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DISCUSSION PAPER

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Short-Term Impacts of Carbon Offsetting on Emissions Trading Schemes: Empirical Insights from the EU Experience

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Short-term Impacts of Carbon Offsetting on Emissions Trading Schemes: Empirical Insights from the EU Experience

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Abstract

The Paris Agreement established a new mechanism by which a country can offset some of its emissions reductions in other countries. Its design is still under negotiation. While taking advantage of cheaper abatement opportunities enables efficiency gains, the impact on the price volatility in the emission trading schemes is unclear. We conduct an empirical analysis of the short-term impacts of these credits on the standard carbon markets, using the European Union experience with accepting credits for compliance in the second phase of its scheme. With vector-autoregressive models allowing regime changes at a priori unknown dates, we analyze the structural relationship between the prices of allowances and credits. Although one might expect that the allowance and credit markets influence one another, we find that, before November 2011, knowing the credit price variations helps to better predict the allowance price variations while, after November 2011, it is the opposite. We explain this by expectations and restrictions regarding credits. For the transmission of shocks and the impact on volatility, the influence is mainly from allowances to credits. The allowance price volatility explains between 56% and 72% of the credit volatility whereas the latter explains less than 2% of the former.

Keywords: Emissions trading; European allowances; international credits; causality analysis.

JEL classification: C32, C58, F18, Q54, Q58.

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1 Introduction

In the Paris Agreement reached within the United Nations Framework Convention on climate change (UNFCCC), the Parties agreed upon principles for voluntary market mechanisms. This includes Article 6.4, which establishes a mechanism that makes it possible for activities contributing to emissions reductions in a host Party to be used by another Party to fulfill its nationally determined contributions (NDCs). The aim of such an offsetting scheme is to support mitigation efforts and foster sustainable development. As we are writing this paper, the exact design of this mechanism is still being negotiated.[1] In particular, the transition from the current international carbon credit system as well as accounting rules are still debated.

Developping international cooperation on emissions trading offers possibilities of efficiency gains by taking advantage of cheaper abatement opportunities. Reducing the cost of abatement in the jurisdictions where abatement is more expensive and avoiding distortions and leakages induced by heterogenous carbon pricing are the main motivations for supporting the development of a worldwide carbon market (Tirole, 2012) or internationally uniform carbon pricing (Gollier and Tirole, 2015). A number of factors may, however, impact the efficiency gains that are expected from linking carbon markets. In the case of bilateral and sequential trading of emissions, Burtraw *et al.* (1998) show that the sequence of trade may undermine the cost savings. The efficiency gains may also be affected by the initial allocation of allowances (Lange, 2012).[2,3] In the case of a project-based supply of international credits, the potential frictions in the permit supply may reduce the cost savings (Liski and Virrankoski, 2004). In the case of a sector-based supply of international credits, the carbon price equalization may actually result in a welfare loss for the developing country involved (Hamdi-Cherif *et al.*, 2011; Gavard *et al.*, 2011). A limit on the amount of permits traded would mitigate this effect (Gavard *et al.*, 2016).[4]

Besides these efficiency considerations, the impact of such international mechanisms on the volatility of the permit prices is an important aspect to consider as it may impact the companies covered by emission trading schemes (ETS). Much of the existing literature consists in theoretical and numerical

[1] No agreement on Article 6 could be finalized at the 24^{th} and 25^{th} Conferences of the Parties (COP) (respectively hold in Katowice in 2018 and in Madrid in 2019). The negotiations will continue in the following UNFCCC sessions. Due to the Covid pandemic, the next Conference of the Parties where a draft decision could be considered will not take place before the end of 2021.

[2] In particular, in the presence of frictions in the emissions market, this initial allocation may impact the market size and trading costs (Liski, 2001).

[3] Habla and Winkler (2018) theoretically show that strategic delegation incentives may also reduce the attractiveness of linkages.

[4] Besides volume restrictions, temporary restrictions are suggested as intermediary steps towards full linkage (Quemin and de Perthuis, 2018).

studies and has focused on linkages between standard ETSs. It suggests that the effect of carbon market coupling is uncertain.[5] A trade-off between efficiency gains and uncertainty increase might be required. There is a need for more empirical analyses of the impact of offsetting on the permit price return[6] and volatility to draw lessons for the design of new international carbon credit schemes.

We empirically examine the short-term impacts of international credits on the standard carbon markets, including with regard to the transmission of shocks and the effect on the permit volatilities. As the European Union Emissions Trading Scheme (EU ETS) is the largest carbon market to have accepted offsets, we take advantage of this experience to analyze the impacts to expect on the ETSs by focusing on the second phase of the EU ETS, which is the time period when credits were accepted in Europe before type restrictions were introduced.[7] Our analysis takes account of potential interactions between the two price series as well as potential regime changes in the relationship. Indeed, in the case of an emissions trading scheme, even if the supply of permits is fixed, set by a cap that is decided at a political level, the demand for allowances might be influenced by the acceptance of international carbon credits. For the latter, the supply is impacted by international energy prices as well as investment support and the demand is a function of the standard ETSs. In addition, policy announcements that took place regarding the acceptance of credits in the second phase of the ETS as well as external factors such as the 2008 economic crisis are also likely to have impacted the interactions between the two price series. For these reasons, our structural analysis of their relationship employs a vector autoregressive model (VAR) and we use the approach developed by Qu and Perron (2007) to allow for potential and a priori unknown regime changes and detect them. For each identified regime, we use the estimations from the VAR analysis to test the existence of a causality relationship in the sense of Granger, examine how shocks are transmitted between the prices of allowances and credits and analyze how their variances mutually influence one another. We discuss the variations across regimes.

While the functioning of the allowance and credit markets described above could imply a bi-directional relationship between their price returns, we actually find a uni-directional structural relationship which changes over time. Before November 2011, knowing the credit price variations enables to better predict

[5]Based on a theoretical and numerical approach, Doda *et al.* (2019) suggest that, while linkage should lower price volatility on average, this may not always be the case for individual jurisdictions. The uncertainty implied by the possibility for couplings to be terminated may induce costs and cause price divergence (Pizer and Yates, 2015). Together with the market sizes, the sunk costs of linking as well as the jurisdiction characteristics, such uncertainties impact the actual economic advantages of linking (Doda and Taschini, 2017).

[6]In this paper, "return" refers to the daily price variation.

[7]Before 2008, no international credits were used in the EU ETS due to the excess of EU allowances in Phase I of the scheme (Ellerman *et al.*, 2016). From 2013 onwards, the type of credits that could be used in the European system were strongly restricted.

the allowance return. After November 2011, it is the opposite. We explain this relationship change by the general economic activity as well as expectations and restrictions regarding international credits. For the transmission of shocks and the impact on the volatility, we find that the influence is in the same direction for both regimes: shocks in the allowance price are immediately transmitted to the credit price but shocks in the credit price are hardly transmitted to the allowance price. The allowance price volatility explains more than 56% of the credit price volatility while the latter explains less than 2% of the former.

Much of the existing empirical literature on the price interactions between emission trading schemes and international offsets has focused on the spread between the prices of European allowances (EUAs) and Certified Emission Reductions (CERs) generated under the Clean Development Mechanism (CDM). Nazifi (2013), Mizrach (2012), Mansanet-Bataller *et al.* (2011) and Chevallier (2010) indicate that the price of international credits has been largely influenced by the European carbon market, due to the fact that it has been the largest in the world to accept international credits for compliance (Ellerman *et al.*, 2010; Mansanet-Bataller *et al.*, 2011). These studies have focused on shorter time ranges than the one we consider. More recently, Hintermann and Gronwald (2019) developed a model to explain the spread between the EUA and CER price series. They indicate that the CER price formation is dominated by the uncertainty about offset demand and supply. The impact of these credits on the standard schemes still needs to be quantified (Trotignon, 2012). The literature on the impact of the carbon market volatility has mostly consisted in theoretical and numerical studies and rather focused on coupling between cap and trade systems. Our empirical analysis is important to better understand the impacts to expect from a more extended use of international carbon credits as a step towards more globally harmonized carbon pricing.

The following section presents the institutional background of this work. Section 3 describes the data we used. Section 4 presents the methodology used for the structural analysis and the detection of regime changes with Qu and Perron's approach. Section 5 discusses the results and Section 6 concludes.

