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## The European Carbon Market in the Financial Turmoil: some empirics in early Phase II

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# The European Carbon Market in the Financial Turmoil: some empirics in early Phase II*


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**Abstract**

We estimate an Error-Correction Model by using dynamic OLS to investigate carbon price drivers in early Phase II. The futures contract negotiated from January to December 2008 on the European Climate Exchange is the focus of our analysis. We consider the allowance price as explained by oil prices, the switching price and the Dow Jones Euro Stoxx 50. The long-term cointegration analysis shows that oil was the main driver of carbon prices in 2008. Technological variables, although statistically significant, had almost no impact on the endogenous variable. The financial index has not been a statistically significant regressor. We find an adjustment speed of 8% in the cointegrating equation. The short-term estimates show a two-tier relationship. Before the financial and economic turmoil, energy inputs were the drivers of carbon prices more than financial assets. After the oil crisis, carbon markets have become sensitive to equity pricing. Brent prices have halved their impact on permit prices. This kind of "equity paradox" in $CO_2$ price drivers represents a new finding in carbon market pricing.


## 1 Introduction

EU Directive 2003/87/EC has established a scheme for trading greenhouse gas emission allowances within the Community (henceforth EU-ETS). The Pilot Phase of the system covered the period


*This paper has benefited from comments by M. Grasso, M. Manera, G. Mascolo and IEFE seminar participants. All errors remain our own responsibility.


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2005-2007. Since January 2008, the scheme has entered Phase II, which will end in 2012 with the achievement of the Kyoto target.

Much has been written so far on the EU-ETS in Phase I, despite the first period was meant to be a learning process. Mansanet-Bataller *et al.* 2008, Convery 2009, Convery *et al.* 2008, and Ellerman *et al.* 2008 have discussed the main lessons learned from the Pilot Phase of the EU-ETS.[1] Together with the environmental effectiveness and the cost-efficiency of the instrument, academics have investigated carbon price patterns in 2005-2007 and discussed either their determinants (Mansanet-Bataller *et al.* 2007; Alberola *et al.* 2008, 2009a, 2009b) or the most suitable stochastic behaviours to forecast such patterns (Benz *et al.* 2009; Daskalakis *et al.* 2007; Paolella *et al.* 2007; Seifert *et al.* 2008).

Very little is known, instead, about Phase II. To the best of our knowledge, this is the first work that documents and analyses the economic and institutional context at the launch of the Kyoto Period in the European market for $CO_2$ permits.

Phase I and II of the EU-ETS differ in terms of market expertise, characteristics (liquidity and depth), and regulation. This is shown, for instance, by simply comparing the main trading indicators for the EU-ETS in 2008 with their counterparts in 2006, which we consider representative of Phase I (see Table I in the Appendix). With an open interest of 140.70 Mton and a daily average of around 5.6 Mton exchanged, 2008 has been a record value for EUAs (European Climate Exchange, 2009). The open interest for 2006 was 14.98 Mton and daily transactions amounted to approximately 1.80 Mton. Among the reasons that have led to this performance, we will mention lower volatility in carbon prices, as compared to 2006 (Point Carbon, 2009), and delays in the allocation procedures for Phase II. The former has attracted financial institutions, the latter has forced regulated installations to trade forward with the aim of reducing risk-exposure and thus uncertainty.

Given these differences, we believe that results for Phase I cannot be automatically extended to Phase II. The objective of our analysis is thus twofold. First, we will shed some light on the determinants of $CO_2$ prices in early Phase II, and we will test whether the carbon price drivers identified so far by the economic literature (Mansanet-Bataller *et al.*,2007; Alberola *et al.* 2008,

[1] These analyses agree in that the Pilot Phase was useful, despite some shortcomings, such as the initial over-allocation and the banking restrictions. The first one led to a structural break in 2006 after the first emission verification, while the second one created a vintage effect of pollution permits expiring in 2007, whose market value fell to zero at the end of the first phase. The overall experience is deemed successful, in that carbon now has a price that must be taken into account by regulated sectors. It is still unclear, however, whether this has led to abatement. Empirical evidence provided by Ellerman *et al.* (2007) and Trotignon (2008) indicates that carbon price has induced some emissions abatement, in the form of intra-fuel substitution (brown to hard coal) in Germany, and improved $CO_2$ efficiency in the UK. Carbon price has had a limited impact on industrial competitiveness, as reported by Demailly et al. (2008). Finally, in an international perspective, the EU-ETS has had external impacts linking different carbon trading mechanisms and has stimulated the flexibility mechanisms included in the Kyoto Protocol .

2009a, 2009b) still hold for the EU-ETS in 2008. Second, our analysis is a bridge between the industrial and financial literature on carbon permits. We will discuss why and to which extent $CO_2$ allowances are a new asset class by investigating whether the erratic behaviour of carbon prices is statistically correlated to the turmoil characterizing financial markets. How emissions in Phase II are affected by the economic and financial recession is indeed a question that actually deserves attention.

We assume that the most reliable indicator for the pollution permit system in 2008 is the European futures market, with the EUA contract expiring in December 2008 being the focus of our analysis. The spot market, instead, was rather illiquid at the beginning of Phase II and is therefore less meaningful to be investigated in our framework. We consider the EUA price as explained by oil prices, the switching price,[2] and, finally, the most important European equity index, the Dow Jones Euro Stoxx 50.

We estimate an Error-Correction Model (henceforth ECM) that combines the long-run cointegrating relationship between levels variables and their short-run relationship. Applied works employing ECM formulations have been mainly used to test for the causal chains implied by the major paradigms in macroeconomic theory (see Masih *et al.* 1995a, 1995b and 1996b for a survey). In the context of $CO_2$ markets, but with a different perspective from ours, ECM models – and especially their generalization with VECM models (Vector ECM models which involve a system of linear cointegration relationships) – have been applied. The VECM approach has been recently used either to test the cointegrating relationship between the price of carbon allowances and the price of energy fuels (including electricity) in the UK and Germany (see Bunn *et al.* 2007, 2009), or the cointegration between spot and future prices in the EU-ETS (Milounovich *et al.* 2007) or in different marketplaces, namely the EU-ETS and the Chicago Climate Exchange.

Since an ECM tests the long-run cointegrating relationship, we would like to stress that this kind of econometric representation becomes meaningless when applied to a time frame different from the year 2008. In fact, the EUA expiring in December 2009 is a different product, therefore including forward observations would bias results. Moreover, as Phase I was not temporally linked to Phase II and was designed as a learning period, we believe that including backward data – that is the price of the EUA December 2008 contract traded in the period 2005-2007 – in the analysis would be misleading. To rule out the possibility of more than one cointegrating relationship – which would have allowed a VECM model –, we have carried out Joahnsen tests whose results indicate that the variables of interest are weakly exogenous. To take into account the possible simultaneity and small-sample biases and obtain consistent estimates in the ECM model, we have used the dynamic OLS technique (Stock and Watson, 1983).

The long-term cointegration analysis shows that oil was the main driver of carbon prices in 2008.

[2] The switching threshold is the theoretical price of allowances which leads to a power operator being indifferent to producing with gas or coal plants.

Technological variables, although statistically significant, had almost no impact on the endogenous variable. Over the full sample period, the financial index has not been a statistically significant regressor. We find an adjustment speed of 8% in the cointegrating equation, which is a noticeable result given the small sample size. The short-term estimates show a two-tier relationship. Before the financial turmoil, energy inputs were the drivers of carbon prices more than financial assets. After the oil crisis, with the economy entering a slowdown, carbon markets have become sensitive to equity pricing. Brent prices have halved their impact on permit prices. This kind of "equity paradox" in $CO_2$ price drivers represents a new (and probably conjunctural) finding in carbon market pricing.

This paper is organized as follows. After presenting the institutional context and the most relevant indicators (Section 2), we will review the empirical literature on carbon price drivers and detail our modelling assumptions (Section 3). The model follows (Section 4), with the description of the dataset (Section 4.1) as well as of the econometric specification (Section 4.2). Results are in Section 5. We will conclude by pointing out the main lesson learned by $CO_2$ trading in 2008. The Appendix complements the statistical analysis; robustness checks were also performed by using alternative estimation techniques (the Autoregressive Distributed Lag Method by Pesaran *et al.*, 1997).

## 2 Carbon Market in early Phase II

As noted in the introduction, 2008 was characterized by an upsurge of carbon forward trading. Daily transactions have remained above 1.80 Mton. Carbon prices have started to decrease as from August 2008. This falling trend resulted in a closing price of 15€/$tCO_2$ when the EUA 2008 expired. The same downward pressure was experienced by future contracts for delivery in 2009, 2010, 2011, and 2012 (Figure 1).

