturity the larger the risk associated with unanticipated changes in the dividend/convenience yield. Chung (1991) also notes that there may exist other forms of risk premiums associated with the cost-of-carry trades. According to Chung, attempting to exploit observed deviations from the no-arbitrage bounds can, *ex ante*, prove to be risky due to order execution lags. That is, traders are not guaranteed execution of their orders at the observed prices. Chung shows that once execution lags and transaction costs are taken into account the size and frequency of boundary violations are substantially smaller than those reported by other studies. In summary, the pricing error from the cost-of-carry relation can be expected to fluctuate between some no-arbitrage boundaries as follows:

$$b_{L,t} < \left(F_{t,T} - S_t e^{(r_t - \delta)(T-t)} \right) < b_{U,t} \quad (2)$$

where are $b_{L,t}$ and $b_{U,t}$ lower and upper no-arbitrage interval bounds created by the above mentioned factors.

Empirical tests of the cost-of-carry model come in two flavours: tests of the lead-lag relationship between spot and futures prices and tests of market efficiency as given by the pricing relation in Eq. (1). The two strands of literature are closely related and complement each other on the basis of the Granger Representation Theorem (Engle and Granger 1987). In particular, when a futures contract is priced efficiently the futures and the underlying spot prices will be linked through Eq. (1) that is they will be cointegrated.⁴ However, if they are cointegrated then according to the Granger Representation Theorem there exists some form of lead/lag relationship between the cointegrating variables. We summarise the literature in the following statement. While the empirical evidence is mixed, it appears that futures prices lead/Granger cause spot prices and that the cost-of-carry model holds in the market for financial assets but not in the market for commodities. A more detailed literature survey is given below.

2.1 *Empirical tests of the cost-of-carry model*

MacKinlay and Ramaswamy (1988) investigate deviations from the cost-of-carry pricing of the S&P 500 index and index futures. They use intraday 15 minute data and find that the departures from the cost-of-carry relationship (i.e. mispricing) increase with time and that they are path-dependent, even after controlling for non-synchronous trading. Harris (1989) also examines the relation between the S&P 500 index and futures contract over a ten-day period surrounding the October 1987 stock market crash. He reports that non-synchronous trading of constituent stocks explains some but not the entire large futures-cash basis (price difference) observed during the crash. Harris attributes the remainder of the basis to disintegration between the two markets. Similarly, Chung (1991) investigates the efficiency of the market for stock index futures and the profitability of index arbitrage for Major Market Index (MMI) contracts. After accounting for transaction costs, execution lags and the uptick rule for short sales, Chung finds

⁴ To see this we take the log of both sides of equation 1: $\ln(F_{t,T}) = \ln(S_t) + (r_t - \delta)(T - t)$. This relationship is not exact and we need to add an error term. The error term is assumed to be stationary. This implies that $\ln(F_{t,T})$, $\ln(S_t)$ and r_t are cointegrated. A deterministic trend should also be included in the relationship.

that the size and frequency of no-arbitrage boundary violations are substantially smaller than those reported in the literature, and have declined over time.

In commodity markets, Bessler and Covey (1991) examine the US cattle markets and report some evidence of cointegration between cash prices and futures contracts closest to delivery. They find, however, no evidence of cointegration for more distant futures contracts analysed over the August 1985–August 1986 period. Schroeder and Goodwin (1991) fail to find cointegration between spot and futures prices for the US live hogs while Fortenbery and Zapata (1993) find support for the cointegration hypothesis between corn and soybeans spot and futures contracts. None of the abovementioned cointegration studies however take into account the possibility that interest rates may be stochastic. Zapata and Fortenbery (1996) correct for this omission and examine a trivariate cointegrating relationship between spot and futures prices in selected commodity markets and interest rates. They find trivariate cointegration for most storable commodities and suggest that previous findings of no cointegration are based on inappropriate specifications due to omissions of interest rates from the cointegration relationships. Further evidence of cointegration between cash and futures markets is given by Fontenbery and Zapata (1997) for anhydrous ammonia (NH₃) and diammonium phosphate (DAP) contracts traded on the Chicago Board of Trade. On the other hand, Fontenbery and Zapata (1997) find no cointegration between the futures and cash markets for cheddar cheese, with or without the specification of interest rates in the hypothesised model, over the June 1993–July 1995 time period. Thraen (1999) addresses this finding using a longer time period of July 1993–October 1997, and shows that the cheddar spot and futures prices, together with interest rates, have established a stable long-run relationship and are cointegrated with at most one cointegrating relation. Thraen thus concludes that the equilibrium relationship between the cheddar cash and futures markets had not been developed over the time period studied by Fontenbery and Zapata, but emerged shortly thereafter.

2.2 *Empirical tests of the lead-lag relationship between spot and futures prices*

Empirical analyses of price discovery typically rely on Granger type causality tests between spot and futures prices or volatility spillover tests. Thus if the futures price (volatility) is found to lead the spot price (volatility) then the futures market is said to dominate the spot market and vice versa. Other approaches exploit the fact that, if the spot and futures prices are cointegrated, the price vector can be decomposed into a permanent and a transitory component (Stock and Watson 1988) where the permanent component is assumed to be an unobserved efficient market price. Information shares to price discovery are then measured as contributions of each of the markets to the efficient price using, for example, Hasbrouck (1995) and Gonzalo and Granger (1995) decompositions. While the empirical evidence is mixed there seem to be more accounts of the futures market dominating the spot market than the other way around.

Kawaller, Koch and Koch (1987) test the lead-lag relationship between the S&P 500 futures and its underlying index using minute-to-minute data over the 1984–1985 time period. They report that the futures price leads the cash index movement by 20 to 45 minutes while the cash index does not affect the futures beyond one minute. This finding is typically interpreted as evidence in support of the hypothesis that the futures market acts as a vehicle of price discovery in the spot market. Stoll and Whaley (1990) note that previously reported lead-lag effects between spot and futures prices could be due to infrequent or non-synchronous trading of stocks included in the index. Using 5-minute returns on the S&P 500 stock market index, the S&P 500 index futures and the Major Market Index (MMI) futures, they report that the futures returns lead the spot returns by five to ten minutes even after adjusting the stock index returns for non-synchronous trading effects. They also report that the lead has diminished over the 1982–1987 period. Ng (1987) and Chan (1992) also report evidence that the futures markets lead spot prices. Chan tests the lead-lag relation between returns of the Major Market cash index and returns of the Major Market Index (MMI) futures and the S&P 500 futures. He also finds strong evidence that the futures market leads the cash index, but weak evidence that the cash index leads the

futures. Furthermore, he reports that when more stocks move together (market-wide information) the futures leads the cash index to a greater extent. Chan interprets this as evidence that the futures market is the main source of market-wide information.⁵

Tse (1999) studies minute-by-minute price discovery and volatility spillovers between the Dow Jones Industrial Average (DJIA) index and the index futures using Hasbrouck's (1995) cointegrating model and a bivariate EGARCH specification. He finds that the price discovery takes place in the futures market but that there are significant bivariate transmissions of volatility. Tse shows that about 88.3% of the information share to the efficient market price is attributable to the futures market. Chan, Chan and Karolyi (1991) and Koutmos and Tucker (1996) examine volatility spillovers between the S&P 500 index and futures prices. While Chan, Chan and Karolyi find bidirectional spillovers, Koutmos and Tucker report that volatility spills only from the futures markets to the spot market. Furthermore, Koutmos and Tucker document an asymmetric volatility spillover effect: bad news increases volatility more than good news.