2 Institutional background

According to the World Bank (WB, 2019), as of April 2019, 57 carbon pricing initiatives, including 28 emissions trading schemes, are in operation or planned in the world. These include, for example, carbon markets in China, Europe, Australia, California, Korea, Canada, New Zealand, Kazakhstan, and Ukraine. The European system started in 2005. It is the largest one after the more recent Chinese scheme. EUA are issued annually at the EU level. Their volume is defined by the European cap and, each year, installations covered by the European trading scheme have to surrender carbon allowances in a volume equivalent to the volume of their verified emissions that year.

Under the United Nations Framework Convention on Climate Change (UNFCCC), offsetting mechanisms have been put in place, initially in the framework of the Kyoto Protocol. With the Clean Development Mechanism, CER issued for approved projects in developing countries (Lecocq and Ambrosi, 2007) can be used by industrialized countries to meet their emission reduction target under the Kyoto Protocol. Under the Joint Implementation, Emission Reduction Units (ERU) from projects in Annex B countries[8] can be used by other Annex B countries to reach their emissions reduction objectives. These international credits can be traded worldwide and there is no limit on the amount of CERs issued annually.

At the Conference of the Parties in Paris in 2015, it was decided to establish a new mechanism by which countries can conduct emission reductions in a host Party and take them into account to fulfil their own nationally determined contributions. This so-called Article 6.4 mechanism could replace the current Clean Development Mechanism.

Historically, the European Union Emissions Trading Scheme (EU ETS) has been the largest carbon market to accept international credits for compliance. Companies covered by the scheme were given the possibility to surrender Kyoto Protocol credits together with the EUAs to cover their emissions. In the first phase of the scheme (2005-2007), given the lack of scarcity of EUAs, this option was not chosen by companies, but, in the second phase (2008-2012), companies used offsets up to the limit of 13% of the amount of EUA issued under the European cap. This experience provides a privileged framework to empirically analyze the impact of international credits on emissions trading schemes.

[8]Countries included in Annex B to the Kyoto Protocol for the first commitment period were Australia, Austria, Belgium, Bulgaria, Canada, Croatia, the Czech Republic, Denmark, Estonia, the European Union, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Latvia, Liechtenstein, Lithuania, Luxembourg, Monaco, the Netherlands, New Zealand, Norway, Poland, Portugal, Romania, Russian Federation, Slovakia, Slovenia, Spain, Sweden, Switzerland, Ukraine, the United Kingdom of Great Britain and Northern Ireland, and the United States of America. Canada withdrew from the Kyoto Protocol in December 2012 and the United States never ratified the Protocol.

3 Data

We focus our analysis on the EUA and CER prices in the second phase of the EU ETS. More specifically, our analysis covers the time period from 14 March 2008 to 12 November 2012. The reason is that this is the phase when offsets were accepted in the EU ETS without type restrictions. The data are taken from the Intercontinental Exchange (ICE). As the market for carbon permits is dominated by contracts for futures (the volume of spot contracts is relatively small), we generate the series by rolling over futures contracts. Given the importance of energy prices for carbon markets, we also use coal and natural gas price data from the ICE.[9] The conversion from $ to € for the coal price and from £ to € for the gas price requires using exchange rates. These are taken from the European Central Bank. Finally, to take into account the effect of the general economic activity on the carbon market, we use the Euro Stoxx 50 Index as a proxy, as done by Creti et al. (2012) and Bredin and Muckley (2011) for other analyses of the EU ETS.[10,11] The descriptive statistics of the daily variations of the logarithmic price series are reported in Appendix A (see Table 10). Several stationarity tests are conducted and reported in Appendix. They show that the series are not stationary but that the daily variations are.[12]

Figure 1 presents the EUA and CER futures price series while Figure 2 shows their daily variations. The carbon market was strongly impacted by the economic and financial crisis in 2008. This is visible in the drop in the EUA and CER price from September 2008 onwards as well as in the relatively high volatility in their return from September 2008 to July 2009. This is correlated with the downward trend and high volatility observed in coal and gas prices from September 2008 to July 2009 (see Figure 3). The carbon market was also affected by the recession in Europe in the third trimester of 2011. This is observable in the drop in the EUA and CER price series as well as in their volatility increase from July 2011 onwards.

In parallel, the CER market was influenced by policy announcements. As early as January 2009, the European Commission announced that there would be restrictions on the type of credits accepted for compliance in the European carbon market, but the list of credit types that would be recognized or not was only published in January 2011. This might have contributed to the observed downward trend in

[9] For coal and natural gas, we employ the price series of month-ahead contracts. For coal, we use the price series of API2 CIF (Cost, Insurance, Freight) with delivery in ARA (Amsterdam, Rotterdam and Antwerp).

[10] The use of this proxy has several justifications. It is difficult to find daily data on the industrial production or on the electricity production or consumption for the whole of the EU. National daily data available for electricity display seasonality patterns and do not well represent the economic activity.

[11] The Euro Stoxx 50 Index and its return are displayed in Figure 8 in Appendix A.

[12] A detailed discussion of these tests is provided in Gavard and Kirat (2018).

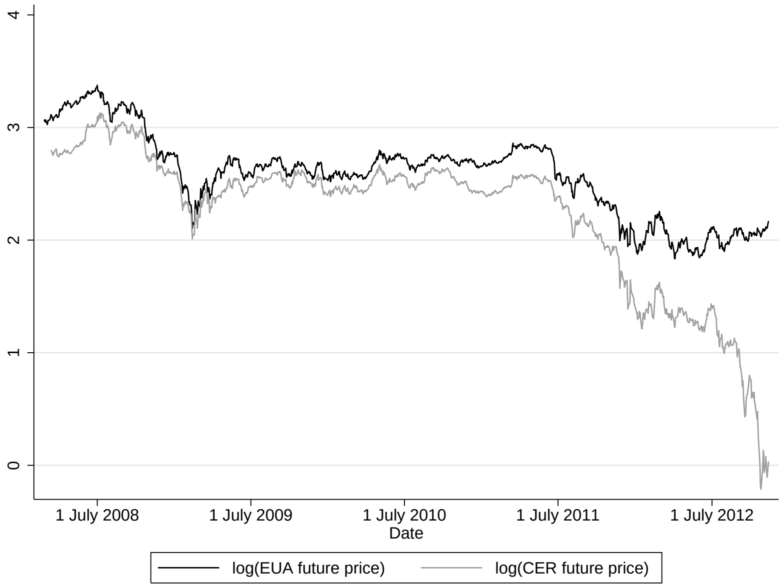
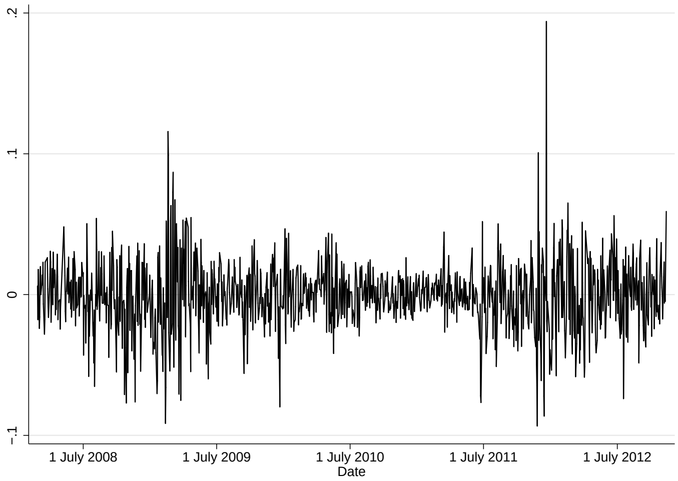


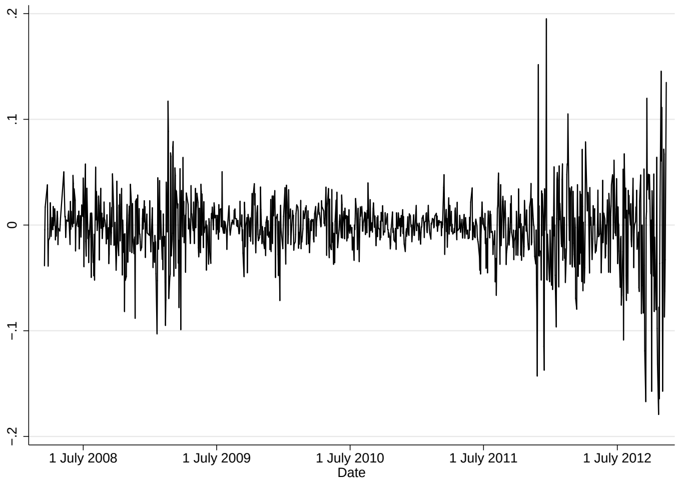
Figure 1: Futures price series of European Emission Allowances and Certified Emissions Reductions (in logarithm).