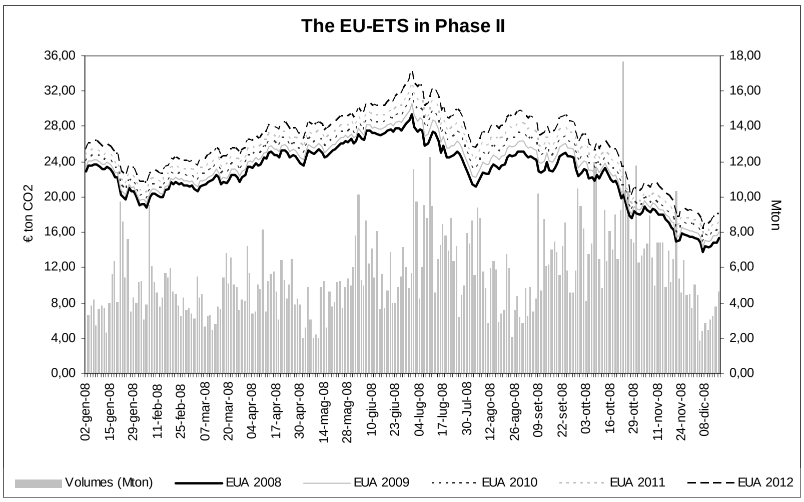


Figure 1. The EU-ETS in Phase II. Source: European climate Exchange, 2009.

The total cap in the EU-ETS Phase II is 2,093Mt$CO_2$/year on average. This includes allocation to existing installations (1,89 Mt$CO_2$) and reserves (197 Mt).[3] As of January 2009, 22 countries have received the European Commission's approval for their national allocation plans, while 23 countries, representing 86% of the total EU-ETS allocation in 2008, have had their installation lists approved. The countries awaiting the final approval of their installation lists are Bulgaria, Cyprus, Hungary, Poland, and Norway. The allocation process has been very lengthy. According to Tendances Carbone, in March 2008, of the overall cap of 2,10 Mt$CO_2$ of allowances to be issued per year, only 550 Mt$CO_2$ were allocated. At the end of 2008, the annual allowances issued amounted to 1,64 Mt$CO_2$. Actually, only minor changes can occur in the overall allocation of allowances and credits. Such changes can only take place in the five countries awaiting the final approval of their national allocation plans or pending court cases.

In nine countries, a share of allowances will be auctioned or sold, totalling 70 Mt$CO_2$/year or 3 percent of the total EU-ETS allocation, well below the 10 percent limit in the directive. In 2008, the only governments to sell allowances were Germany and the UK, with each choosing different approaches to the process. Whereas in Germany a bank was tasked with selling allowances on a daily basis in the brokered market and through exchanges, in the UK a centralized auction was held.

On sector level, the power and heat sector bears the highest reduction burden, while the oil and gas sector is the only industry not to have excess allowances.

To complete the institutional context, let us also mention that the legislative proposal for the post-Kyoto period, reinforcing the role of carbon markets, was presented in January 2008 and finally approved in December. It is also worth noting that since August 2008, credits from emission reduction projects carried out by using Clean Development Mechanisms have been officially quoted. However, the formal interconnection between the European Emissions Registry and the United Nation one was completed only at the end of October 2008.

Emissions reported in 2007, and published in April 2008, were 2,20 Mt$CO_2$, indicating a slight increase from the previous year. Market analysts anticipate 2008 emissions in the EU-27 to have decreased by 4.1%, or 90 Mt$CO_2$, to 2,11 Mt$CO_2$ (Société Générale, 2009). $CO_2$ emissions from EU-ETS installations should be 3% down from the previous year, according to New Carbon Finance (2009), which has released estimates two months ahead of the official release of verified emissions data for the EU by the European Commission that will come in April 2009. Although European power production has remained unchanged overall (+0.2%), emissions from utilities have fallen due to greater use of nuclear power, good hydro availability and the further deployment of renewables (Tendances Carbone, 2009). The drop in emissions has also been driven by substantial cutbacks

[3] The size of the New Entrant Reserve (NER) is estimated at 118 Mt/year, or around 6 percent of the overall allocation. The Joint Implementation Reserve totals 9.8 Mt/year, and only six countries have established such reserves. The total credit import potential is around 281 Mt/year (1,41 Mt over Phase II), representing 13 percent of the EU's total allocation.

in industrial production (almost -11%) during the second half of 2007. Even though this points to a EUA shortfall of 30 $MtCO_2$ in 2008, in 2009 it could be possible to have an over-allocation of 60 $MtCO_2$ driven by a further emissions drop of 90 $MtCO_2$, reaching a level of 2,020 Mt (Point Carbon, 2009).

## 3 The literature on carbon price drivers

The literature on the determinants of carbon prices in Phase I is relatively extensive. Academics have started investigating this issue by the early stages of the regulatory process which has led to the entering into force of the EU-ETS (Springer 2003; Christiansen *et al.* 2005).[4] This Section, which reviews the industrial and financial approaches to the issue, provides a better understanding of the originality of the econometric assessments in this paper.[5]

The focus of the Industrial Organization literature has been the determinants of carbon prices, which may be grouped into the following categories: *institutional factors* (i.e. tightness of national targets, sectoral coverage, temporal/spatial restrictions to trading, etc.),[6] *market structure* (i.e. macroeconomic factors - such as GDP level and economic growth - and energy variables - namely the price of energy sources, energy substitutability and the costs of abatement efforts/technologies)[7] and *exogenous factors* (i.e. temperatures, precipitations, clouds and wind speed).[8]

The models most closely related to ours are Mansanet-Bataller *et al.* (2007) and Alberola *et al.* (2008, 2009a, 2009b). The former was a pioneer in providing an empirical assessment of theoretical expectations. In particular, Mansanet-Bataller *et al.* (2007) have carried out a multivariate regression OLS model (corrected by the Newey-West covariance matrix, to deal with heteroskedasticity and autocorrelation) running from the formal launch of the EU-ETS, $1^{st}$ January 2005, to the $30^{th}$ November 2005. According to this analysis, energy prices are the determinants of EUA prices, whose explanatory power is around 50%. Average weather data are not statistically significant. Alberola *et al.* (2008) have substantially extended Mansanet-Bataller *et al.* (2007) by including the full Pilot Phase of the EU-ETS and looking at proxies for policy factors. Interestingly, Alberola *et al.* (2008)'s econometric specification explains one-third of EUA price patterns. Sub-periods investigations make clear that policy proxies are the main driver of carbon prices before

[4] Theoroetical models which analyse the determinants of $CO_2$ prices can be divided into five categories, depending on the approach adopted: (1) integrated assessment models, (2) computable general equilibrium models, (3) emission trading models, (4) neo-Keynesian macroeconomic models and (5) energy system models. For a more detailed description of each category see, among others, Springer (2003).

[5] The most updated review of industrial and financial models is in Bonacina *et al.* (2009).

[6] The issue is investigated in Springer *et al.* (2004), Nordhaus *et al.* (1999) and Nordhaus (2001).

[7] For further details on the argument see Kainuma *et al.* (1999), Van der Mensbrugghe (1998), Burniaux (2000), McKibbin *et al.* (1999), Ciorba (2001), Springer *et al.* (2004), Bahn *et al.* (1999, 2001), Kanudia *et al.* (1998), Zhang *et al.* (1998), Grubb *et al.* (1993), Sijm *et al.* (2000), Kanen (2006), Bunn *et al.* (2007) and Convery *et al.* (2007).

[8] For further assessments see, among the others Kainuma *et al.* (1999) and Ciorba *et al.* (2001).

the compliance break (June 2005-April 2006), while energy fundamentals govern carbon price trajectories henceforth (May 2006-April 2007). The role of market structure and industrial sectors is further investigated in Alberola *et al.* (2009a and 2009b) where paper, metal, coke oven, chemical, glass, cement and power (i.e. heating from electricity and gas) industrial production indexes are considered. Results are controversial as they depend on the level of the analysis (EU-27 versus country-specific) and the sample period (before versus after the compliance break).

The focus of the financial literature has been the forecasting of carbon price patterns. The issue has been investigated by Benz *et al.* (2009), Daskalakis *et al.* (2007), Paolella *et al.* (2007), Seifert *et al.* (2008) and Chevallier (forthcoming). The common root of this branch of research is that, as explained by Benz *et al.* (2006), "*there are several price determinants of $CO_2$ allowances which have stochastic behaviours (the changes of policy directives and regulations, the weather data, fuel prices, economic and sectorial growth, etc.). In particular, unexpected environmental events and sudden large variations in the fuels spreads represent the most affecting sources of EUA price uncertainty. As a consequence, $CO_2$ allowances prices and returns present a stochastic behaviour characterized by price jumps, spikes as well as phases of high volatility and heteroskedasticity in returns*". The large variety of models considered (AR-GARCH, regime-switching, Brownian augmented with jumps, GAt-GARCH, etc.) attests the difficulties encountered in analysing the complexity of the permit market due to the presence of several sources of uncertainty. However, one message is clear: carbon returns are not correlated with the returns of traditional financial assets. As a consequence, $CO_2$ allowances increase the diversification of a financial portfolio and reduce the overall investment risk (Mansanet-Bataller *et al.* 2008).