In the commodities market, Figuerola-Ferretti and Gilbert (2005) examine the price discovery process for aluminium using an extended version of the Beveridge and Nelson (1981) decomposition of four different aluminium price series. They find that the start of aluminium futures trading in 1978 resulted in greater price transparency in the sense that the information content of transactions prices increased. Fontenbery and Zapata (1997) find that the futures market for fertilizer products lead the underlying cash markets.

A number of studies, on the other hand, suggest that the spot market leads the futures market. Amongst them are Green and Joujon (2000) who find evidence that the French CAC-40 spot stock market index leads its futures contract, and Silvapulle and Moosa (1999) who find bi-directional non-linear Granger causality between futures and spot prices in the crude oil market. Quan (1992) examines the

⁵ Additional evidence that the futures market dominates the spot market can be found in Abhyankar (1995) for the UK stock market; Fleming *et al.* (1996) for the US market; Twite (1991) for Australia; Tang *et al.* (1992) for Hong Kong; and Puttonen (1993) for Finland.

price discovery function in the crude oil market and finds that the futures market does not contribute significantly to the price discovery process.

3. Econometric methodology

The no-arbitrage condition, when applied to an imperfect market with frictions such as transaction costs, stochastic interest rates, etc., suggests that Eq. (1) holds in the long-run but not necessarily in the short-term. After taking natural logarithms Eq. (1) can be expressed as a cointegrating relation of the following form:

$$f_t = s_t + r_t(T-t) - \delta(T-t) + u_t \quad (3)$$

where $s_t \equiv \ln S_t$, $f_t \equiv \ln F_t$, T is time when the futures contract matures (in years), t is the current time (in years) and u_t is a stationary zero-mean innovation term whose variance is determined by the extent of market imperfections. Note also that the terms in brackets of the above equation represent reverse time trends that start at T (years to maturity) and end at zero as t approaches T . In order to test the cost-of-carry model we can re-write Eq. (3) as:

$$\begin{aligned} f_t &= as_t + b(r_t(T-t)) - d(T-t) + \eta_t \\ &= -dT + as_t + b(r_t(T-t)) + dt + \eta_t \end{aligned} \quad (4)$$

and test the following hypotheses:

1. Eq. (4) forms a cointegrating relationship, i.e. η_t is stationary.
2. $H_0 : a = b = 1$ i.e. restrictions implied by the cost-of-carry model.

The above hypotheses 1 and 2 have the following interpretations. Should we find that both 1 and 2 hold, then there exists a long-run relationship between the spot and futures carbon prices and the interest rates that is consistent with the cost-of-carry model. This finding would imply that the futures market is efficient and that there are no arbitrage opportunities between the spot and futures prices. Futures contracts would also be a suitable risk-mitigation instrument. An alternative finding would be to discover that hypothesis 1 holds while hypothesis 2 does not. In this situation Eq. (3) would represent a cointegrating relationship

but the restrictions $a = 1$ and $b = 1$ would be rejected and we would conclude that while the three markets are linked in the long-run, the relationship is not provided by arbitrage activity and hence arbitrage opportunities persist in the market. However, even though arbitrage opportunities may exist in this case, the finding of a long-run relation (with or without the cost-of-carry restrictions) implies that futures contracts can still be regarded as an adequate instrument for risk management. Lastly, failure to find any kind of long-run relation would imply two things. First, it would mean that the futures contract price behaves independent of the underlying spot price and thus should not be used for hedging. Second, it would mean that there are arbitrage opportunities in this market.

Besides testing the above hypothesis we are also in a position to estimate the convenience yield d . Thus we also test the zero-convenience yield hypothesis for any contract in its last year to expiry using the following hypothesis:

$$3. \quad H_0 : d = 0 \text{ for } (T-t) < 1$$

where $(T-t)$ is a number of days to expiry divided by 360. At the time of writing this paper, only the December 2006 futures contract was eligible for this test.

3.1 Model specification

In order to formulate tests of the above hypotheses we follow Engle and Granger⁶ (1987) and respecify the model in differenced log prices as a Vector Error Correction Model (VECM):

$$\begin{aligned} \Delta s_t &= \alpha_0 + \alpha_1 (\eta_{t-1}) + \sum_{i=1}^{M-1} \alpha_{2i} \Delta s_{t-i} + \sum_{j=1}^{M-1} \alpha_{3j} \Delta f_{t-j} + \sum_{k=1}^{M-1} \alpha_{4k} \Delta r_{t-k} + e_{1t} \\ \Delta f_t &= \beta_0 + \beta_1 (\eta_{t-1}) + \sum_{i=1}^{M-1} \beta_{2i} \Delta s_{t-i} + \sum_{j=1}^{M-1} \beta_{3j} \Delta f_{t-j} + \sum_{k=1}^{M-1} \beta_{4k} \Delta r_{t-k} + e_{2t} \end{aligned} \quad (5)$$

where $\eta_t = dT + f_t - as_t - b(r_t(T-t)) - dt$ is derived from Eq. (4). We assume that the risk-free interest rate is stochastic and exogenously determined so we do not

⁶ Engle and Granger (1987) show that if two prices are cointegrated, i.e. exhibit a stable long-run relationship, then there exists an error correction representation of this relation.

model its process. Within this framework we examine temporal links between the spot and futures returns by testing the following hypotheses:

1. Spot market Granger does not cause Futures market, i.e.

$$H_0 : \beta_1 = \beta_{2i} = 0 \text{ for all } i .$$

2. Futures market Granger does not cause Spot market, i.e.

$$H_0 : \alpha_1 = \alpha_{3j} = 0 \text{ for all } j$$

Given that we conduct Granger causality tests in a cointegration framework we implement them as Mosconi and Giannini (1992) and Toda and Yamamoto (1995) tests. Mosconi and Giannini's (1992) causality test is derived by explicitly applying the cointegration restrictions both under the null and the alternative hypotheses. It is an extension of the Johansen's (1991) likelihood ratio tests for linear restrictions on cointegrated systems. The distributions of these tests are χ^2 . Toda and Yamamoto (1995) show that having chosen a VECM of lag length M we can test for causality in a VAR model specified in levels where the lag length is $M+d$, d being the maximal order of integration of any process in the system.