the CER price from July 2011 onwards[13] but two additional mechanisms are likely to also explain this drop in the offset price. On the one hand, the European institutions set a limit on the volume of credits that each installation covered by the scheme could use for compliance. As installations progressively reached their respective limits, the demand for these carbon permits gradually decreased. On the other hand, there was also a supply effect: the CDM was becoming more mature and, as underlying projects finally generated emission reductions, CER issuance increased sharply (oversupply due to the time lag between investment decisions and CER issuance). This increase in the supply of CER is confirmed by the figures of the volume of CERs issued annually under the Kyoto Protocol, as reported by Ellerman *et al.* (2016) (see Table 1). In the years from May 2008 to April 2009 and from May 2009 to April 2010, there were less than 150 million CERs issued annually (148 and 119 million respectively for each year). From 2010 onwards, as the gas price was increasing and recovering its pre-crisis level (see Figure 3a), we observe that the volume of CERs issued annually kept rising: it was close to 200 million for

[13]Visually, the downward trend in the CER price seems to start in July 2011. This date coincides with the launch of the Sandbag's report *Buckle Up! 2011 Environmental Outlook for the EU ETS*, on the occasion of which the Climate Action Commissioner reminded her audience of several reforms regarding the use of offsets in the EU-ETS in a speech she pronounced at the European Parliament (EC, 2011b). These reforms include the limitation of the use of offsets in the EU ETS from 2013 onwards, the focus on credits coming from projects in least developed countries, and the ban of controversial industrial gas projects. Her speech also indicated that the EU would push for a reform of this Kyoto Protocol offset mechanism. Statistically, Gavard and Kirat (2018) detect a break in the CER price trend only in November 2011.

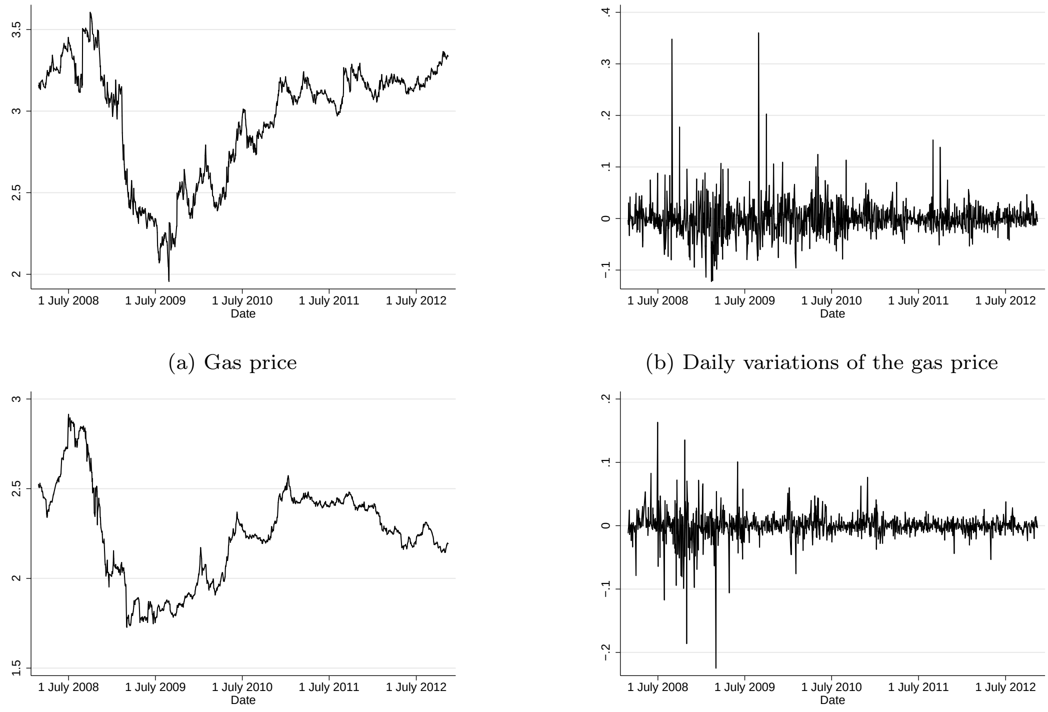


(a)



(b)

Figure 2: Daily variations of the (a) EUA and (b) CER price.



(a) Gas price

(b) Daily variations of the gas price

(c) Coal price

(d) Daily variations of the coal price

Figure 3: Energy prices (in logarithm) and corresponding returns (first differences of the logarithmic series)

May 2010 - April 2011, more than 300 million for May 2011 - April 2012 and nearly 400 million for May 2012 - April 2013. This increase in supply in addition to a reduction in demand is likely to have contributed to the drop in the CER price from July 2011 onwards. Finally, we should also note that, at the 17^{th} COP in Durban in November 2011, the decision was taken to review existing market-based mechanisms and to develop new ones to help emerging and developing countries in their emission reduction efforts (KPMG, 2011). Together with the announcements and mechanisms described above, this might have enhanced the uncertainty regarding offsets and hence their price volatility.

We test the existence of a long-term relationship between the EUA and CER price series. Several authors have found no evidence of cointegration between EUA and CER price series: Nazifi (2013) for the time period from March 2008 to May 2009, Mizrach (2012) for the time period from June 2007 to April 2010.[14] The observation of the EUA and CER prices over time (Figure 1) suggests the absence of cointegration. Even if EUA and CER prices have common drivers and might influence one

[14] Mansanet-Bataller *et al.* (2011) and Chevallier (2010) find some cointegration between EUA and CER prices, but Mizrah (2012) suggests that this is due to the fact that they use the Reuters index for the CER data and that this index averages prices from different expiries.

Table 1: Volume of CER issued under the Kyoto Protocol and surrendered in the EU ETS

Year	CER surrendered in the EU ETS (millions)	CER issued under the Kyoto Protocol (millions)
Pre-2008	0	139
2008	84	148
2009	78	119
2010	117	198
2011	178	314
2012	220	388
Total	667	1,308

Adapted from Ellerman *et al.* (2016) based on UNEP Risø Centre (2013) and EC (2015).
Note: ETS compliance years are from May of the year indicated to April of the following year. They nearly coincide with the years considered for CER issuance.

another, the Engle Granger test (see Schaffer, 2010), which takes account of breaks in the series, finds no evidence of a long-term relationship between the EUA and CER prices on the time period from March 2008 to November 2012. The results reported in Table 2 suggest that we cannot reject the null hypothesis of no cointegration.

The Johansen test results presented in Table 3[15] requires more discussion. The trace and maximum eigenvalue statistics do not allow to reject the null hypothesis of no cointegration when one uses a specification with trend in the cointegration equation and the error correction equation, but this null hypothesis can be rejected if one uses a specification with a constant and without trend for the cointegration relationship. However, in that case, the estimation of the corresponding error correction model indicates that the coefficient associated with the error correction term is not significant and cannot bring the long-term relationship back to equilibrium. We hence conclude that the EUA and CER price series are not cointegrated.[16] We explain this result by the fact that, even if EUA and CER prices are driven by similar factors, their long-term dynamics are different. The absence of evidence of cointegration is a justification for using the first differences of the permit prices in the analysis that follows.

[15] According to the Bayesian and Hannan-Quinn information criteria, we select one lag for inclusion in the error correction model. We present the Johansen test results for several specifications. The likelihood ratio tests select the model without constant as the most appropriate model.

[16] The graphical representation of the residuals of the regression of the CER price on the EUA price is reported in Figure 9 in Appendix A. Its random walk aspect corroborates this conclusion.

Table 2: Results of the Engle-Granger cointegration test.

Null hypothesis	Test statistic	1% Critical value	5% Critical value
P^{CER} and P^{EUA} are not cointegrated	3.801	-3.906	-3.341

Note: the null hypothesis of no cointegration is rejected if the test statistic is below the critical value. Critical values are taken from MacKinnon (1990, 2010).

Table 3: Results of the Johansen's cointegration tests

Null hypothesis	Trace Stat	5% critical value	Maximum eigenvalue stat	5% critical value
Specification with trend in both the cointegration equation and the error correction equation				
None	13.09	18.17	11.45	16.87
At most 1	1.64	3.74	1.64	3.74
Specification with trend in the cointegration equation				
None	26.38**	25.32	22.28**	18.96
At most 1	4.10	12.25	4.10	12.52
Specification with a constant in both the cointegration equation and the error correction equation				
None	22.62***	15.41	22.23***	14.07
At most 1	0.39	3.76	0.39	3.76
Specification with a constant in the cointegration equation				
None	28.00***	19.96	25.65***	15.67
At most 1	2.34	9.42	2.34	9.24
Specification without constant				
None	26.16***	12.53	24.19***	11.44
At most 1	1.96	3.84	1.96	3.84

Note: *** and ** respectively refer to the rejection of the null hypothesis at the 1% and 5% significance levels.