Analyses and findings concerning the EU-ETS in Phase I could not hold for modelling the carbon price in Phase II since the framework is actually different (i.e. enhancement of market maturity and liquidity, exclusion of banking restrictions, etc.), as also argued in Daskalakis *et al.* (2007). Indeed testing whether the carbon market determinants at the beginning of Phase II show continuity with the trends observed during Phase I is an interesting question that we will address. We will focus on the role of energy prices as determinants of the EUA price, as the literature has shown their overall importance. In particular, we are going to test whether the oil market and the coal-to-gas switching price - as a proxy to short-term abatement efforts - drive $CO_2$ prices. The oil market is included as it is the most mature and liquid product in the fossil fuel markets.

Regulatory uncertainty will not be explicitly included in our model, but this has motivated the choice of futures market over the spot market. In fact, the uncertainty on the National Allocation Plans, smoothly lasting throughout 2008, has prompted trades on the futures market. We share the interest in futures transactions with the carbon finance literature. Moreover, the topic of non-diversification will deserve special attention in our analysis. We will investigate whether the hedging property is likely to hold during the period of the financial crisis, at the end of 2008. This is of crucial importance when looking at the futures market, as we do, which encompasses trading activity by both regulated firms and financial intermediaries.

An innovative aspect of our analysis is that we will distinguish the long-term impact from the short-term impact of carbon price drivers. We believe that clarifying these drivers in a neat time frame provides a useful complement to the existing analyses that instead always pool them.

# 4 The dataset and the econometric specification

In what follows, we will describe data for $CO_2$ allowances, Brent, coal, natural gas, and Euro Stoxx prices used to determine the fundamentals of carbon prices in 2008 (Section 4.1); hence we will detail the econometric specification (Section 4.2).

## 4.1 The dataset

Since the installations covered by the EU-ETS have not an hourly or daily but a yearly need for allowances, we have opted to use futures prices to provide an assessment which moves apart from short-run changes in industrial expectations. The dataset encompasses five variables and consists of 241 observations.

**The carbon price.** The EUA price is determined on several markets: OTC (over-the-counter), spot, and futures markets. The first one is the most liquid, but price data are confidential. Among the remaining platforms where carbon allowances are traded, the most liquid is the ECX (European Climate Exchange). Therefore we will use the EUA 2008 price (*EUA* in €/ton of $CO_2$) negotiated from January $2^{nd}$ to December $15^{th}$ 2008 on the ECX. The sample period is consistent with the expiration of our endogenous variable. As for the descriptive statistics, the minimum is 13.72 €/ton of $CO_2$, the maxium is 29.33 €/ton of $CO_2$, with a standard deviation of 3.36.

**Energy prices.** As for energy market data, the following series are used. The oil price (*brent* in \$/barrel) is the daily Brent Front Month Contract quoted on the ICE's platform and referred to the North Sea hub. The natural gas price (*gas* in £/Btu) is the daily futures Summer 2009 price negotiated at the National Balancing Point.[9] The price of coal (*coal* in €/ton) is the daily Platt's Assessment on the API#2 CIF ARA contract and represents a price estimate of a representative quantity of standard steam coal delivered into the ARA range within 90 days. To ensure that all energy price series are traded with the same currency, both the oil and the natural gas price series are converted to € using the daily exchange rate provided by the European Central Bank. Figure 2 provides the plot of energy data (in €/MWh) over the sample period (Jan. $2^{nd}$ - Dec. $15^{th}$, 2008).

[9] We consider this gas price, since it is the most liquid gas trading point and it is the one that better represents the current market conditions, as usually gas is priced with a time lag of 6 to 9 months mainly with formulas that include baskets of crude and distillate products.

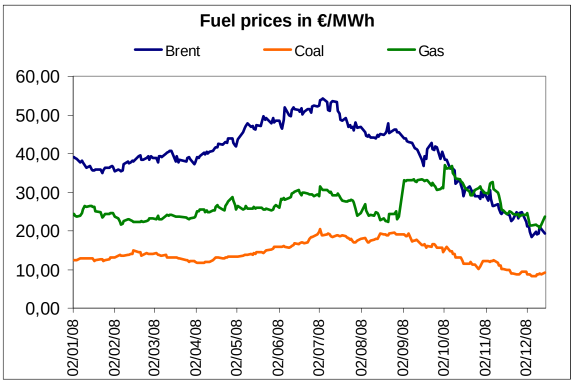


Figure 2. Energy data in 2008 (€/MWh).

Clearly, the *brent* price series is the most affected by the economic downturn as from August and shows the highest standard deviation, with a maximum that is three times its minimum (Table 1).

TABLE 1. STATISTICS FOR ENERGY DATA, 2008 (€/MWh)

| | Average | Std.Dev. | Min | Max |
|---|---|---|---|---|
| Brent | 39.81 | 8.52 | 18.46 | 54.21 |
| Coal | 14.36 | 2.94 | 8.26 | 20.38 |
| Gas | 26.86 | 3.63 | 20.67 | 36.96 |

Furthermore, to take into account abatement options for power producers, data on *coal* and *gas* are combined to get the switching price, *i.e.* the price of allowances at which the marginal costs of gas and coal-fired power plants are equal. As reported by Tendance Carbone (2007), in order to gauge the impact of fuel-switching, we need to make assumptions on the average plant efficiency and on the emission coefficients at the European level. These assumptions are reported in Table 2.

TABLE 2. TECHNOLOGICAL PARAMETERS

| | Efficiency | Emission coefficient |
|---|---|---|
| Coal | 36% | 0.86 $tCO_2$ |
| Gas | 50% | 0.36 $tCO_2$ |

Source: Tendence Carbone, 2007

Formally the switching price (*switch* in €/ton $CO_2$) is a result of the following relationship:

$$0.36 \times switch + 50\% \times gas = 0.86 \times switch + 36\% \times coal,$$

where the LHS measures the marginal cost of producing electricity with a gas-fired plant in a carbon-constrained framework and the RHS is the same for a coal-fired unit. Profit maximizing power producers use the least marginal cost technology. Assuming that each operator has the possibility to switch from high to low cost inputs, we get that fuel price differentials affect the technology used to produce electricity, and consequently fuel demand. During 2008, *gas* has always been above *coal*; thus leading to a non-negative *switch*. As for the relationship with $CO_2$ prices, except for January 2008, the EUA price has remained below the switching threshold thus signalling an economic advantage to use coal-fired power plants even if in carbon if in carbon-constrained frameworks (Figure 3).

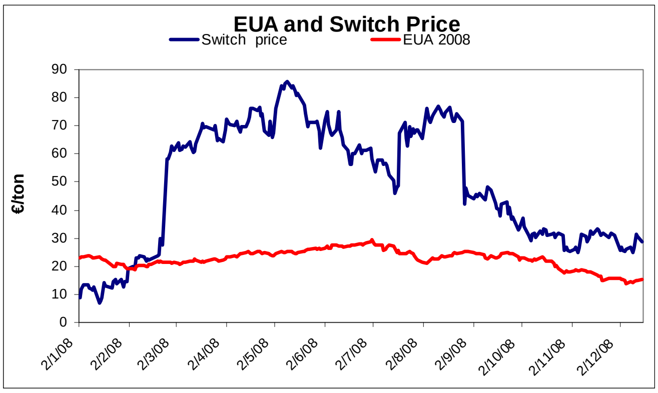


Figure 3. EUA prices and fuel-switching incentives in 2008.

**The financial index.** The Dow Jones Euro Stoxx50 Index (*eurostoxx* in €) is the Europe's leading equity index for the Eurozone and covers 50 stocks from 12 countries (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain). As noted in the introduction, the advantage from adding of *eurostoxx* is twofold. On the one side, the variable studies the likelihood of considering EUA 2008 as a financial asset. On the other, it controls for the economic downward trend that has characterized 2008 and, in particular, the second half of the year (Figure 4).

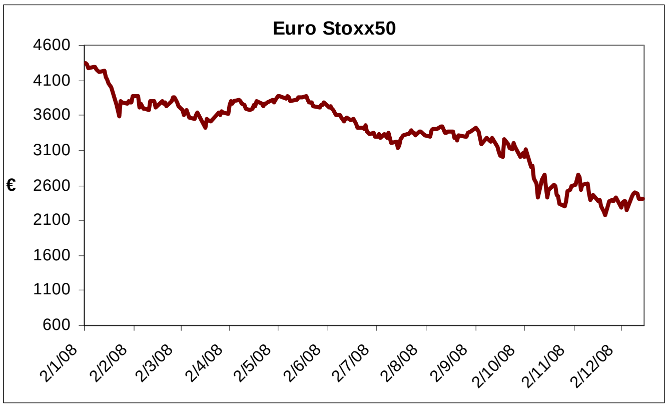


Figure 4. The Euro Stoxx Index in 2008.