Additional to the above causality tests that we conduct on mean equations we also perform tests for information spillovers in conditional variance equations, i.e. volatility spillovers⁷. Here we specify the conditional variance matrix \mathbf{H}_t of the spot and futures return vector process $(\Delta s_t \ \Delta f_t)'$ as a BEKK(1,1) multivariate GARCH model (Engle and Kroner, 1995):

$$\mathbf{H}_t = \mathbf{\Gamma}'\mathbf{\Gamma} + \mathbf{\Lambda}' \mathbf{e}_{t-1} \mathbf{e}_{t-1}' \mathbf{\Lambda} + \mathbf{\Phi}' \mathbf{H}_{t-1} \mathbf{\Phi} \quad (6)$$

where $\mathbf{\Gamma}$ is an upper triangular (2×2) matrix while $\mathbf{\Lambda}$ and $\mathbf{\Phi}$ are (2×2) matrices of parameters. Thus, our last set of hypothesis tests takes the following form:

1. The Spot market does not transfer volatility to the Futures market

$$H_0 : \lambda_{22} = 0, \phi_{22} = 0$$

⁷ Volatility spillovers are typically defined as transmissions of risk/turbulence between financial assets/markets.

2. The Futures market does not transfer volatility to the Spot market

$$H_0 : \lambda_{21} = 0, \phi_{21} = 0$$

4. Data description

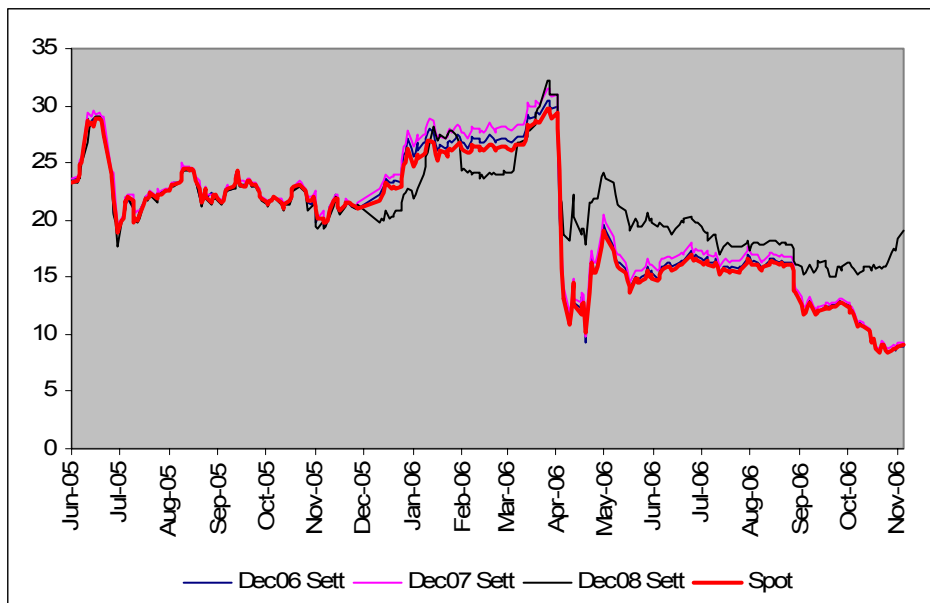
We use spot and futures prices of carbon allowances that trade under the European Union Emissions Trading Scheme (EU ETS). The European Union allowances (EUAs) are allocated to carbon emitting installations in quantities predetermined by the EU member state governments and trade on both the over-the-counter markets and organised exchanges. The trading occurs between firms that over-emit and under-emit relative to their allocations and other market participants such as speculators and arbitrageurs. The total number of permits issued is consistent with the path undertaken by the EU towards its overall Kyoto commitment (reduction of greenhouse gas emissions 8% below its 1990 level by the year 2012). The EUAs are legally binding and the affected installations are required to surrender permits for emissions produced each calendar year in the month of April of the following year. In the case where an insufficient number of permits are surrendered, the affected institution is charged a penalty and is still liable to surrender the deficit permits. Thus the penalty charged does not represent an upper limit of the price of carbon emissions.

The EU ETS itself is designed to operate in two phases. Phase I (2005–2007) is a pilot program that is expected to prepare the EU for compliance with the Kyoto Protocol. In its first phase the EU ETS covers about 40% of the total EU CO₂ emissions produced mainly by the electricity sector. During the course of the second phase (2008–2012) the EU ETS will broaden the coverage to other industries and will include other types of greenhouse gases (e.g. methane, nitrous oxide, etc). The EU ETS is expected to achieve full compliance with the Kyoto Protocol in Phase II. Although some EU member countries have suggested that unused Phase I permits will be valid GHG offsetting instruments in Phase II, the general opinion among the EU states is that the two are inconsistent and that Phase I permits will expire at the end of 2007. Although the EU ETS is a relatively new market, launched on 1 January 2005, it has grown at an

unprecedented rate recording a total trading volume in excess US\$6.5 billion in the first quarter of 2006 alone.

Our dataset consists of daily observations on interest rates, carbon allowance spot prices and December 2006, December 2007 and December 2008 carbon futures contracts. The EUA spot contract is traded on a number of European exchanges including the Nordpool® and Powernext®. In this paper, we use Powernext⁸ data due to its free access. The EU futures carbon contracts trade on the European Climate Exchange® (ECX) and we source this data from the ECX's website⁹. The time period covered is 24 June 2005–27 November 2006. The data on short-term Euribor interest rates and longer term swap rates is obtained from DataStream©. In order to match the interest rate maturity to the maturity of each of the futures contracts effectively, i.e. track the futures contracts through time, we interpolate the interest rates to maturities of every month within the data sample. For example, we match a two-year futures contract with a two-year swap interest rate in the first month of the contract life, a twenty-three month interest rate in the second month of its life and so on. Figure 1 depicts the evolution of the spot and futures prices over this time period.

Figure 1. Spot and Futures Time Series in natural logarithms: Jun '05–Nov '06.



⁸ www.powernext.com

⁹ www.ecx.com

We make two observations about the above graph. First, although the four series exhibit nonstationary behaviour, they appear to move in a relatively loose union. This observation supports our empirical modelling technique (i.e. cointegration framework) outlined in Section 3. The exception is the December 2008 futures contract that departs from the common trend at the end of 2005. The second observation is that we can see a sharp decline in all four price series over the period 24 April 2006–12 May 2006. The decline coincides with the release of the first official carbon emissions report that showed lower than expected carbon emissions over the 2005 reporting period. Thus, we will need to be careful to account for this break in our empirical model. Table 1 presents summary statistics for log return series.