4 Method

In the case of an emissions trading scheme, the supply of permits is fixed, set by a cap that is decided at a political level, while the demand for permits is a function of the general economic activity and energy prices and might be influenced by the acceptance of international carbon credits. In contrast, for the latter, the supply is influenced by international energy prices as well as investment support and the demand is a function of the standard ETSs.

We first apply a standard VAR methodology to model the interrelationship between EUA and CER price variations over the whole time period of the analysis. We then introduce the more sophisticated methodology developed by Qu and Perron (2007) to allow for potential and a priori unknown regime changes in this multivariate system and identify them. Indeed, given the policy announcements that took place regarding CER acceptance in the second phase of the ETS and given the external factors (e.g. the 2008 economic crisis) that impacted the market, we expect changes in the regime of interactions between the two series during this phase. In a third step, we perform a structural analysis

(Granger causality tests, impulse response functions and forecast-error variance decomposition) inside each regime and discuss causality variations across regimes.

In the following, we present the methodology we use for the structural analysis of the causality relationship between the EUA and CER markets. We then explain how we employ Qu and Perron's approach to allow and detect regime changes in this relationship as well as select the best model in each identified regime.

4.1 Structural analysis of the EUA and CER relationship

The causality relationship between the EUA and CER prices (in logarithm) is first tested on the whole sample duration with a two-dimensional VAR model with two lags.[17] In principle, we can indeed expect bi-directional interactions between the two price series, as the demand for CER is likely to be largely driven the European carbon market, while the demand for EUA is also influenced by the use of CERs in the EU scheme. Given the impact of energy prices and the economic activity on the carbon price, we included these variables as exogenous controls, as presented below.

$$\begin{cases} \Delta P_t^{EUA} = \alpha_1 + \beta_1 \Delta P_{t-1}^{EUA} + \gamma_1 \Delta P_{t-2}^{EUA} + \delta_1 \Delta P_{t-1}^{CER} + \lambda_1 \Delta P_{t-2}^{CER} + \zeta_1 \Delta P_t^{gas} + \eta_1 \Delta P_t^{coal} + \theta_1 \Delta G_t + \varepsilon_{1t} \\ \\ \Delta P_t^{CER} = \alpha_2 + \beta_2 \Delta P_{t-1}^{EUA} + \gamma_2 \Delta P_{t-2}^{EUA} + \delta_2 \Delta P_{t-1}^{CER} + \lambda_2 \Delta P_{t-2}^{CER} + \zeta_2 \Delta P_t^{gas} + \eta_2 \Delta P_t^{coal} + \theta_2 \Delta G_t + \varepsilon_{2t} \end{cases}$$

where ΔP_t^{EUA}, ΔP_t^{CER}, ΔP_t^{gas}, ΔP_t^{coal} are respectively the price variations of EUA, CER, gas and coal in period t, ΔG_t the variation in the economic activity in period t, and ε_{1t} and ε_{2t} the error terms corresponding to each relationship.

We tested specifications with energy prices and the economic activity as endogenous or exogenous variables, and variations with one or two lags for these variables. On the basis of the Akaike and Bayesian information criteria, we found that including the energy prices and economic activity as endogenous variables does not provide additional information compared to the specifications for which they are included as exogenous variables. In particular, the estimation with the alternative specifications show

[17] The number of lags is chosen according to the Akaike and Hannan-Quinn information criteria.

that the impact of EUA and CER prices on the energy prices and economic activity does not provide any justification for including them as endogenous terms. We keep the simplest specification (with the contemporaneous terms of energy prices and the economic activity included as exogenous variables) and report the results for this.[18]

We employ the estimation results from the VAR analysis to perform a Granger causality test, conduct an impulse response analysis and a decomposition of the forecast error variance (volatility decomposition). The Granger causality test indicates whether the lagged values of the price of one type of carbon permit improve the forecasting performance of the price of the other type. It informs on the effect of the past values of one variable on the current value of the other one.

We then simulate a shock on the EUA price and look at the impact on the CER price, and, symmetrically, simulate a shock on the CER price and examine the impact on the EUA price. In order to perform this impulse-response analysis, we employ the Cholesky decomposition to orthogonalize ε_1 and ε_2.

Finally, the forecast error variance decomposition tells us the proportion of the movement in the price sequence of one permit type that is due to its own shocks or due to shocks to the other variable. The variance decomposition of the EUA and CER prices indicates the share of the CER price volatility that is explained by the EUA price volatility and, symmetrically, the share of the EUA price volatility that is explained by the CER price volatility.

After detection of potential regime changes and model selection with Qu and Perron's approach, we perform a comparable analysis (Granger causality test, impulse-response functions and forecast-error variance decomposition) on each specific regime.

4.2 Detection of potential regime changes and model selection in a multivariate system

There are several well-documented methods to identify structural changes in a single regression model.[19] Only a limited number of studies adress the issue of structural breaks for multivariate systems. We follow Qu and Perron's approach (2007) to detect potential regime changes in the short-term relationship between the EUA and CER prices and select the relevant model for these multivariate regressions.

[18] The results with the alternative specifications are available upon request.

[19] For example, Bai and Perron (1998) develop a methodology to detect multiple breaks occurring at unknown dates in a single linear equation.

This method considers the number of breaks and their dates a priori unknown. They are detected rather than imposed in an ad hoc manner. Ahamada and Diaz Sanchez (2013) provide a very good summary of the method. We use the same structure and notations to summarize the method below but refer to the original paper for the details.

4.2.1 The model

We consider a VAR model with two real components, $y_t = (y_{1t}, y_{2t})' \epsilon \Re^2$, and one lag.Below, T and m respectively represent the sample size and the total number of structural changes. The vector $\Gamma = (T_1, ..., T_m)$ is composed of the unknown break dates and we set $T_0 = 1$ and $T_{m+1} = T$. This means that there are $m+1$ unknown subperiods : $T_{j-1} + 1 \leq t \leq T_j$, with $1 \leq j \leq m+1$. We can present the model as:

$$y_t = \pi_{j0} + \pi_{j1} y_{t-1} + \varepsilon_t \tag{1}$$

where:

$y_t = (y_{1t}, y_{2t})' \epsilon \Re^2$,

$\pi_{j0} = (\pi_{j0}^{(i)})_{i=1,2}\ \epsilon \Re^2$ is the vector of constant parameters and varies between subperiods,

$\pi_{j1} = (\pi_{j1}^{(kl)})_{k=1,2;\ l=1,2}$ indicates the 2×2 matrix of the VAR parameters and varies between subperiods,

and $\varepsilon_t\ \epsilon \Re^2$ are the residuals with mean zero and covariance matrix denoted by Σ_j.

We aim at estimating $\Lambda = \left\{ \widehat{m}, \widehat{T}_1, ..., \widehat{T}_m, \widehat{\beta}_{j=1,..,m+1}, \widehat{\Sigma}_{j=1,...,m+1} \right\}$ where:

$\beta_j = \left(\pi_{j0}^{(1)}, \pi_{j1,}^{(11)} \pi_{j1}^{(12)}, \pi_{j0}^{(2)}, \pi_{j1,}^{(21)} \pi_{j1}^{(22)} \right)' \epsilon \Re^6$.

We first assume that m is known and discuss its estimation in Section 4.2.2. To estimate the model introduced in Equation (1), we present it as:

$$y_t = x_t' \beta_j + \varepsilon_t \tag{2}$$

where $x_t' = (I_2 \otimes (1, y_{1t-1}, y_{2t-1}))$. We employ the restricted quasi-maximum likelihood estimation method. Given the break dates $\Gamma = (T_1, ..., T_m)$, the Gaussian quasi-likelihood function is:

$$LR_T = \frac{\Pi_{j=1}^{m+1} \Pi_{t=T_{j-1}+1}^{T_j} f\left(y_t | x_t; \beta_j, \Sigma_j\right)}{\Pi_{j=1}^{m+1} \Pi_{t=T_{j-1}^0+1}^{T_j^0} f\left(y_t | x_t; \beta_j^0, \Sigma_j^0\right)} \tag{3}$$

where $f\left(y_t | x_t; \beta_j, \Sigma_j\right) = (2\pi)^{-n/2} \left|\Sigma_j\right|^{-1/2} \exp\left\{-\frac{1}{2}\left[y_t - x_t'\beta_j\right]' \Sigma_j^{-1}\left[y_t - x_t'\beta_j\right]\right\}$ with $n = 2$.