The usual unit root test (Augmented Dikey Fuller, henceforth ADF) is performed for all price series. All of them are I(1) and converted to stationary taking first natural logarithm differences. *EUA, brent, switch* first logarithmic differences are not too highly correlated (Appendix, Table II). *EUA, brent, switch and eurostoxx* are not normal (the Jarque-Bera test rejects of the null hypothesis of normality at any reasonable level), their distributions are slightly platykurtic and negatively skewed.

### 4.2 Econometric specification: the error-correction model

The role played by energy and financial variables on EUA prices in both the short and the long-run is estimated using the ECM. Engle and Granger (1987) have demonstrated that once a number of variables are found to be cointegrated, there always exists an ECM implying that changes in the endogenous variable depend on the level of disequilibrium in the cointegrating relationship, which is captured by the error-correction term, as well as of changes in other explanatory variable(s). Therefore the long-run equilibrium acts as "attractor" towards which the system converges when there is a divergence from it due to non-stationarity (caused by stochastic trends). The cointegrating equations write as follows:

$$EUA_t = \alpha_0+\beta_0 brent_t+\gamma_0 switch_t+\delta_0 eurostoxx_t+\varepsilon_t \quad (1)$$

$$r_{t,EUA} = \alpha_1+\beta_1 r_{t,brent}+\gamma_1 r_{t,switch}+\delta_1 r_{t,eurostoxx}+\omega_1 v_{t-1}+\eta_t \quad (2)$$

where $t$ is the time period under consideration, $EUA_t$ is the log-EUA price series, $brent_t$ is the log-brent price series, $switch_t$ is the log-switch price series, $eurostoxx_t$ is the log-eurostoxx price series, $r_{t,\cdot}$ is the log-return of the relevant price series (i.e. $r_{t,i} = i_t/i_{t-1}$ and $i = EUA$, $brent$, $switch$, $eurostoxx$), $\varepsilon_t$ (and $\eta_t$) is the error term and $v_{t-1}$ is the error-correction term.

When variables are cointegrated, short-term deviations from long-term trends feed back on changes in the dependent variable to force the re-assessment towards the long-term path. As argued in the introduction, we assume that the EUA 2008 reaches the long-term equilibrium at the time of its expiry. Therefore our analysis provides information about the maturity of carbon markets. If $EUA$ is driven directly by this long-term equilibrium error, represented by $v_{t-1}$, then it is responding to this feedback; otherwise it is reacting to short-term shocks in the stochastic environment.

The significance tests of the differenced explanatory variables give us an indication of the short-term effects, whereas the long-term causal relationship is implied through the significance of the lagged error-correction term. This latter contains the long-term information since it is derived from the long-term cointegrating relationship. The coefficient of $v_{t-1}$, however, is a short-term adjustment coefficient and represents the proportion by which the long-term disequilibrium (or imbalance) in the dependent variable is being corrected in each short period. Non-significance or elimination of the lagged error-correction term affects the implied long-term relationship and may be a violation of theory. The non-significance of any of the differenced variables which reflect only a short-term relationship, however, does not involve such violations (Thomas 1993).

It is worth noting that, before choosing the two-equation model described by (2) and (3), we have tested the suitability of a VECM structure.[10] The usual cointegration tests (Johansen Max-Eigenvalue) have been applied. These tests, besides confirming the existence of one cointegrating vector in the majority of cases, have proved to be sensitive to lag specifications and trend assumptions, thus stressing the unsuitability of multi-equation techniques, due to the small size of the sample (as explained by Cheung, 1993).

Moreover, from an economic perspective, it seems more reasonable to find a long-term causal relationship where fuels, and eventually financial variables, are the main carbon price drivers. This is better investigated by an econometric technique, like the ECM, concentrated on a single equation, with EUA designated as the dependent variable, explained by other variables that are assumed to be weakly exogenous for the parameters of interest.

# 5 Results and interpretation

The full period results of eq.(1) are commented in Section (5.1); the same for eq.(2) are in Section (6). Please note that the quality of each regression is verified through several diagnostic tests (i.e. adjusted $R^2$, $p$-value of the $F$-test statistics, Akaike Information Criterion and Schwarz Criterion).

## 5.1 Estimation of the cointegration equation

To deal with the likely bias due to the sample's size and to correct for possible simultaneity bias amongst regressors, we will estimate eq.(1) by using the method proposed by Stock and Watson (1993): dynamic OLS (DOLS).[11] This parametric approach is suggested for estimating long-run equilibria in systems which may involve variables integrated of different orders but still cointegrated. The potential of simultaneity bias and small-sample bias among the regressors is dealt with by including lagged and led values. The procedure outperforms Phillips *et al.* (1991) and Saikkonen (1991) in that DOLS is much more practically convenient to implement and estimate. For variables that cointegrate, DOLS generates asymptotically efficient estimates of the regression coefficients (by using Newey-West standard errors and covariance), whose economic rationale can be exploited. Notice that DOLS does not estimate the short-run dynamics because this is not necessary for an asymptotically efficient estimation of the cointegrating relationship. The DOLS

[10] Further details on VECM are in Johansen (1988), Johansen and Juselius (1990), Juselius (2006).

[11] The ADF test related to the residuals of the long run-equation estimated by using standard OLS technique shows that ECM is not stationary at the 10% level (absence of cointegration). However, the error-correction test shows that the coefficient of the short-run equation is significant at the 5% level and negative. This discrepancy can be explained by the fact that ADF imposes a possibly invalid common factor restriction (Kremers *et al.*, 1993) which biases the estimation in small samples.

equivalent of eq.(1) is as follows:[12]

$$\begin{aligned} EUA_t = &\alpha_0 + \beta_0 brent_t + (\beta_1 r_{brent,t-1} + \beta_2 r_{brent,t} + \beta_3 r_{brent,t+1}) + \gamma_0 switch_t + \\ &+ (\gamma_1 r_{switch,t-1} + \gamma_2 r_{switch,t} + \gamma_3 r_{switch,t+1}) + \delta_0 eurostoxx_t + \\ &+ (\delta_1 r_{eurostoxx,t-1} + \delta_2 r_{eurostoxx,t} + \delta_3 r_{eurostoxx,t+1}) + \varepsilon_t \end{aligned} \quad (3)$$

Table 3 shows the results of eq.(3). The adjusted $R^2$ is above 90%.

TABLE 3. LONG-TERM RESULTS.

DOLS Long-run equation.

| | | | |
|---|---|---|---|
| $\alpha_0$ | 0.4032 (0.5615) | $r_{switch,t-1}$ | -0.0002 (0.0248) |
| $brent_t$ | 0.3896 (0.0481) | $r_{switch,t}$ | -0.0118 (0.039) |
| $switch_t$ | 0.0659*** (0.0101) | $r_{switch,t+1}$ | 0.0055 (0.0240) |
| $eurostoxx_t$ | 0.0889 (0.0878) | $r_{eurostoxx,t-1}$ | -0.2941** (0.1429) |
| $r_{brent,t-1}$ | -0.0822 (0.1397) | $r_{eurostoxx,t}$ | -0.2807* (0.1459) |
| $r_{brent,t}$ | -0.1422 (0.1522) | $r_{eurostoxx,t+1}$ | -0.4349*** (0.1323) |
| $r_{brent,t+1}$ | 0.0297 (0.1292) | | |

Main statistics

| | | | |
|---|---|---|---|
| $R^2$ | 0.9085 | Durbin Watson | 0.2593 |
| Adj. $R^2$ | 0.9036 | Akaike Info. Criteria | -3.1243 |
| F-stat[a] | 0.0000 | Schwarz Criterion | -2.9341 |
| ADF | -4.0632 | | |

[a]MacKinnon p-values: no-trend (240 obs.), 5% = -4.22; 10% = -3.85 (MacKinnon, 1991).

***, **, * indicate significance at 1%, 5% and 10% level, respectively.

The ADF test does not reject at the 10% significance level (MacKinnon, 1991) the null hypothesis of no residual cointegration for this model.[13] Unlike Kanen (2006) and Alberola *et al.* (2008), but consistently with economic intuitions, *brent* and *switch* variables are statistically significant at the 1% level. Their coefficients can be interpreted as long-term elasticities. Notice that the Brent coefficient $\hat{\beta}_0$ is quite large and positive. The rationale is that when oil prices increase, natural gas prices follow. Gas-fired units become more expensive, coal demand increases thus favouring the upsurge of carbon emissions and this in turn yields a higher price for allowances. The switching coefficient ($\hat{\gamma}_0$) is close to zero, because for most of the time in 2008 (see Section 4), coal has been less expensive than gas, and EUA has remained below the switching threshold. The carbon market

[12]The number of lags and leads is chosen using the Bayesian information criterion (Schwartz, 1978). The Schwarz Criterion suggests 0 leads and lags; however, to exploit the beneficial properties of the DOLS model one lead and one lag have been inserted in the long-run equation.