Table 1. Summary statistics

	Spot	Dec-06	Dec-07	Dec-08
Mean	-0.276	-0.284	-0.277	-0.060
Std. Dev.	0.000	0.000	0.218	0.126
Median	4.867	5.066	5.047	3.585
Skewness	-1.887	0.779	0.633	-1.668
Kurtosis	28.197	36.517	34.744	18.537
Jarque-Bera p-value	0.000	0.000	0.000	0.000
Jarque-Bera statistic	9223.318	15995.750	14339.890	3588.071

Summary statistics are calculated for weekly log return $r_i = \ln(P_i/P_{i-1}) \times 100$ series over the period June 2005–November 2006.

We see that all four series exhibit negative average returns over the sample period. This observation is most likely due to large negative returns realised over the period 24 April 2006–12 March 2006 as indicated by the median measures which are less susceptible to extreme observations. The December 2008 futures contract appears to be least volatile followed by the spot contract and then the December 2007 and December 2006 futures. All four indices exhibit excess kurtosis (thick tails) and varying signs of skewness. In summary, all contracts

display significant departures from normality as indicated by the Jarque-Bera statistics and this is most likely due to large kurtosis.

4.1. Unit root and cointegration tests

As noted above, the price series appear to exhibit nonstationary behaviour, i.e. contain a unit root. We test this hypothesis more formally using two alternative tests: the Augmented Dickey Fuller (1979, ADF) unit root test and the Kwiatkowski *et al.* (1992, KPSS) unit root test. While the first procedure tests the null hypothesis of a unit root in the autoregressive representation of the series, the KPSS technique tests the null of stationarity. We conduct the tests on variables in natural logarithms and in first log differences. As indicated by the results presented in Table 2 below, all series appear to exhibit a unit root in log levels, while the first log differences are stationary at the 1% level.

Table 2. Unit root tests

	ADF	KPSS
Variables in (log) levels		
Spot	-2.648	1.376***
Futures Dec 06	-2.622	1.419***
Interest Rate (Matching Dec 06)	-2.476	1.793***
Futures Dec 07	-1.941	1.476***
Interest Rate (Matching Dec 07)	-2.401	1.583***
Futures Dec 08	-2.907	1.419***
Interest Rate (Matching Dec 08)	-1.798	1.706***
Variables in (log) differences		
Spot	-12.691***	0.114
Futures Dec 06	-8.840***	0.128
Interest Rate (Matching Dec 06)	-13.058***	0.213
Futures Dec 07	-8.989***	0.141
Interest Rate (Matching Dec 07)	-13.204***	0.194
Futures Dec 08	-11.002***	0.035
Interest Rate (Matching Dec 08)	-17.280***	0.206

Note: *** denotes statistical significance at 1% level, ** at 5% level and * at 10% level. Tests are performed with optimal lag lengths chosen by SIC criteria, including a trend and an intercept for variables in levels and an intercept for variables in differences. However, findings are not sensitive to alternative specifications.

Next we test for cointegration, i.e. a long-run relationship between each of the futures prices, the spot price and the interest rates. Cointegration tests are most commonly performed using maximum likelihood tests developed in Johansen (1988, 1991). In our study however we need to take particular care of accounting for possible trends and/or breaks predicted by the cost-of-carry model as well as visual analysis of the raw data in Figure 1. Potentially we have three breaks to model.

First, there is a level shift in all series related to price declines in April 2006 (the settlement month of the first year of carbon trading). Such a level shift is easy to model by including indicator dummies for the relevant observations. Second, a break associated with our hypothesis of zero-convenience yield in the last year of any carbon futures life needs to be included. This break was only applicable to the December 2006 contract at the time of writing this paper. For this case two sub-samples need to be considered: the first sub-sample from 24 June 2005–20 December 2005 and the second sub-sample from 3 January 2006–27 November 2006. Third, a graphical representation of the mis-pricing relationship presented in Appendix 1 (figures A.1–A.3.) suggest that all three futures contracts have undergone structural shifts in their relation to the spot contract in December 2005, i.e. the end of the first compliance period. It seems that market participants have significantly altered their expectations about the carbon price following this period. The same sub-samples as defined earlier will need to be considered. Thus we can anticipate a difficulty in separating the effects of the structural break associated with the end of the 2005 reporting period and the zero-convenience yield hypothesis for the December 2006 contract.

In order to test for cointegration in the presence of the three breaks/shifts we use Johansen, Mosconi, and Nielsen's (2000) maximum likelihood cointegration test method. This method generalizes the likelihood-based cointegration analysis developed by Johansen (1988, 1991) to the case where structural breaks exist at known points in time. For all the contracts and the two sub-samples defined above we consider that some or all of the time series follow a trending pattern in each sub-sample and the cointegrating relations are trend stationary in each sub-

sample; a trend break is allowed both in the cointegrating relations and in the variables in levels. A level shift is also allowed. We summarise the results of the Johansen, Mosconi, and Nielsen (2000) cointegration tests in Table 3, and present more detailed analysis in Appendix 2.

Table 3. Cointegration rank tests

Hypothesis	Trace Statistic	p-value*	95% simulated critical value	95% critical value Osterwald-Lenum (1992)
December 2006: Model with a trend break				
R = 0	73.071	0.000	35.845	25.731
R ≤ 1	12.963	0.228	18.478	12.448
December 2007: Model with a trend break				
R = 0	74.771	0.000	32.785	25.731
R ≤ 1	11.272	0.272	17.135	12.448
December 2008: Model with a trend break				
R = 0	22.964	0.482	34.779	25.731
R ≤ 1	6.936	0.721	18.513	12.448

Note: Johansen, Mosconi, and Nielsen (2000) tests for cointegration between a carbon futures contract, the spot price and interest rates. * p-value for the simulated distribution.

According to the Johansen, Mosconi, and Nielsen (2000) cointegration tests, December 2006 futures and spot prices and interest rates are cointegrated with one cointegrating vector. Similarly, December 2007 futures and spot prices and interest rates are cointegrated with one cointegrating vector but December 2008 futures price, the spot price and interest rates are not cointegrated.

5. Empirical findings

This section presents market-efficiency and price-discovery tests under three headings based on our analysis of the December 2006, December 2007 and December 2008 futures contract respectively.