$\Gamma^0 = (T_1^0, ..., T_m^0)$, β_j^0 and Σ_j^0 represent the true unknown parameters. The denominator of (3) is supposed to be constant. The Qu and Perron's approach allows restriction in $\beta = (\beta_1', ..., \beta_{m+1}')'$ and $\Sigma = (\Sigma_1, ..., \Sigma_{m+1})$. For instance, imposing that a subset of β_j is kept identical for all j would imply a partial structural change model. We represent by $g\left(\beta, vec(\Sigma)\right) = 0$ the form of some restrictions in β and/or Σ where $g(.)$ is an r-dimensional vector and r the number of restrictions. We deduct the restricted log-likelihood ratio:

$$rlr_T = \log(LR_T) + \lambda' g\left(\beta, vec(\Sigma)\right) \tag{4}$$

and the estimates:

$$\{\widehat{T}_1, ..., \widehat{T}_m, \widehat{\beta}, \widehat{\Sigma}\} = \arg \max_{(T_1, ..., T_m, \beta, \Sigma)} rlr_T \tag{5}$$

The maximization (5) is conducted over all partitions $\Gamma = (T_1, ..., T_m)$ such that $|T_j - T_{j-1}| \geq [\delta T]$ and $T_m \leq [T(1-\delta)]$ where δ is an arbitrary small positive number and [] represents the integer part of argument. As trimming parameter, δ imposes a minimal time-length for each regime. A major implication is that, "under more general assumptions, the estimates of the break dates $\Gamma = (T_1, ..., T_m)$ and the coefficients (β, Σ) are asymptotically independent and valid restrictions on the latter do not affect the distribution of the former." This result is useful for the model selection and the inclusion of exogenous regressors in our application.

4.2.2 Selection of the number of breaks

In order to identify m, the number of breaks, we use the likelihood ratio test of no structural change versus a particular number of changes k. Following Qu and Perron (2007), this can be defined as:

$$\sup LR_T\left(k, p_b, n_{bd}, n_{bo}, \varepsilon\right) = 2\left[\log \widehat{L}_T\left(\widehat{T}_1, ..., \widehat{T}_k\right) - \log \widetilde{L}_T\right]$$

where $\log \widehat{L}_T\left(\widehat{T}_1, ..., \widehat{T}_k\right)$ is the maximum log-likelihood found with the optimal partition $\{\widehat{T}_1, ..., \widehat{T}_k\}$,

$\log \widetilde{L}_T$ is the maximum "log-likelihood under the null hypothesis of no structural change, p_b the total number of coefficients that can change (not including the coefficients of the variance-covariance matrix), n_{bd} and n_{bo} indicate, respectively, the number of parameters that can vary in the diagonal and off-diagonal coefficients of the variance-covariance matrix." ε imposes a minimal time-length for each regime. The limiting distribution of $\sup LR_T$ is discussed in more details by the authors and is function of the parameters described above. "This testing procedure can adjust to a diversity of types of structural changes: (a) changes only in the coefficients of the conditional mean ($n_{bd} = 0, n_{bo} = 0$); (b) changes only in the coefficients of the covariance matrix of residuals ($p_b = 0$); (c) changes in all of the coefficients ($p_b \neq 0, n_{bd} \neq 0, n_{bo} \neq 0$).

The test for no change versus an unknown number of breaks can also be considered given some upper-bound M for k. These types of tests are called double maximum tests and the statistic is defined for some fixed weights $W = \{a_1, ..., a_M\}$ as:"

$$D \ \max \ LR_T(M) = \max_{1 \leq k \leq M} \left[a_m \sup LR_T\left(k, p_b, n_{bd}, n_{bo}, \varepsilon\right)\right].$$

The weights $W = \{a_1, ..., a_M\}$ reflect the imposition of some priors on the likelihood of various numbers of structural breaks. Following Bai and Perron (1998), uniform double maximum tests (UD) ($a_i = 1$ for $1 \leq i = 1 \leq M$,) and weighted double maximum tests (WD) can be employed. For more explanations about these types of tests, we refer to the originql paper.

Another possibility is to use the so-called sequential test based on the null hypothesis of l breaks dates versus $l + 1$ breaks. The statistic can be written as:

$$SEQ_T\left(l + 1|l\right) = \max_{1 \leq j \leq l+1} \ \sup_{\tau \epsilon \Lambda_{j,\varepsilon}} lr_T\left(\widehat{T}_1, ..., \widehat{T}_{j-1}, \tau, \widehat{T}_j, ..., \widehat{T}_l\right) - lr_T\left(\widehat{T}_1, ..., \widehat{T}_l\right)$$

where:

$lr_T(.)$ denotes the log of the likelihood ratio,

$\{\widehat{T}_1, ..., \widehat{T}_l\}$ is the optimal partition if we assume l breaks,

and $\Lambda_{j,\varepsilon}$ is the set of possible additional break dates given $(\widehat{T}_1, ..., \widehat{T}_l)$:

$\Lambda_{j,\varepsilon} = \{\tau; \widehat{T}_{j-1} + (\widehat{T}_j - \widehat{T}_{j-1})\varepsilon \leq \tau \leq \widehat{T}_j - (\widehat{T}_j - \widehat{T}_{j-1})\varepsilon\}$.

The limiting distribution of the test is a function of the number of coefficients that can vary. In practice, the preferred strategy to define the number of breaks is to first look at the $UD \max LR_T(M)$ or $WD \max LR_T(M)$ tests to know if there is at least one structural break. We then find the number of breaks by examining the $SEQ_T\,(l+1|l)$ statistics. We select the number of breaks m such that the tests $SEQ_T\,(l+1|l)$ are non-significant for any $l \geq m$. Bai and Perron (2003) recommend this method for empirical analyses.

4.2.3 Selection of the number of lags

The presence of structural changes in multivariate models introduces nonlinearities. The Akaike and Byesian information criteria (AIC and BIC) usually employed for model selection are hence not valid. To overcome this issue, we use a practical result from Qu and Perron (2007) which states that "the limited distribution of the estimates of the break points is not affected by the imposition of valid restrictions on the parameters." It follows that a non-linear VAR model with p lags can be considered as a valid restriction of a non-linear VAR model with $p+1$ lags, if the break-dates in the two models remain unchanged. We apply this result to select the number of lags in our empirical analysis.[20]

5 Results and discussion

5.1 Structural analysis on the whole time period

The results of the Granger causality tests on the whole time period are presented in Table 4. They provide indications on how knowing the price variations of one type of permits enables to better predict the price variation of the other type. While there would be reasons to expect a bi-directional relationship between the two price series (see Section 4.1), the analysis on the whole time period suggests a unidirectional influence from the EUA to the CER price. We find that the short-term variations in the EUA price cause variations in the CER price, but that the opposite is not true. The null hypothesis that variations in the price of EUAs do not cause variations in the price of CERs is rejected, while the hypothesis that variations in the price of CERs does not cause variations in the price of EUAs is not. These results seem to indicate that knowing the EUA price variations allows to

[20]Together with this practical result, one may use the modified Akaike (MAIC) and Bayesian (MBIC) information criteria developed by Kurozumi and Tuvaandorj (2011). These criteria enable to select the number of regressors and the number of structural breaks in multivariate regression models, possibly with lagged dependent variables as regressors and multiple structural changes in both the coefficients and the variance matrices. The modified BIC consistently selects the regressors and the number of breaks whereas both modified information criteria perform well in finite samples.

better predict the CER price variations whereas the knowledge of the past values of the CER price variations does not reduce the forecast error variance of the EUA price variations.

Table 4: Results of the Granger causality tests.

Null hypothesis	LR statistic	Granger causality test (Prob $>\chi^2$)
ΔP^{EUA} does not Granger cause ΔP^{CER}	17.21	0.000***
ΔP^{CER} does not Granger cause ΔP^{EUA}	3.75	0.153

Note: *** and ** respectively refer to rejection of the null hypothesis at the 1% and 5% significance levels.