[13]We choose the option of MacKinnon p-values without trend as we assume that the economic trend is captured by the *eurostoxx* variable which has a more macroeconomic interpretation than the other regressors. The 10% level significance level may be due to the small (temporal) size of the sample

has resulted in no incentives for fuel-switching. Our results indicate that the market for allowances has come to maturity and reacts to fundamentals.

*Eurostoxx* is not statistically significant. This result is in line with both the financial (see Benz *et al.* 2006; Daskalakis *et al.* 2007; Paolella *et al.* 2007; Seifert *et al.* 2008) and the industrial (see Kanen 2006; Mansanet-Bataller *et al.* 2007; Alberola *et al.* 2008) literature on carbon markets. However the assessment of the short-term adjustment unveils different findings, as we will explain in Section 6.

## 6 Signals of downturn in the $CO_2$ market: the short-term adjustment

During 2008 *brent* and *eurostoxx* have shown very erratic behaviours. Within a semester they have experienced a cut of 65% and 40% respectively. In particular, *brent* has dropped from 54.21€/MWh in July 2008 to 18.46€/MWh in December, while *eurostoxx* has passed from an average of 3527.83€ in June to an average of 2165.91€ in December. The path of *EUA* has been quite similar. After reaching its high of 29.33€/ton in July, the average price of carbon permits has collapsed to 14.80€/ton in December: a price cut of 50%. The upsurge of erratic trends since mid-2008 is confirmed by the visual inspection of the one-week-rolling volatility of *brent*, *eurostoxx* and *EUA* (Figure 5).

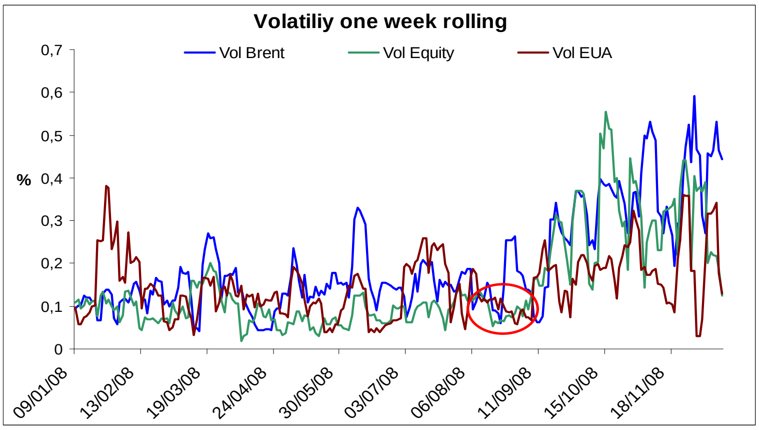


Figure 5. One-week rolling volatilities of EUA, brent and eurostoxx.

These indicators would have suggested the occurrence of a *structural break* in 2008 as a consequence of the oil downturn and the financial crisis. However, the Quandt-Andrew test has rejected such assumption at any reasonable level.

Even if the break does not seem to have fundamentally changed the market, the impact on *EUA* is non negligible. Since July $22^{nd}$, when *brent* decreased for the first time, both the correlation and the descriptive statistic for each variable have dramatically changed (Table IIIa and IIIb in the

Appendix). Given econometric and empirical evidences, we have parted the dataset in two sub-sample periods - *before the financial turmoil* ($2^{nd}$ January - $11^{th}$ August) and *after the financial turmoil* ($18^{th}$ August - $15^{th}$ December) - assuming that each of them is characterized by specific short-term adjustments. Hence, we have re-assessed the short-run eq.(2) for the periods *before* and *after the financial turmoil*. Notice the former sub-sample period consists of 153 observations, while the second one of 83 observations. The results of eq.(2) referring to the full sample, the period *before the financial turmoil, after the financial turmoil* are presented in Table 4, columns (2), (3) and (4) respectively.[14]

TABLE 4. SHORT-TERM RESULTS.

DOLS Short-run equations

| | Full sample | Before the financial turmoil | After the financial turmoil |
|---|---|---|---|
| $\alpha_0$ | -0.0005 (0.0013) | -0.0012 (0.0016) | -0.0006 (0.0025) |
| $r_{brent,t}$ | 0.2380*** (0.0437) | 0.3125*** (0.0733) | 0.1821*** (0.0637) |
| $r_{switch,t}$ | 0.0495*** (0.0120) | 0.0536*** (0.01349 | 0.0418 (0.0265) |
| $r_{eurostoxx,t}$ | 0.2160*** (0.0586) | 0.1436 (0.1011) | 0.2882*** (0.0824) |
| $v_{t-1}$ | -0.0841*** (0.0272) | -0.0888*** (0.0365) | -0.0824*** (0.0450) |
| Main statistics | | | |
| $R^2$ | 0.3281 | 0.2302 | 0.4417 |
| Adj. $R^2$ | 0.3165 | 0.2090 | 0.4131 |
| F-stat[a] | 0.0000 | 0.0000 | 0.0000 |
| Durbin Watson | 1.7522 | 1.8510 | 1.6293 |
| Akaike Info. Criteria | -4.9510 | -5.0095 | -4.7572 |
| Schwarz Criterion | -4.8778 | -4.9091 | -4.6114 |
| Breush-Godfrey SCLM | 0.2789 | 0.3426 | 0.1318 |
| ARCH | 0.6857 | 0.1126 | 0.6098 |

[a]MacKinnon p-values: no-trend (240 obs.), 5% = -4.22; 10% = -3.85 (MacKinnon, 1991).

***, **, * indicate significance at 1%, 5% and 10% level, respectively.

**Full sample** The adjusted $R^2$ is reasonably high (32.8%) for a first difference model, and, as judged by the $F$-test $p$-value, the joint significance of results is accepted at the 1% level. All the variables in the short-run equation (Table 4, column 2) are significant at the 1% level, except the constant. The ARCH and the Breush-Godfrey Serial Correlation Lagrange Multiplier tests[15] show

[14]Robustness checks for the estimates were also performed by using ARDL techniques. Overall, the results are similar (see the Appendix for details).

[15]Breusch-Godfrey test: it is a test for autocorrelation in the residuals from any kind of statistical regression. It is the generalization of the Durbin-Watson test which is applicable only for the first order autoregressive model AR(1). It is formulated by regressing the sample residuals of the OLS analysis applied on the initial stochastic process on their lagged values (together with the initial regressors). The resulting $R^2$ statistic is then used asymptotically to

that residuals are homoskedastic and not serially correlated. The Akaike Information Criterion and the Schwarz Criterion gave us the optimal number of lags. Most importantly, residuals are negative and significant which indicates that carbon prices reach their long-term equilibrium with an 8.4% speed of adjustment. For instance, by perturbing the EUA with a shock equal to its standard deviation (0.049), the endogenous variable goes back to its long-run trend in approximately three months (70 observations).

By comparing short-term results with long-run elasticities, we see that the coefficients of *brent* and *switch* decrease while that of *eurostoxx* becomes significant. The latter effect indicates that, although long-term trends are governed by market fundamentals, there should be some rationale, in the short-run, for looking at carbon permits as an asset class. Undoubtedly, this unexpected result deserves further analysis, which is an issue analysed by dividing the sample in two sub-periods.

**Before the financial turmoil ($2^{nd}$ January - $11^{th}$ August).** The joint significance of results is accepted at the 1% significance level ($F$-test $p$-value). The adjusted $R^2$ is 20.90%, slightly below their full-sample counterpart (Table 4, column 3). Similar results apply to the Akaike Information Criterion and the Schwartz Criterion, thus indicating an acceptable goodness of fit. Compared to full-period - short-run - estimates and excluding *eurostoxx*, regressors remain statistically significant at the 5% level but their explanatory power changes: the coefficient of *brent* neatly increases, that of *eurostoxx* decreases. Consistently with the literature on the subject, *before the financial turmoil*, carbon allowances echo energy inputs more than financial assets.

**After the financial turmoil ($18^{th}$ August - $15^{th}$ December).** The joint significance of results is accepted at the 1% significance level ($F$-test $p$-value). The adjusted $R^2$ (41.32%, Table 4, column 4) is higher than their full-sample equivalent. *Brent* and *eurostoxx* are statistically significant at the 1% level; confirming the absence of abatement (fuel-switching) incentives *switch* loses explanatory power. Compared to full-period - short-run - estimates, the coefficient of *brent* neatly decreases while that of *eurostoxx* increases. *After the financial turmoil*, carbon allowances lose part of their input value and become a financial asset. Please note that the error correction term is statistically significant at the 10% level, and that the Akaike Information Criterion and the Schwartz Criterion are lower than those in the full-sample assessment. These findings are mainly due to the limited number of observations in the second sub-sample (83 observations).