5.1 December 2006 futures contract

We model the long-run relationship between the December 2006 futures price, the spot price and the interest rates by specifying a cointegrating vector that includes a linear trend restricted to the cointegrating relationship and a trend break dated

from January 2006 until the end of the sample. The dummy variable which takes the value 1 from January 2006 and 0 before is denoted by D_2 in Table 4.

Table 4. The Long-Run Models for December 2006 futures, spot, interest rates

The Cointegrating Vector (t-ratios in parentheses)	
s_t	-1.003 (-256.774)
f_t	1
r_t	-0.750 (-0.811)
$tD_{2,t}$	-0.002 (-0.117)
t	0.017 (1.668)
The speed of Adjustment Coefficients (t-ratios in parentheses)	
s_t	0.375 (1.041)
f_t	-0.686 (-1.742)

- LR test on the restriction that the coefficient of the trend is equal to minus the trend break coefficient: χ^2 statistic = 0.359, p-value = [0.549].
- LR test on the cost-of-carry restrictions: $H_0: a = b = -1$
 χ^2 statistic = 0.469 p-value = [0.791]

The cointegrating vector takes the following form $f_t = as_t + b(r_t(T-t)) + dt + d_2tD_{2,t} + \xi_t$ where the dummy variable $D_{2,t}$ is defined in Appendix 2.

Although we cannot reject the null of cost-of-carry model according to the LR test presented in the above table, we argue that this is most likely due to the large standard error associated with the interest rate variable. Although the joint test of parameters on s_t and r_t variables being equal to -1 is not rejected, as shown in Table 4, we also cannot reject the null that the coefficient on r_t is equal to zero. The coefficients of the trend and the dummy variable times the trend, are not significantly different from zero either. These findings imply that the convenience yield d in equation (4) is zero over the whole sample, with the dummy variable showing no difference between the first part of the sample and the post-January 2006 period. However, given that the structural break identified in the previous section occurs at about the same time it is difficult to interpret our results on the trend break variable. We thus cautiously conclude that although the December

2006 futures contract price, the spot contract price and the risk-free interest rate are linked in the long-run, the long-run relation is not given by a cost-of-carry model, and that there is no evidence of convenience yield in this pricing equation.

We now turn to Mosconi and Giannini (1992) and Toda and Yamamoto (1995) Granger Causality tests which will shed light on the price discovery process in the market for carbon permits. Tables 5 and 6 present these tests.

Table 5. Mosconi and Giannini (1992) Causality Tests for December 2006

R1	test-statistic	p-value
Null Hypothesis: Spot is not caused by Dec 06 Futures		
0.000	58.850	0.000
1.000	16.810	0.030
Null Hypothesis: Dec 06 Futures is not caused by Spot		
0.000	137.430	0.000
1.000	92.700	0.000

Table 6. Toda and Yamamoto (1995) Causality Tests for December 2006

Null Hypothesis	Chi-Squared (4) test statistic	p-value
Spot is not caused by Dec 2006 Futures	27.325	0.000
Dec 2006 Futures is not caused by Spot	14.289	0.014

According to the above tables we find bi-directional causality between the December 2006 carbon futures price and the carbon spot price and thus conclude that price discovery occurs in both the spot and futures markets. In contrast with these results, the volatility spillover tests presented below suggest that there are no statistically significant volatility transfers between the 2006 futures contract and the spot price.

Table 7. Volatility Spillover between December 2006 futures and spot contracts

Volatility Spillover	ARCH	GARCH
Spot to Futures	0.002 (0.010)	-0.047 (-0.125)
Futures to Spot	0.222 (0.725)	0.020 (0.049)

T-ratios (in brackets) are calculated using Bollerslev-Wooldridge (1992) robust standard errors.

5.2 *December 2007 futures contract*

Cointegration tests presented in Table 3 suggested that there is a long-run cointegrating relationship between December 2007 futures contract, the spot futures price and interest rates. Estimates of this relationship are presented in Table 8 below. The first restrictions we would like to test involve the coefficients on the interest rate and the spot variable. In particular, in order to be consistent with the cost-of-carry model these coefficients need to be equal to one. This restriction is tested by using a likelihood ratio statistic with a $\chi^2(2)$ distribution. This restriction is strongly rejected (pvalue = 0.011). We thus conclude that although there is a long-run link between the December 2007 futures contract and the spot carbon price (and hence the futures contract can be effectively used for hedging purposes), the link is not given by the cost-of-carry model.

Table 8. The Long-Run Model: December 2007 futures, spot, interest rates

The Cointegrating Vector (t-ratios in parentheses)	
s_t	-0.991 (-135.013)
f_t	1
r_t	-3.053 (-4.906)
$tD_{2,t-k}$	-0.077 (-3.075)
t	0.111 (5.836)
The speed of Adjustment Coefficients (t-ratios in parentheses)	
s_t	-0.084 (-0.405)
f_t	-0.689 (-3.380)

- LR test on the restriction that the coefficient of the trend is equal to minus trend break coefficient $\chi^2(1) = 4.178$, p-value = [0.041]
- LR test on the cost-of-carry restrictions: $H_0: a = b = -1$ χ^2 statistic = 9.073
p-value = [0.011]

The cointegrating vector takes the following form: $f_t = as_t + b(r_t(T-t)) + dt + d_2tD_{2,t} + \xi_t$ where the dummy variable $D_{2,t}$ is defined in Appendix 2.

The likelihood ratio test of the cost-of-carry model strongly rejects the null of cost-of-carry relationship for the December 2007 contract. This departure from the cost-of-carry is an indicator of unexploited arbitrage opportunities associated with the December 2007 contract over the period. Judging by the statistical significance of the coefficients on the time trend variable we can conclude that there is a statistically significant convenience yield associated with this contract, which becomes smaller following December 2005 but does not disappear completely. This is in line with our hypothesis in which the convenience yield is expected to disappear only after December 2006, i.e. with the start of the last year of this contract's life.

Mosconi and Giannini (1992) and Toda and Tamamoto (1995) causality tests indicate bi-directional causality between December 2007 futures and spot prices. These results are presented in Tables 9 and 10.

Table 9. Mosconi and Giannini (1992) Causality Tests for December 2007 futures contract

R1	test-statistic	p-value
Null Hypothesis: Spot is not caused by Dec 07 Futures		
0.000	71.730	0.000
1.000	35.070	0.000
Null Hypothesis: Dec 07 Futures is not caused by Spot		
0.000	64.500	0.000
1.000	16.530	0.000

Table 10. Toda and Yamamoto (1995) Causality Tests for December 2007

Null Hypothesis	Chi-Squared (4) test statistic	p-value
Spot is not caused by Dec 2008 Futures	24.419	0.000
Dec 2008 Futures is not caused by Spot	27.141	0.000

Lastly, volatility spillover tests given below indicate that volatility spills in both directions between the December 2007 futures contract and the spot price. This is in contrast to what we found in the case of the December 2006 futures contract.