Regarding the transmission of shocks between the EUA and CER prices, we use the Cholesky decomposition to proceed to an impulse-response analysis (see results in Figures 4 and 5) and to the variance decomposition of the two series (see results in Table 5). We observe that a shock on the EUA price is immediately transmitted to the CER price. This effect is amortized in two days and it disappears after four days. On the contrary, a shock on the CER price has no significant impact on the EUA price.

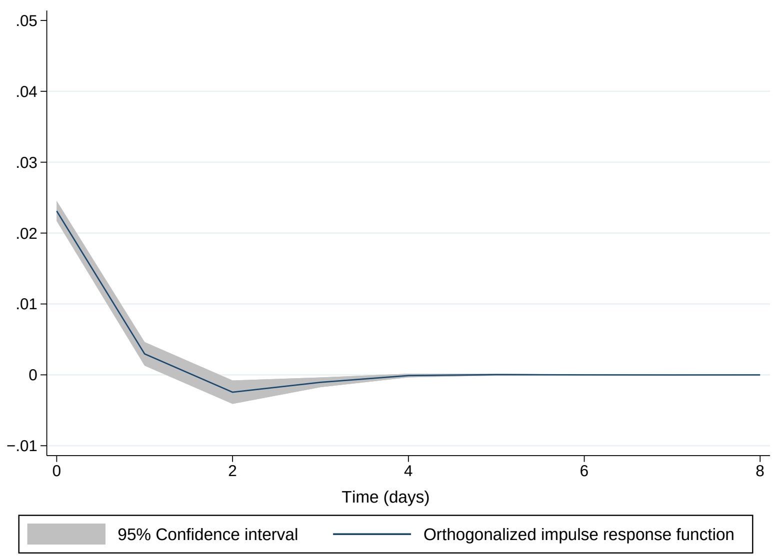


Figure 4: Response in the variation of the logarithmic CER price to an impulse in the variation of the logarithmic EUA price.

The variance decomposition of the EUA and CER prices indicates the share of the CER price volatility that is explained by the EUA price volatility and, symetrically, the share of the EUA price volatility that is explained by the CER price volatility. The results for the analysis on the whole time period are presented in Table 5. We find that the EUA price volatility explains 60% of the CER price volatility,

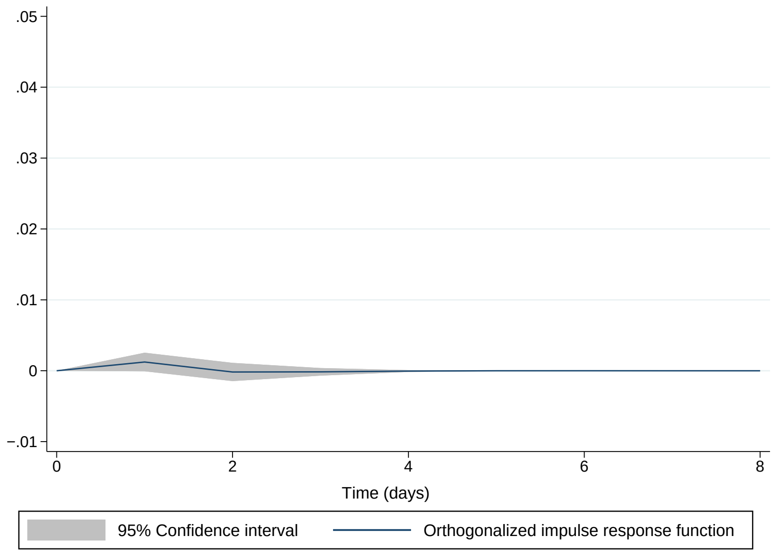


Figure 5: Response in the variation of the logarithmic CER price to an impulse in the variation of the logarithmic EUA price

while the CER price volatility has no impact on the EUA price volatility. In line with what is observed in the impulse-response analysis, the CER price variations do not seem to have a contemporaneous impact on the EUA price variations. This would lead to the conclusion that the EUA price variations evolve independently of the CER price variations.

Table 5: Variance decomposition of the forecasted errors for the whole time period.

	Variance decomposition of ΔP^{EUA}		Variance decomposition of ΔP^{CER}	
Days	ΔP^{EUA}	ΔP^{CER}	ΔP^{EUA}	ΔP^{CER}
1	100%	0%	60.48%	39.52%
2	99.72%	0.28%	58.96%	41.04%
3	99.72%	0.28%	59.00%	41.00%
4	99.71%	0.29%	59.03%	40.97%
5	99.71%	0.29%	59.02%	40.98%
6	99.71%	0.29%	59.02%	40.98%
7	99.71%	0.29%	59.02%	40.98%
8	99.71%	0.29%	59.02%	40.98%

All these results on the whole time period seem to suggest a unidirectional influence of the EUA on the CER. There are three possible explanations for this. First, the EUA market is much larger than the CER market: the number of EUA issued annually (more than 2 billion in 2013) is in the same order of magnitude as the cumulative number of CER generated since the mechanism was established (more than 1.3 billion indicated on the CDM pipeline[21] at the end of Phase II). Second, the demand for CER has come mainly from the EU ETS. Third, the volume of CER that could be used for compliance in the EU ETS was limited to 13% of the overall cap in the second phase of the scheme.

However, given the policy announcements regarding the acceptance of CER in the EU ETS during the second phase and given the external factors that impacted the European carbon market or the credit market (e.g. the 2008 economic crisis or changes in global fossil energy prices), we expect changes in the regime of interactions between the two price series. For this reason, in the following, we conduct a VAR analysis allowing for a priori unkown regime changes.

5.2 Detection of regime changes and model selection with the Qu and Perron's approach

We now apply the method developed by Qu and Perron to detect potential regime changes in the relationship between the EUA and CER price. For this, we consider the following VAR model with p lag length:

$$\begin{pmatrix} \Delta P_t^{EUA} \\ \Delta P_t^{CER} \end{pmatrix} = \begin{pmatrix} \pi_{j0}^{1} \\ \pi_{j0}^{2} \end{pmatrix} + \sum_{i=1}^{p} \begin{pmatrix} \pi_{ji}^{11} & \pi_{ji}^{12} \\ \pi_{ji}^{21} & \pi_{ji}^{22} \end{pmatrix} \begin{pmatrix} \Delta P_{t-i}^{EUA} \\ \Delta P_{t-i}^{CER} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \tag{6}$$

where:

$T_{j-1} + 1 \leq t \leq T_j,\ j = 1, ..., m+1,$

ΔP_t^{EUA} and ΔP_t^{CER} are respectively the EUA and CER price variations,

and $\begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}$ is the vector of the residuals with zero mean and constant covariance-matrix.

We only consider the case of a structural change in the coefficients of the VAR model equations. All the tests are carried out with $m = 5$. We first estimate a VAR model with one lag only. We also choose a trimming of 0.15.[22] This imposes a minimum time length of 177 observations for each regime. Since

[21] Detailed information on CDM projects is provided by the Centre on Energy, Climate and Sustainable Development: UNEP DTU CDM/JI Pipeline Analysis and Database (UNEP Risø Centre, 2013).

[22] The trimming value imposes a minimal length for each regime and the limiting distributions of the different tests are affected by this value. As a robustness exercise, we experiment with three different values: 0.10, 0.15, and 0.20. These values are considered in many empirical papers and Monte Carlo simulations. We find that results are statistically unaffected.

Table 6: The sequential test of (l+1) breaks against (l) breaks for a VAR(1) with no linear restrictions

	Statistic	5% critical value	Significant break	95% confidence interval for break dates
Seq(2|1)	3.85	22.64	November 22, 2011	[September 27, 2011; January 20, 2012]
Seq(3|2)	18.58	23.56		
Seq(4|3)	8.41	24.18		
Seq(4|5)	0.00	24.67		

we do not know the true lag length, we first estimate the number of breaks with the testing procedure presented in Section 4.2.2, a procedure which is robust to heteroskedasticity and autocorrelation. We then estimate the lag length using the estimated number of breaks. More precisely, we first test for the null of no break using the double maximum (WD max $LR_T(M)$) tests at the 5% significance level allowing different second moments of the regressors as well as the heterogeneity of the variances. If the null hypothesis is rejected, we use the sequential test based on the null hypothesis of l breaks dates versus $l+1$ breaks until it cannot reject the null hypothesis. Once the number of breaks m is estimated, the lag length is estimated through model selection stage which includes two main steps. The first step is based on an important result of the Qu and Perron's approach stating that the limiting distribution of the estimates of the break dates is unaffected by the imposition of valid restrictions on the other parameters of the model. In order to filter out models with invalid restrictions, we check whether the estimated break dates in reduced VARs are within the confidence intervals of those in the full VAR. The second step serves as a robustness check analysis and minimizes modified information criteria to select the most appropriate model among the candidates. We minimize the modified Akaike and Bayesian information criteria (MAIC and MBIC respectively), developed by Kurozumi and Tuvaandorj's (2011) with a number of breaks m already estimated. [23]

The results of the WD max $LR_T(M)$ test on a VAR model with lag length equal to 1 suggest the rejection of the null hypothesis of no structural change at the 5% level. The value of the WD max $LR_T(M)$ statistic is 60.14. It exceeds 21.83, the critical value at the 5% level. Hence, there is at least one structural break in the VAR model with lag length equal to 1. Table 6 shows the results of the $SEQ_T\ (l+1|l)$ test along with the estimated break date with its 95% confidence interval. We cannot reject the null hypothesis of one break against the alternative of two breaks as $SEQ_T(2|1) = 3.85$ with the 5% critical value being 22.64. The performed tests thus lead us to conclude that there is only one significant structural break in the considered VAR system with one lag.