We would like to remark that results are robust to the estimation technique used (see the Appendix). The speed of adjustment increases in the first sub-sample period and decreases in the second one, but with tiny differences (Table 4, columns 3-4). As confirmed by the ARCH and the Breusch-Godfrey Serial Correlation LM tests, residuals are not serially correlated.

test the correlation between the sample residuals and all their lagged terms. The p-values are associated to the Breusch-Godfrey test for the short-run equation, and one lag is close to the 5% bound. If more lags are considered, the rejection of the autocorrelation hypothesis is less problematic.

There are several possible explanations to econometric findings. It is self-evident that *the recent trend of carbon prices reflects the rapid deterioration of the economic situation*, which is massively reducing the demand for allowances. As indicated by Alberola *et al.* (2009b), both the energy sector and industries such as steel, cement and glass have a non-negligible impact on $CO_2$ prices. In 2008 two trends have emerged. On the one side, production cutbacks in steel, cement and glass industries were more pronounced than in other sectors thus resulting in a non-negligible cut in the demand - and hence in the price - for permits (Tendances Carbone, 2009). On the other, the stability of energy demand (at least at the European level) and the absence of fuel-switching incentives, which has led the use of the most carbon intensive technologies by power producers, have acted the other way round, especially by mid-2008. We believe that the former effect has prevailed, thus causing a downward pressure on carbon prices.

A further - less evident but equally important - factor is that *the economic crisis is changing the behaviour of carbon market players*. The sudden decline of demand (and hence of profits) has led several companies to short-term cash unbalances. Due to the financial turmoil, credit markets have become illiquid. By mid-2008, both capital markets and banks had tightened up corporate lending. In the same period, carbon constraints relaxed, as a consequence of production cutbacks. Therefore companies have started selling their unused carbon stocks to cover short-run cash unbalances. This monetization of permits is the main reason for the high volumes of trade at the end of 2008.

Each of the above mentioned mechanisms is consistent with the drop in EUA prices, as they describe a decrease in permit demand and an increase in supply. However, they do not explain the change in the short-run - adjustment - equation. Consistently with Alberola *et al.* (2008), which investigated the issue for the Pilot Phase of the EU-ETS, the likely *structural break* reflects the *impact of regulatory uncertainty* on $CO_2$ trading. As for Phase II, the allocation of the cap has been a very lengthy process. At the end of 2008, while the allocation of 275 Mton was still pending, the European Parliament approved upon first reading the draft Directive governing the EU-ETS for the period 2013 to 2020. In the absence of certainty concerning future rules, there is an increasing risk of wait-and-see attitudes. Sectors covered by EU-ETS, waiting for the environmental objective to become clear, have traded mostly for speculative purposes, as banks or other intermediaries also do, making EUA more similar to financial assets. This would explain why *eurostoxx* is significant in the short-run equation exclusively and in particular in the second sub-sample, when the technological variable (i.e. *switch*) becomes uninfluential and *brent* loses explanatory power.

However this strategy has been played at the wrong period - that is when the financial market has collapsed - therefore deeply affecting also $CO_2$ markets.

# 7 Conclusions

The core objective of this paper was to stress carbon price drivers at the beginning of the second phase of the EU-ETS. We have performed a cointegration analysis with ECM techniques, suitable to disentangle long- and short-term adjustments. The results point out that the EU carbon market has reached maturity, with energy fundamentals being the main long-term drivers. By confirming its pivotal role in the energy framework, *brent* is the key determinant in the long-run. The short-term speed of adjustment toward this long-term path is estimated at 8% in our sample. The short-term relationship also unveils that as from August 2008, $CO_2$ fundamentals have slightly changed, leaving room for financial price drivers.

Although the effectiveness of marketable allowances as an environmental tool was not an issue here, our work discloses several interesting aspects. First, carbon caps provide a useful economic shock absorber. Allowance prices automatically soften as soon as the economy enters a recession. In principle, carbon taxes could be adjusted to become countercyclical too. However, the efficient use of tax systems would require an unconventional level of sophistication and independence on the part of policy makers. Second, regulatory uncertainty - *i.e.* the repartition of carbon constraints among the sectors covered by the scheme - has affected the operation of $CO_2$ markets, by influencing its main price drivers. As Montero (1997) argues, in a marketable permit system, uncertainty on trade approval will lead to an outcome that does not reach the least-cost equilibrium in terms of environmental protection. Indeed, it is likely that the drop in carbon prices, together with the shift in the behaviour of market players, will delay investment decisions at the expense of the environment. This is a risky attitude, as the energy-climate package maintains a tight cap for the 2013-2020 period. As industries will be able to bank unused allowances to Phase III, we will probably see an increase in $CO_2$ prices by the end of 2012, when permit market shortages could be very costly to cover.

# References

[1] Abeysinghe, T. and Boon, T.K., 1999, Small Sample Estimation of Cointegratin Vector: an Empirical Evaluation of Six Estimation Techniques, Applied economics letters 6, 645-648.

[2] Alberola, E., Chevallier, J., Cheze, B., 2009a, European Carbon Price Fundamentals in 2005-2007: the Effects of Energy Markets, Temperatures and Sectorial Production, forthcoming in Journal of Policy Modeling.

[3] Alberola, E., Chevallier, J., Cheze, B., 2009b, The EU Emission Trading Scheme: Disentangling the Effects of Industrial Production and $CO_2$ Emission on Carbon Prices, forthcoming in Journal of International Economics.

[4] Alberola, E., Chevallier, J., Cheze, B., 2008, Price Drivers and Structural Breaks in European Carbon Prices 2005-2007, Energy Policy 36 (2), 787-797.

[5] Bahn, O., Barreto, L., Kypreos, S., 2001. Modelling and Assessing Inter-Regional Trade of $CO_2$ Emission Reduction Units. Environmental Modeling and Assessment 6, 173–182.

[6] Bahn, O., Kypreos, S., Büeler. B., Lüthi, H-J., 1999, Modeling an International Market of $CO_2$ Emission Permits, International Journal of Global Energy Issues 12(1–6), 283–291.

[7] Benz, E. and Trück, S., 2009, Modeling the Price Dynamics of $CO_2$ Emission Allowances, Energy Economics 31, 4–15.

[8] Benz, E. and Trück, S., 2006, $CO_2$ Emission Allowances Trading in Europe — Specifying a New Class of Assets. Problems and Perspectives in Management 3(3).

[9] Bonacina M. and Cozialpi, S., 2009, Carbon Allowances as Inputs or Financial Assets: Lesson Learned from the Pilot Phase of the EU-ETS, IEFE Working Paper n. 19

[10] Bunn, D. and Fezzi, C., 2007, Interaction of European Carbon Trading and Tnergy Prices, FEEM Working Paper 123, Milan.

[11] Bunn, D. and Fezzi, C., 2009, A Vector Error Correction Model of Interaction among Gas, Electricity and Carbon Prices: an application to the cases of Germany and the United Kingdom, Chapter 6 in Markets for Carbon and Power Pricing in Europe, edited by F. Gulli, Edward Elgar Publishing, London, UK.

[12] Burden Sharing Agreement, 1998, Commission Communication to the Parliament and the Council, Preparing for Implementation of the Kyoto Protocol, COM(1999) 230.

[13] Burniaux, J.-M., 2000, A Multi-Gas Assessment of the Kyoto Protocol, OECD Economics Department, Paris.

[14] Christiansen, A., Arvanitakis, A., Tangen, K., Hasselknippe, H., 2005, Price Determinants in the EU Emissions Trading Scheme, Climate Policy 5, 15–30.

[15] Cheung, Y.W. and Lai, K., 1993, Finite-sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration, Oxford Bulletin of Economics and Statistics 55, 313-328.

[16] Chevallier, J., *forthcoming*, Carbon Futures and Macroeconomic Risk Factors: A View from the EU-ETS, accepted manuscript, Energy Economics

[17] Ciorba, U., Lanza, A., Pauli, F., 2001, Kyoto Protocol an Emission Trading: Does the US make a Difference?, FEEM Working Paper 90, Milan.

[18] Convery, F. 2009, Reflections - The Emerging Literature on Emission Trading in Europe, Review of Environmental Economics and Policy, 3(1), 121-137.

[19] Convery, F., Ellerman, D., De Perthuis, C., 2008. The European Carbon Market in Action: Lessons from the First Trading period, Interim Report. Report 162, Cambridge, Massachusetts Institute of Technology, Center Energy and Environmental Policy Resources.