Table 11. Volatility Spillover between December 2007 futures and spot contracts

Volatility Spillover	ARCH	GARCH
Spot to Futures	1.096 (9.351)	-0.397 (-2.418)
Futures to Spot	-0.681 (-9.341)	0.124 (1.003)

T-ratios (in brackets) are calculated from Bollerslev-Wooldridge (1992) robust standard errors.

5.3 December 2008 futures contract

Since our cointegration tests indicate that there are no long-run links between the December 2008 futures contract, the spot contract and the interest rates we only present Granger causality and volatility spillover analysis. Granger causality conducted in a vector autoregression framework on differenced log series suggest that the spot contract is caused by December 2008 futures return but not vice versa.

Table 12. Granger Causality Tests for December 2008 futures contract

Null Hypothesis	Chi-Squared (4) test statistic	p-value
Spot is not caused by Dec 2008 Futures	11.201	0.024
Dec 2008 Futures is not caused by Spot	3.003	0.557

Using VAR model in first difference including intercept break but no trend/trend break.

On the other hand, volatility spillover tests indicate that there are no statistically significant volatility spillovers between the two contracts.

Table 13. Volatility spillover between December 2008 futures and spot contracts

Volatility Spillover	ARCH	GARCH
Spot to Futures	0.066 (0.793)	-0.054 (-1.177)
Futures to Spot	0.016 (0.191)	-0.073 (-1.069)

T-ratios (in brackets) are calculated from Bollerslev-Wooldridge (1992) robust standard errors.

6. Conclusion

In this paper we have examined the issues of price discovery and market efficiency in the European Union carbon allowance market.

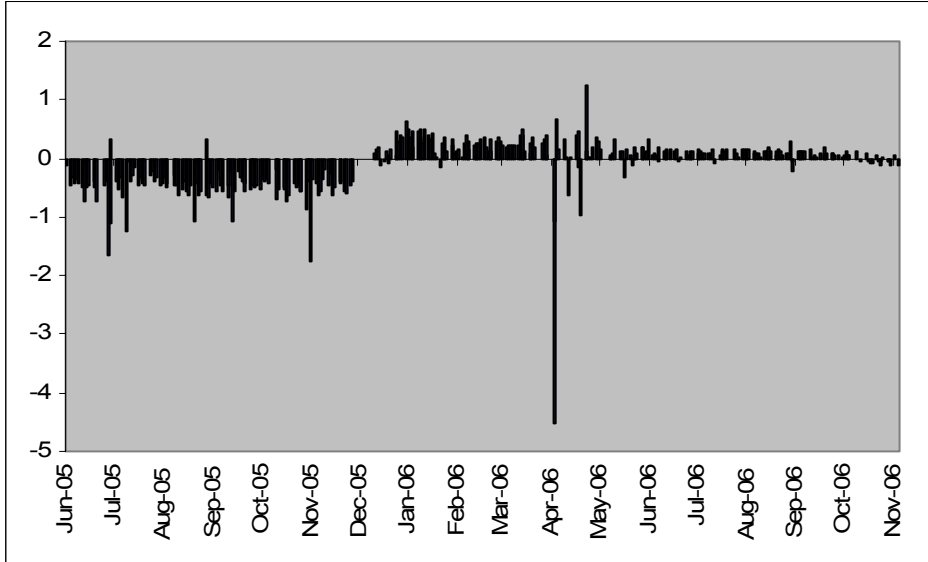
None of the carbon futures contracts studied here are priced according to the cost-of-carry model, although a likelihood ratio test for the December 2006 cannot reject the null of cost-of-carry relation. We argue that this is due to a large standard error associated with the estimated parameter on the interest rate variable. While the December 2006 and December 2007 futures contracts form a long-run link with the carbon spot contract, the December 2008 contract does not. We find bi-directional Granger causality between the December 2006 futures and the spot contract using Mosconi and Giannini (1992) and Toda and Yamamoto (1995) tests. After estimating the conditional covariance matrix on VECM residuals we find no volatility spillovers between the December 2006 futures contract and the spot contract. Granger causality tests as well as a volatility spillover test detect bi-directional causality/volatility spillovers between the December 2007 futures and the spot contract. Although there is no long-run link between the December 2008 futures price, the spot price and the interest rates, and we find no volatility spillovers between the December 2008 futures and spot contracts, this futures contract Granger causes the spot return.

Our findings have important implications for a number of financial applications including carbon arbitrage trading, risk management and portfolio selection and provide directions for future research. For example, the finding that none of the futures contracts follows a cost-of-carry relationship with the spot price and interest rates suggests the existence of arbitrage opportunities in the market for carbon permits. This proposition needs to be tested on historical data possibly using simulated trading strategies; we leave this as a topic of future research. Similarly, our finding that both the December 2006 and December 2007 contracts form a long-run relationship with interest rates and the spot price implies that these two contracts can both be used effectively for risk mitigation purposes. Again, in order to test this empirically one would need to form a hedge position, perhaps utilising dynamic hedge ratios within the GARCH framework, and

measure how effective the hedge positions are. Further, as the December 2008 futures contract does not form a long-run relationship with the interest rates and the Phase I carbon spot price, we suggest that this contract should not be used for hedging pre-2008 exposures in the spot market. We explain this finding by observing that the current (Phase I of the EU ETS) spot contract does not provide a carbon price relevant to the 2008–2012 period (Phase II). Thus, the December 2008 contract should be used for hedging post 2008 carbon exposures and in fact fulfils the important role of spot price discovery for the Phase II period.

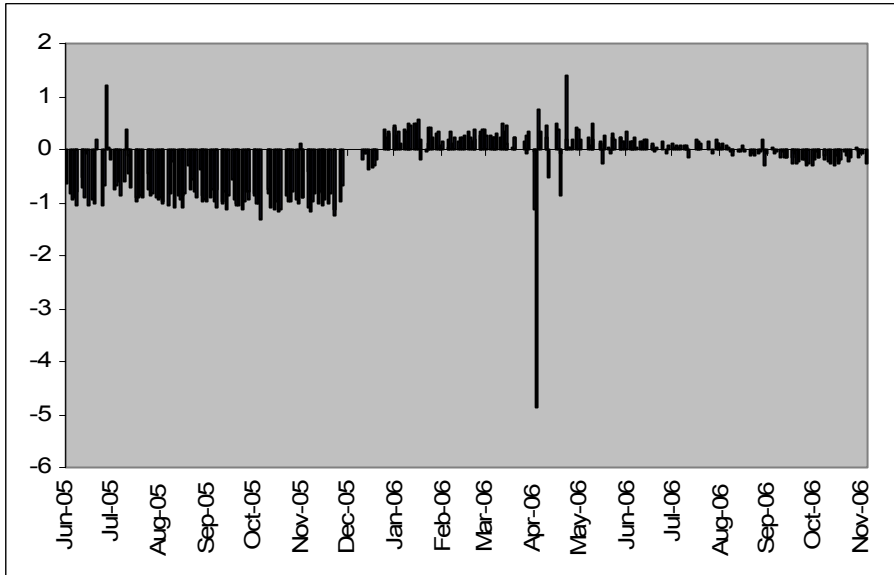
Appendix 1: Empirical mispricings from the cost-of-carry model