[23]Kurozumi and Tuvaandorj's (2011) modified the Akaike and Bayesian information criteria developed by Akaike (1973) and Schwarz (1978) to allow comparing multivariate models with structural changes.

Afterwards, we extend the model to include additional lags. We apply the tests using a VAR with a lag fixed at $p = 2$. The value of the WD max $LR_T(M)$ statistic is 64.44, whereas the critical value at the 5% level is 28.72. We thus reject the null hypothesis of no structural change in favor of at least one break. According to the sequential $SEQ_T(l+1|l)$, we cannot reject the null hypothesis of one break against the alternative of two breaks as $SEQ_T(2|1) = 9.17$ is lower than the 5% critical value of 30.28. We find that the number and location of the break-date remain unchanged. The VAR model with one lag can then be considered as a valid and parsimonious restriction of the VAR model with two lags. The modified Akaike information criteria are $MAIC(1,1) = -12284.20$ and $MAIC(1,2) = -12284.06$ for VAR models respectively with one lag and two lags. The corresponding modified Bayesian information criteria are $MBIC(1,1) = -12199.95$ and $MBIC(1,2) = -12159.23$. Both modified information criteria suggest to select the VAR model with one lag and consequently confirm our previous conclusion about model selection. Table 6 reports the estimated break date and its corresponding 95% confidence interval.

The detected date of regime change, November 22, 2011 can be explained by the economic situation as well as policy announcements. At that time, as presented in Section 2, the carbon market was affected by the recession in Europe (third trimester of 2011). This induced a drop in the EUA and CER price series as well as an increase in their volatility from July 2011 onwards. In addition, at the 17^{th} COP in Durban in November 2011, the decision was taken to review existing market-based mechanisms and to develop new ones to help emerging and developing countries in their emission reduction efforts (KPMG, 2011). This followed announcements of stricter rules of acceptance of CDM credits in the EU ETS in the course of the year 2011. In January 2011, the European Commission published the list of credit types that would continue to be accepted under the EU ETS. This is part of stricter limitations on the use of offsets in the EU ETS from 2013 onwards, including a focus on credits coming from projects in least developed countries, and a ban of controversial industrial gas projects. In July 2011, on the occasion of the launch of the Sandbag's report *Buckle Up! 2011 Environmental Outlook for the EU ETS*, the Climate Action Commissioner indicated to the European Parliament (EC, 2011b) that the EU would push for a reform of this Kyoto Protocol offset mechanism.[24]

[24] All this happened in a context of decreasing demand for permits (due to limitations on the volume of credits that each installation covered by the European scheme could use for compliance in Phase II) and increasing supply, as explained in Section 2 (see Table 1). All together these events and announcements may explain the observed downward trend in the CER price from July 2011 onwards, the enhanced uncertainty regarding offsets and hence their larger price volatility.

5.3 Structural analysis with regime changes

The results of the Granger causality test on the time periods before and after the regime change are reported in Table 7. They show that, although for the whole time period, the dominant regime suggests a unidirectional influence of the EUA on the CER price variations, this is not the case for all subperiods. Not taking into account possible regime changes would result in neglecting non-linear effects. Before 22 November 2011, knowing the price variations of CERs helps to better predict the EUA price variations, but the opposite is not true. After 22 November 2011, the causality relationship follows the opposite direction. We provide an explanation for these differences between regimes based on the changes in the general economic activity, the policy announcements regarding CERs, and the changes in their demand and supply.

Table 7: Results of the Granger causality tests per subperiod.

Time period	Null hypothesis	LR statistic	Granger causality test (Prob $>\chi^2$)
Before 22 November 2011	ΔP^{EUA} does not Granger cause ΔP^{CER}	2.59	0.108
	ΔP^{CER} does not Granger cause ΔP^{EUA}	18.92	0.000***
After 22 November 2011	ΔP^{EUA} does not Granger cause ΔP^{CER}	6.10	0.014**
	ΔP^{CER} does not Granger cause ΔP^{EUA}	0.066	0.797

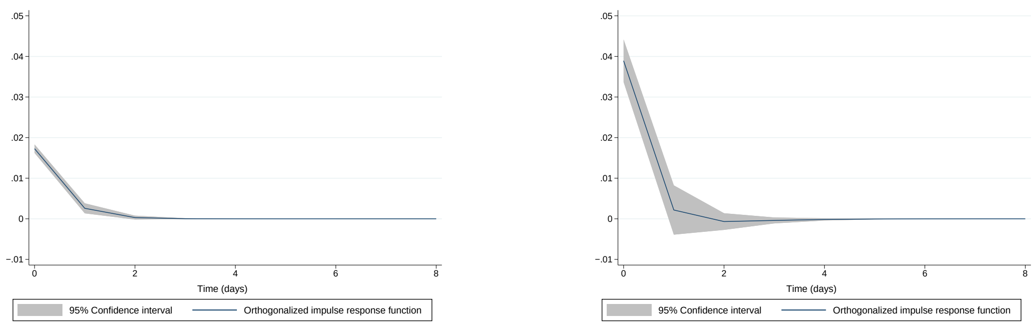
Note: *** and ** respectively refer to rejection of the null hypothesis at the 1% and 5% significance levels.

While the link between the EU ETS and the international credit market could involve a bi-directional influence between the EUA and CER price variations, we are now examining why the Granger causality relationship is uni-directional, from the CER to the EUA price before November 2011 and from the EUA to the CER price after November 2011.

For the time period after 22 November 2011 (T2), the Granger causality test indicates a unidirectional influence from the EUA to the CER price. As explained above, the recession in Europe in the third trimester of 2011, together with policy announcements regarding these Kyoto credits, reduced demand from EU ETS installations[25] and increased supply (as explained in Section 2) is likely to have caused the CER price drop and its volatility increase. This uncertainty regarding the value of the offset credits resulted in the EUA price becoming more independent from the CER price variations, hence the uni-directional influence from the EUA to the CER.

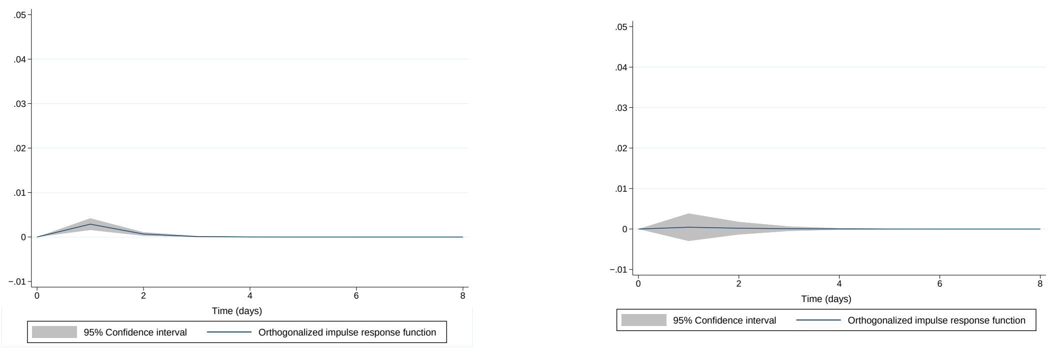
[25] Back then it was uncertain whether the limit of offsets that could be accepted for compliance for the second phase of the EU ETS would be reached.