[20] Convery, F.J. and Redmond, L., 2007, Market and Price Developments in the European Union Emissions Trading Scheme, Review of Environmental Economics and Policy 1(1), 88-111.

[21] Daskalakis, G., Psychoyios, D., Markellos, R. N., 2007, Modeling $CO_2$ Emission Allowance Prices and Derivatives: Evidence from the European Trading Scheme, Working Paper, Athens University of Economics and Business.

[22] Demailly, D. and Quirion, P., 2008, Changing the Allocation Rules in the EU ETS: Impact on Competitiveness and Economic Efficiency, FEEM Working Paper 89, Milan.

[23] Diakoulaki, D. and Mandaraka, M., 2007, Decomposition Analysis for Assessing the Progress in Decoupling Industrial Growth from $CO_2$ Emission in the EU Manufacturing Sector, Energy Economics 29, 636–664.

[24] Directive 2003/87/EC of the European Parliament and of the Council of 13 October 2003 establishing a scheme for greenhouse gas emission allowance trading within the Community and amending Council Directive 96/61/EC.

[25] Directive 2008/101/EC of the European Parliament and of the Council of 19 November 2008 amending Directive 2003/87/EC so as to include aviation activities in the scheme for greenhouse gas emission allowance trading within the Community.

[26] Ellerman, A. and Bucher, B., 2007, The European Union Emission Trading Scheme: Origins, Allocation and Early Results, Review of Environmental Economics and Policy, 1, 66-87.

[27] Ellerman, A. and Joskow, P., 2008, The European Emission Trading System in Perspective, PEW Center on Global Climate Change.

[28] Engle, R.F. and Granger, C.W.J., 1987, Co-integration and error-correction: Representation, Estimation and Testing, Econometrica 55, 251-276.

[29] Grubb, M., Edmonds, J., Ten Brink, P., Morrison, M., 1993, The Costs of Limiting Fossil-Fuel $CO_2$ Emissions, Annual Review of Energy and the Environment 18, 397–478.

[30] Juselius, K., 2006, The Cointegrated VAR model: Methodology and Applications. Oxford University Press, Oxford.

[31] Johansen, S., 1988, Statistical Analysis of Cointegration Vectors, Journal of Economic Dynamics and Control 12, 231-254.

[32] Johansen, S. and Juselius, K., 1990, Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money, Oxford Bulletin of Economics and Statistics 52, 169-209.

[33] Kainuma, M., Matsuoka, Y., Morita, T., 1999, Analysis of Post-Kyoto scenarios: the Asian-Pacific integrated model. In: Weyant, J. (Ed.), The costs of the Kyoto Protocol: A Multi-Model Evaluation, Energy Journal, Special Issue, 207-220.

[34] Kanen, J.L.M., 2006, Carbon Trading and Pricing, Environmental Finance Publications, London.

[35] Kanudia, A. and Loulou, R., 1998, Joint Mitigation under the Kyoto Protocol: A Canada-USA-India Case Study, GERAD discussion paper G, 98-40, Montreal.

[36] Kremers, J., Ericsson, N., Dolado, J., 1993, The Power of Cointegration Tests. Oxford Bulletin of Economics and Statistics 54, 349-367.

[37] Liaskas, K., Mavrotas, G., Mandaraka, M., Diakoulaki, D., 2000, Decomposition of Industrial $CO_2$ Emissions: the case of European Union, Energy Economics 22, 383–394.

[38] MacKinnon, J.G., 1991, Critical Values for Co-integration Tests, in R.F. Engel and C.W.J. Granger (eds.), Long-run econometric relationships, Oxford, Oxford University Press.

[39] Mansanet-Bataller, M. and Pardo, A., 2008, $CO_2$ Prices and Portfolio Management, Working Paper, Department of Financial Economics, University of Valencia.

[40] Mansanet-Bataller, M., Pardo, A., Valor, E., 2007, $CO_2$ Prices, Energy and Weather, The Energy Journal 28(3), 67–86.

[41] Masih, R. and Masih, A.M.M., 1996b, Stock-Watson Dynamic OLS (DOLS) and Error-Correction Modelling Approaches to Estimating Long and Short-Run Elasticities in a Demand Function: New Evidence and Methodological Implications from an Application to the Demand for Coal in Mainland China, Energy Economics 18, 315-334.

[42] Masih, A.M.M. and Masih, R., 1995a, Temporal Causality and the Dynamic Interactions among Macroeconomic Activity within a Multivariate Cointegrated system: Evidence from Singapore and Korea, Weltwirtschaftliches Archiv 131.

[43] Masih, A.M.M. and Masih, R., 1995b, Investigating the Robustness of Rests of the Market Efficiency Hypothesis: Contributions from Cointegration Techniques on the Canadian Floating Dollar, Applied Financial Economics 5, 139-150.

[44] McKibbin, W.J., Shackleton, R., Wilcoxen, P.J., 1999, What to Expect from an International System of Tradable Permits for Carbon Emissions, Resource and Energy Economics 21, 319-346.

[45] Milunovich, G. and Joyeux, R., 2007, Testing Market Efficiency and Price Discovery in European Carbon Markets. Research Papers 0701, Macquarie University, Department of Economics.

[46] Montero, J.-P, 1997, Marketable Pollution Permits with Uncertainty and Transaction Costs, Resource and Energy Economics 20, 27-50.

[47] New Carbon Finance, 2009, Emission from EU ETS down 3% in 2008, Research Notes February 13th.

[48] Nordhaus, W.D., 2001, Global Warming Economics, Science 294, 1283-1284.

[49] Nordhaus, W.D. and Boyer, J.G., 1999, Requiem for Kyoto: an Economic Analysis of the Kyoto Protocol. In: Weyant, J. (Ed.). The Costs of the Kyoto Protocol: A Multi-Model Evaluation, The Energy Journal, Special Issue, 93-130.

[50] Paolella, M.S. and Taschini, L., 2007, An Econometric Analysis of Emission-Allowance Prices, Journal of Banking and Finance 32(10), 2022-2032.

[51] Pesaran, M.H., Shin, Y., Smith, R.J., 2001, Bounds Testing Approaches to the Analysis of Level Relationships, Journal of Applied Econometrics 16, 289–326.

[52] Pesaran, M H. and Smith, R.J., 1998, Structural Analysis of Cointegrating VARs, Journal of Economic Surveys 12(5), 471-505.

[53] Pesaran, M.H. and Pesaran, B., 1997, Working with Microfit 4.0: An Interactive Econometric Software Package (DOS and Windows versions), Oxford University Press, Oxford.

[54] Phillips, P.C.B. and Loretan, M., 1991, Estimating Long-Run Equilibria, The Review of Economic Studies 58(3), 407-436.

[55] Point Carbon, 2009, All you need to know about EU ETS Phase 2, Research Report, January.

[56] Saikkonen, P., 1991. Asymptotically Efficient Estimation of Cointegration Regressions, Econometric Theory 7, 1-21.

[57] Sanstad, A.H. and Greening, L.A., 1998, Economic Models for Climate Policy Analysis: a Critical Discussion, Environmental Modeling and Assessment 3, 3-18.

[58] Schwarz, G., 1978, Estimating the Dimension of a Model, Annals of Statistics 6, 461-464.

[59] Seifert, J., Uhrig-Homburg, M., Wagner, M.W., 2008, Dynamic Behavior of $CO_2$ Spot Prices, Journal of Environmental Economics and Management 56, 180-194.

[60] Sijm, J.P.M., Ormel, F.T., Martens, J.W., Van Rooijen, S.N.M., Voogt, M.H., 2000, Kyoto Mechanisms. The Role of Joint Implementation, the Clean Development Mechanism and Emissions Trading in Reducing Greenhouse Gas Emissions, ECN report C-00-026, Petten, The Netherlands.

[61] Société Générale, 2009, Carbon Drivers, Orbeo Research Report, 9th March.

[62] Springer, U. and Varilek, M., 2004, Estimating the Price of Tradable Permits for Greenhouse Gas Emissions in 2008-2012, Energy Policy 32, 611-621.

[63] Springer, U., 2003, The Market for Tradable GHG Permits under the Kyoto Protocol: a Survey of Model Studies, Energy Economics 25(5), 527-551.

[64] Stock, J.H. and Watson, M., 1993, A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, Econometrica 61, 783–820.

[65] Tendances Carbone, 2007, Methodological Notes.

[66] Tendances Carbone, 2009, Monthly Bulletin n. 32.

[67] Thomas, R.L., 1993, Introductory Econmetrics: Theory and Applications. London, Longman.

[68] Trotignon, R. and Delbosc, A., 2008, Allowance Trading Patterns During the EU-ETS Trial Period: What does the CITL reveal? Research Report 13, Mission Climat Caisse des Dépôts, Paris.