Figure A.1: Mispricing Relationship – December 2006 carbon futures contract



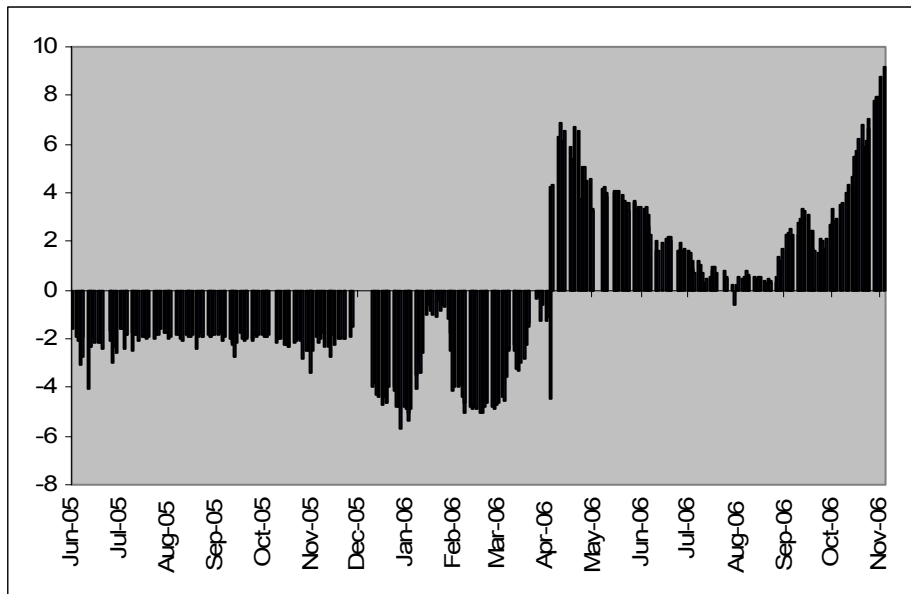
Mispricing is defined as a deviation of the realised futures price from its cost-of-carry price given by $Mispricing_t = F_{t,T} - S_t e^{(r_t - \delta)(T-t)}$.

Figure A.2: Mispricing Relationship – December 2007 carbon futures contract



Mispricing is defined as a deviation of the realised futures price from its cost-of-carry price given by $Mispricing_t = F_{t,T} - S_t e^{(r_t - \delta)(T-t)}$.

Figure A.3: Mispricing Relationship – December 2008 carbon futures contract



Mispricing is defined as a deviation of the realised futures price from its cost-of-carry price given by $Mispricing_t = F_{t,T} - S_t e^{(r_t - \delta)(T-t)}$.

Appendix 2: Cointegration test diagnostics

Cointegration analysis of a (3×1) process $\mathbf{Y}_t' = (s_t, f_t, r_t)$ is generally conducted (without taking account of structural breaks) in a VECM framework:

$$\Delta \mathbf{Y}_t = \mathbf{\Pi} \mathbf{Y}_{t-1} + \mathbf{\Pi}_1 t + \boldsymbol{\mu} + \sum_{i=1}^{k-1} \mathbf{\Gamma}_i \Delta \mathbf{Y}_{t-i} + \boldsymbol{\varepsilon}_t \quad (\text{A.1})$$

where $\boldsymbol{\varepsilon}_t$ is a normal, independent and identically distributed (3×1) vector stochastic process with zero mean and a variance matrix $\boldsymbol{\Omega}$. We assume that although some or all of the three time series in \mathbf{Y}_t may have a time trend, none have a quadratic trend. The hypothesis of cointegration can then be formulated as a reduced rank problem of the $\mathbf{\Pi}$ matrix, in which case $\mathbf{\Pi} = \boldsymbol{\alpha} \boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $(3 \times r)$ full rank matrices, and \mathbf{Y}_t has a quadratic trend. If none of the p time series displays a quadratic trend it is necessary to assume that $\mathbf{\Pi}_1 = \boldsymbol{\alpha} \boldsymbol{\gamma}'$, where $\boldsymbol{\gamma}$ is a $(1 \times r)$ full rank matrix.

On the other hand, if one break is present (and consequently there are 2 sub-samples), conditionally on the first k observations of each sub-sample the model can be rewritten as 2 equations:

$$\Delta \mathbf{Y}_t = (\mathbf{\Pi} \ \mathbf{\Pi}_j) \begin{pmatrix} \mathbf{Y}_{t-1} \\ t \end{pmatrix} + \boldsymbol{\mu}_j + \sum_{i=1}^{k-1} \mathbf{\Gamma}_i \Delta \mathbf{Y}_{t-i} + \boldsymbol{\varepsilon}_t \quad (\text{A.2})$$

where $j = 1, 2$ and $\mathbf{\Pi}_j$ and $\boldsymbol{\mu}_j$ are (3×1) vectors. Under the null hypothesis of cointegration, the trend is restricted to the cointegrating relationships to exclude the possibility of quadratic trends in any time series. This means that $\mathbf{\Pi}_j = \boldsymbol{\alpha} \boldsymbol{\gamma}_j'$. However, instead of writing 2 equations, we define the following matrices: $\mathbf{D}_t = (1 \ D_{2,t})'$, $\boldsymbol{\mu} = (\mu_1 \ \mu_2)'$, $\boldsymbol{\gamma} = (\gamma_1' \ \gamma_2)'$ of dimensions $(2 \times 1), (3 \times 2), (2 \times r)$ respectively. Equation (A.2) can then be rewritten as:

$$\Delta \mathbf{Y}_t = \boldsymbol{\alpha} \begin{pmatrix} \boldsymbol{\beta} \\ \boldsymbol{\gamma} \end{pmatrix}' \begin{pmatrix} \mathbf{Y}_{t-1} \\ t\mathbf{D}_{t-k} \end{pmatrix} + \boldsymbol{\mu}\mathbf{D}_{t-k} + \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \Delta \mathbf{Y}_{t-i} + \sum_{i=0}^{k-1} \boldsymbol{\kappa}_i \mathbf{I}_{t-i} + \boldsymbol{\varepsilon}_t \quad (\text{A.3})$$

where the $\boldsymbol{\kappa}_i$ are (3×1) vectors and the dummy variables $D_{2,t}$, and I_t are defined as:

$$D_{2,t} = \begin{cases} 1 & \text{January 1}^{st} 2006 \leq t \leq T \\ 0 & \text{otherwise} \end{cases} \quad (\text{A.4})$$