[69] Van der Mensbrugghe, D., 1998. A (Preliminary) Analysis of the Kyoto Protocol: Using the OECD GREEN Model, Economic Modelling of Climate Change, OECD, Paris.

[70] Zhang, Z.X. and Folmer, H., 1998, Economic Modelling Approaches to Cost Estimates for the Control of Carbon Dioxide Emissions, Energy Economics 20, 101–120.

# 8 Appendix

TABLE I. TRADING INDICATORS: 2008 versus 2006

| | Mean | Min | Max | Std.Dev. |
|---|---|---|---|---|
| EUA 2008 (1y) (€/t$CO_2$) | 23.41 | 14.36 | 30.53 | 3.45 |
| EUA 2008 (2y) (€/t$CO_2$) | 24.14 | 14.97 | 31.71 | 3.56 |
| EUA 2006 (1y) (€/t$CO_2$) | 18.26 | 6.60 | 31.50 | 7.27 |
| EUA 2006 (2y) (€/t$CO_2$) | 20.47 | 15.05 | 32.25 | 4.00 |
| Volumes 2008 (Mton daily) | 5.61 | 1.82 | 17.67 | 2.34 |
| Volumes 2006 (Mton daily) | 1.80 | 0.17 | 7.48 | 1.80 |

SOURCE: European Climate Exchange, 2009

**Correlation matrixes and descriptive statistics.**

TABLE II. CORRELATION MATRIX (variables in returns). FULL SAMPLE.

| | EUA | Switch | Eurostoxx | Brent |
|---|---|---|---|---|
| EUA | 1 | | | |
| Switch | 0.274 | 1 | | |
| Eurostoxx | 0.396 | 0.081 | 1 | |
| Brent | 0.455 | 0.090 | 0.465 | 1 |

TABLE IIIa. DESCRIPTIVE STATISTICS (EUA, €/t$CO_2$)

| | Full sample | Before the financial turmoil | After the financial turmoil |
|---|---|---|---|
| Mean | 22.67 | 23.85 | 20.45 |
| Std.Dev. | 3.36 | 2.45 | 3.67 |
| Min | 13.72 | 18.84 | 13.72 |
| Max | 29.33 | 29.33 | 25.19 |

TABLE IIIb. CORRELATION MATRIX (variables in returns).

| Before the financial turmoil | | | | | After the financial turmoil | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | EUA | Switch | Eurostoxx | Brent | | EUA | Switch | Eurostoxx | Brent |
| EUA | 1 | | | | EUA | 1 | | | |
| Switch | 0.314 | 1 | | | Switch | 0.224 | 1 | | |
| Eurostoxx | 0.157 | 0.079 | 1 | | Eurostoxx | 0.571 | 0.100 | 1 | |
| Brent | 0.304 | 0.023 | 0.059 | 1 | Brent | 0.566 | 0.171 | 0.614 | 1 |

**Autoregressive Distributed Lags (ARDL).** This Section proves the robustness of the results in the paper by carrying out an ARDL model (0,0,1,1).[16] The method has several advantages: it

[16] For an extensive assessment of ARDL models see Pesaran *et al.* (1997, 1998 and 2001).

produces consistent estimates with small samples; it can be applied to both I(0) and I(1) variables; it can replicate an ECM by a simple linear transformation. Econometric theory and empirical applications have demonstrated that ARDL and ECM techniques should provide almost the same results. First of all we apply the "bounds testing procedure". The test consists in estimating a regression where the $r_{t,EUA}$ is the dependent variable, and verifying whether the coefficients of *EUA*, *brent*, *eurostoxx* are jointly different from 0. The hypothesis of no cointegration is rejected at the 10% level (F-test).

TABLE IVa. BOUNDS TESTING PROCEDURE

| | | | |
|---|---|---|---|
| $\alpha_0$ | -0.0860 (0.1188) | $r_{eua,t-1}$ | 0.1323* (0.0786) |
| $EUA_{t-1}$ | -0.1209*** (0.0315) | $r_{switch,t-1}$ | 0.0020 (0.0149) |
| $brent_{t-1}$ | 0.0444*** (0.0168) | $r_{eurostoxx,t-1}$ | -0.0862 (0.0721) |
| $switch_{t-1}$ | 0.0085** (0.0043) | $r_{brent,t-1}$ | -0.0061 (0.0560) |
| $eurostoxx_{t-1}$ | 0.0286 (0.0191) | | |
| Main statistics | | | |
| $R^2$ | 0.0808 | Akaike Info. Criteria | -4.6084 |
| Adj. $R^2$ | 0.0487 | Schwarz Criterion | -4.4771 |
| F stat (p-value) | 0.0122 | Durbin Watson | 2.0006 |

***, **, * indicate significance at 1%, 5% and 10% level, respectively.

The F test on no co-integration is rejected in almost all the cases at least at 10% level, as the following table shows (Table IVb).

TABLE IVb.

| | |
|---|---|
| F statistics | 3.8422 (0.0048) |
| Chi-square | 15.3687 (0.0040) |

Results for the long-run equation are in Table V.

TABLE V. ARDL FULL SAMPLE RESULTS: long-run.

ARDL long-run equation

| | | | |
|---|---|---|---|
| $\alpha_0$ | 0.3853 (0.2498) | $r_{brent,t-1}$ | 0.1035 (0.1119) |
| $brent_{t-1}$ | 0.2843*** (0.1122) | $r_{switch,t-1}$ | 0.0102 (0.0304) |
| $switch_{t-1}$ | 0.0552** (0.0304) | $r_{eurostoxx,t-1}$ | 0.2200 (0.1478) |
| $eurostoxx_{t-1}$ | -0.1274 (0.1493) | $r_{eua,t-1}$ | 0.1035 (0.1119) |
| $EUA_{t-1}$ | 0.3853 (0.2498) | | |

Main statistics

| | | | |
|---|---|---|---|
| $R^2$ | 0.9035 | Durbin Watson | 0.2092 |
| Adj. $R^2$ | 0.9010 | Akaike Info. Criteria | -3.1060 |
| F-stat[a] | 0.0000 | Schwarz Criterion | -3.0041 |
| Wald test (4 variables) | 3.84 | | |

[a]MacKinnon p-values: no-trend (240 obs.), 5% = -4.22; 10% = -3.85 (MacKinnon, 1991).

***, **, * indicate significance at 1%, 5% and 10% level, respectively.

Compared to ECM, ARDL long-term estimates are characterized by a lower coefficient for *brent* (0.2843 versus 0.3897) and a lower statistical significance of *switch* (10% versus 5% level). However, results for short periods (*full sample*, *before the financial turmoil* and *after the financial* *turmoil*) are similar to those in Table 4 (see Table VI).

TABLE VI. ARDL SHORT-TERM RESULTS.

ARDL Short-run equations

| | Full sample | Before the financial turmoil | After the financial turmoil |
|---|---|---|---|
| $\alpha_0$ | -0.0004 (0.0013) | -0.0013 (0.0016) | -0.0003 (0.0026) |
| $r_{brent,t}$ | 0.2457*** (0.0437) | 0.3128*** (0.0729) | 0.2055*** (0.0668) |
| $r_{switch,t}$ | 0.0490*** (0.0120) | 0.0537*** (0.0134) | 0.0391 (0.0271) |
| $r_{eurostoxx,t}$ | 0.1938*** (0.0593) | 0.1409 (0.1047) | 0.2490*** (0.0846) |
| $r_{brent,t-1}$ | 0.0753* (0.0440) | 0.1044 (0.0745) | 0.0929 (0.0661) |
| $r_{switch,t-1}$ | 0.0129 (0.0120) | 0.0173 (0.0133) | 0.0008 (0.0274) |
| $r_{eurostoxx,t-1}$ | -0.0053 (0.0583) | 0.0899 (0.1012) | -0.0551 (0.0835) |
| $v_{t-1}$ | -0.0817*** (0.0264) | -0.0874*** (0.0264) | -0.0774* (0.0444) |
| Main statistics | | | |
| $R^2$ | 0.3431 | 0.2554 | 0.4587 |
| Adj. $R^2$ | 0.3231 | 0.2189 | 0.4082 |
| F-stat[a] | 0.0000 | 0.0000 | 0.0000 |
| Durbin Watson | 1.7480 | 1.8225 | 1.6635 |
| Akaike Info. Criteria | -4.9527 | -5.0101 | -4.7158 |
| Schwarz Criterion | -4.8360 | -4.8503 | -4.4827 |
| Breush-Godfrey SCLM | 0.2598 | 0.3057 | 0.1719 |
| ARCH | 0.7005 | 0.2231 | 0.7588 |

[a]MacKinnon p-values: no-trend (240 obs.), 5% = -4.22; 10% = -3.85 (MacKinnon, 1991).

***, **, * indicate significance at 1%, 5% and 10% level, respectively.