Correspondingly indicator dummies for the break points can be defined:

$$I_t = \begin{cases} 1 & \text{for } t = \text{first observation after January 1}^{st} 2006 \\ 0 & \text{otherwise} \end{cases} \quad (\text{A.5})$$

Having formulated our specification as in above, we apply Johansen, Mosconi, and Nielsen (2000) maximum likelihood cointegration test method which is based on the squared sample canonical correlations, $\hat{\boldsymbol{\lambda}}_g$ of $\Delta \mathbf{Y}_t$ and $(\mathbf{Y}'_{t-1} \quad t\mathbf{D}'_{t-k})$ corrected for the regressors:

$$\mathbf{D}_{t-k}, \Delta \mathbf{Y}_{t-i}, \quad i = 1, \dots, k-1, \quad \mathbf{I}_{t-i}, \quad i = 0, \dots, k-1.$$

The likelihood ratio test statistic for the hypothesis of at most r cointegrating relations is given by:

$$LR = -T \sum_{i=r+1}^3 \ln(1 - \hat{\lambda}_g) \quad (\text{A.6})$$

In our case, we consider that some or all of the time series follow a trending pattern in each sub-sample and the cointegrating relations are trend stationary in each sub-sample; trend breaks are allowed both in the cointegrating relations and in the non-stationary series. We also include additional indicator dummy variables¹⁰ that correspond to the observations around the end of April 2006.

¹⁰ Since these dummies are indicator dummies they do not affect the critical values of the usual cointegration tests.

The maximum lag length k is selected on the basis of the usual information criteria and the autocorrelation tests. For the December 2006 futures k is selected to be 5. The tests¹¹ for cointegration are reported in Table A.1 using two possible specifications: a trend break restricted to the cointegrating relationship and an intercept break restricted to the cointegrating relationship. Both the 95% simulated critical values (Johansen *et al.*, 2000) and the Osterwald-Lenum (1992) critical values are presented. As expected, the simulated critical values are much larger, due to the use of indicator variables. Using the Osterwald-Lenum (1992) critical values, we would actually conclude that the spot and futures prices are stationary. Using the simulated critical values we conclude, at the 5% significance level, that there is only one cointegrating vector. Having tested for cointegration we also test the interest rate for weak exogeneity (Table A.2). As mentioned in Section 3, we would not expect the interest rate to endogenous with respect to the carbon prices. We find that the interest rate variable is indeed weakly exogenous and thus re-specify the model without the r_t equation as such. Single stationarity tests of VECM residuals are reported in Table A.3 and normality tests in Table A.4. The unit root tests are carried out by testing restrictions on the cointegrating vectors. In all cases, the tests were done under the assumption that the trend breaks were present in the cointegrating relationships (as expected, excluding the trends does decrease the p-values dramatically). In all cases, we reject the null of stationarity at the 1 percent level. Table A.4. indicates some problems with non-normality.

Table A.1. Rank tests in CATS (version 2)

Hypothesis	Trace Statistic	p-value ^a	95% simulated critical value	95% critical value Osterwald-Lenum (1992)
Model with a Trend Break				
R = 0	73.071	0.000	35.845	25.731
R ≤ 1	12.963	0.228	18.478	12.448
Model with Intercept Break				
R = 0	71.036	0.000	30.309	20.164
R ≤ 1	12.218	0.148	15.603	9.142

^a p-value for the simulated distribution

¹¹ We conduct cointegration analysis in CATS in RATS version 2 (Dennis, 2006).

Table A.2. Weak exogeneity test $\chi^2(1)$ (p-values in brackets)

r_t
1.407 [0.236]

The test is done under the assumption that $r = 1$ in the model with a trend break.

Table A.3. Stationarity Tests (p-values in brackets)

Variable	$\chi^2(1)$
Model with Trend Break	
s_t	44.382 [0.000]
f_t	43.767 [0.000]
Model with Intercept Break	
s_t	42.820 [0.000]
f_t	42.114 [0.000]

All tests are done under the assumption that $r = 1$.

Table A.4. Jarque-Bera Normality Tests (p-values in brackets)

Equation	Skewness	Kurtosis	Normality
Model with Trend Break			
s_t	0.860	6.644	53.465 [0.000]
f_t	1.144	7.886	62.346 [0.000]
Model with Intercept Break			
s_t	-0.948	6.842	53.106 [0.000]
f_t	-1.202	8.172	64.367 [0.000]

The analysis is repeated for the December 2007 ($k = 3$) and December 2008 ($k = 4$) futures contracts in Tables 5 through 10.

Table A.5. December-2007 Futures Rank Tests in CATS (simulated critical values)

Hypothesis	Trace Statistic	p-value ^a	95% simulated critical value	95% critical value Osterwald-Lenum (1992)
Model with Trend Break				
$R = 0$	74.771	0.000	32.785	25.731
$R \leq 1$	11.272	0.272	17.135	12.448

^a p-value for the simulated distribution

Table A.6. Weak exogeneity test $\chi^2(1)$ (p-values in brackets)

r_t
0.663 [0.416]

The test is done under the assumption that $r = 1$ in the model with a trend break.

Table A.7. Jarque-Bera Normality Tests (p-values in brackets)

Equation	Skewness	Kurtosis	Normality
Model with Trend Break			
s_t	-0.851	6.752	56.559 [0.000]
f_t	-0.780	7.768	90.556 [0.000]

Table A.8. Stationarity tests (p-values in brackets)

Variable	$\chi^2(1)$
Model with Trend Break	
s_t	51.387 [0.000]
f_t	50.704 [0.000]

All tests are done under the assumption that $r = 1$.

Table A.9. December 2008 Futures Rank Tests in CATS (simulated critical values)

Hypothesis	Trace Statistic	p-value ^a	95% simulated critical value	95% critical value Osterwald-Lenum (1992)
Model with Trend Break				
$R = 0$	22.964	0.482	34.779	25.731
$R \leq 1$	6.936	0.721	18.513	12.448

^a p-value for the simulated distribution

Table A.10. Jarque-Bera Normality Tests (p-values in brackets)

Equation	Skewness	Kurtosis	Normality
Model with Trend Break			
s_t	-0.851	6.752	56.559 [0.000]
f_t	-0.780	7.768	90.556 [0.000]

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