For the time period before 22 July 2011 (T1), i.e. for the majority of the second phase of the EU ETS, the Granger causality test indicates that knowing the CER price variations helps to better predict the EUA price changes, while the opposite is not true. We explain the absence of a Granger causality relationship from the EUA to the CER by the impact of the 2008 economic crisis on the EU ETS. The economic crisis of 2008 impacted energy and carbon prices. This as particularly visible in the drop of the coal and gas prices until summer 2009 as well as in their respective volatilities which remained high until summer 2009 for coal and even until summer 2010 for gas (see Figure 3). The gas price and volatility were back to their pre-crisis level only during the course of the year 2011. While this induced a lack of scarcity for carbon permits, the expectations regarding CERs were still high as the second phase of the EU ETS marked the start of their acceptance in the European scheme. This would explain the uni-directional Granger causality relationship from the CER to the EUA.



(a) Before 22 November 2011 (b) After 22 November 2011

Figure 6: Response in the variation of the logarithmic CER price to an impulse in the variation of the logarithmic EUA price



(a) Before 22 November 2011 (b) After 22 November 2011

Figure 7: Response in the variation of the logarithmic EUA price to an impulse in the variation of the logarithmic CER price

We now proceed to the Cholesky decomposition and perform the impulse-response analysis as well as the volatility decomposition. For both subperiods, we observe that a shock on the EUA price is immediately transmitted to the CER series and amortized in two days (see Figure 6). The effect is the strongest for the time period after 22 November 2011, which coincides with the time when the Granger causality link is unidirectional from the EUA to the CER price series. On the contrary, an impulse in the CER price is not transmitted to the EUA series in the second regime (see Figure 7), while there is a slight transmission with a one-day delay in the first one. The shock is then absorbed in one day. We note that the first regime is the one during which we observe a Granger causal influence of the CER price on the EUA.

Table 8: Variance decomposition of the forecasted errors for the time period before 22 November 2011.

	Variance decomposition of ΔP^{EUA}		Variance decomposition of ΔP^{CER}	
Days	ΔP^{EUA}	ΔP^{CER}	ΔP^{EUA}	ΔP^{CER}
1	100%	0%	71.65%	28.35%
2	98.10%	1.90%	70.72%	29.28%
3	98.00%	2.00%	70.69%	29.31%
4	98.00%	2.00%	70.69%	29.31%
5	98.00%	2.00%	70.69%	29.31%
6	98.00%	2.00%	70.69%	29.31%
7	98.00%	2.00%	70.69%	29.31%
8	98.00%	2.00%	70.69%	29.31%

Table 9: Variance decomposition of the forecasted errors for the time period after 22 November 2011.

	Variance decomposition of ΔP^{EUA}		Variance decomposition of ΔP^{CER}	
Days	ΔP^{EUA}	ΔP^{CER}	ΔP^{EUA}	ΔP^{CER}
1	99.98%	0.02%	59.72%	40.28%
2	99.97%	0.03%	57.02%	42.98%
3	99.97%	0.03%	56.73%	43.27%
4	99.97%	0.03%	56.70%	43.30%
5	99.97%	0.03%	56.70%	43.30%
6	99.97%	0.03%	56.70%	43.30%
7	99.97%	0.03%	56.70%	43.30%
8	99.97%	0.03%	56.70%	43.30%

The forecast error variance decomposition shows a trend that is common to all subperiods and consistent with the observation on the whole time period (see Tables 8 and 9). Between 56% and 72% of the CER price volatility is explained by the EUA price volatility, while less than 2% of the EUA volatility is explained by the CER price volatility.

In conclusion, we observe a clear influence of the EUA price on the CER price with regard to the transmission of shocks and the effect on volatility. A shock on the EUA price is immediately transmitted to the CER price and absorbed within two days. The volatility of the CER price is largely explained by the EUA price volatility. On the contrary, we observed nearly no transmission of CER price shocks to the EUA price and the volatility of the CER price has a very limited influence on the volatility in the CER price. These observations are likely due to the market size difference between EUA and CER: the volume of EUAs issued each year during the second phase of the EU ETS was larger than the total volume of CERs issued since the establishment of the corresponding Kyoto Protocol mechanism. In addition, the limit on the volume of CER accepted in the EU ETS as well as the fact that the demand for CER had come mainly from the EU ETS reinforced the impact of this asymmetry between the two markets on the price interaction.

However, the Granger causality relationship changes between the two regimes. It is unidirectional from CERs to EUAs in first regime and from EUAs to CERs in the second one. The direction depends on the general economic activity as well as the expectations and potential restrictions regarding offsets. Before November 2011, we find a Granger causal impact from the CER to the EUA price variations. This can be explained by the fact that, while the 2008 economic crisis induced a lack of scarcity for permits, expectations regarding offsets were still high as the EU ETS phase II was the first time such credits were accepted in the European carbon market. On the contrary, after November 2011, we find a Granger causal effect from the EUA to the CER price variations. In 2011, the EU published the list of restrictions on the use of credits in European scheme and reviews of the corresponding Kyoto mechanism were decided in the international climate negotiations at the end of the year. As expectations regarding CERs were severely affected, the demand for credits from European installations declined. As the CER price dropped and its volatility increased, the EUA market became more independent of the international offset market.

These results may have policy implications both for the international climate negotiations and for ETS regulation. At a regional policy level, stricter restrictions on the acceptance of credits would involve a larger independence of the emission trading schemes. For the design of new market mechanisms under Article 6 of the Paris Agreement, stricter conditions for credit issuance (limit on the volume or higher quality requirements) would improve the reliability of such permits, increase their expected value, and also their potential impacts on the daily price variations in the standard emissions trading systems. However, in all cases, if the volume of credits remains small in comparison with the volume of certificates in the emission trading schemes, the price shocks of the credits do not have much influence on the volatility in the standard carbon markets. It is rather the latter which largely drives the offset price volatility.

6 Conclusion

In this paper, we conduct an empirical analysis of the short-term impacts of international carbon credits on standard emissions trading schemes. We take advantage of the European experience with accepting offsets for compliance in the second phase of the EU ETS. We employ vector-autoregressive models and allow for regime changes, which we detect with the approach developed by Qu and Perron. For each identified regime, we characterize the structural relationship between the price variations of each carbon permit type as well as the transmission of shocks between them and the impact of the volatility of each type of permit on the other one. To the best of our knowledge, this is the first time the structural relationship between daily prices variations of credits and European allowances is empirically tested taking into account possible regime changes. More generally, this work is the first empirical analysis of the impact of offsetting on the volatility in standard ETSs.

In principle, we could expect that the allowance and credit prices mutually influence one another. However, we find that the structural relationship between the price variations of allowances and credits is uni-directional and changes over time as a consequence of policy annoucements and expectations regarding credits. Before November 2011, the price causality relationship is from credits to allowances, while it is from allowances to credits afterwards. Before November 2011, while the 2008 economic crisis involved a lack of scarcity for permits, expectations regarding credits were still high as the second phase of the EU ETS was the first time that these credits were accepted in such a large trading scheme. At the end of the year 2011, this reversed. The fall 2011 corresponds to a time of economic recession in Europe but also to the COP17 UNFCCC meeting where the decision was made to review existing market-based mechanisms and to develop new ones to help developing countries in their emission reduction efforts. This follows announcements by the European Commission in the course of the year 2011 of stricter rules of acceptance of offsets in the European carbon market.

Regarding the transmission of shocks and the impact on the permit price volatilities, we find that the influence is mainly from allowances to credits. A shock on the allowance price is always transmitted to the credit price and absorbed in two days. A shock on the credit price is not transmitted to the allowance price in the regime after November 2011. It is slightly transmitted in the first regime and also absorbed in two days. The allowance price volatility explains between 56% and 72% of the credit price volatility, whereas the latter explains less than 2% of the former.

As the design of new carbon offseting mechanisms is being negotiated, in particular in the framework of Article 6.4 of the Paris Agreement, our work has the following policy implications. The absence of transmission of shocks from the offset market to the ETSs is a rather positive point. This is likely due to the difference in market size between the two types of permits. Regarding the interactions between their returns, setting a limit on the volume of international credits that can be accepted in an ETS would tend to reduce their influence. This would, however, reduce the demand for offsets and, in fine, undermine the support that such carbon offseting mechanisms are expected to provide for low-carbon projects in developing countries. Strict conditions for the issuance of these credits (potentially together with a global limit on the volume of international credits generated annually) would increase their environmental quality and improve their acceptability by ETSs. A significant demand from these would help avoiding the price of these credits to drop. It might increase their influence on ETSs but it would tend to mitigate the risk of offsets contributing to a reduction of the carbon price in the latter. This would hence reinforce the effectiveness of both types of climate policy instruments.

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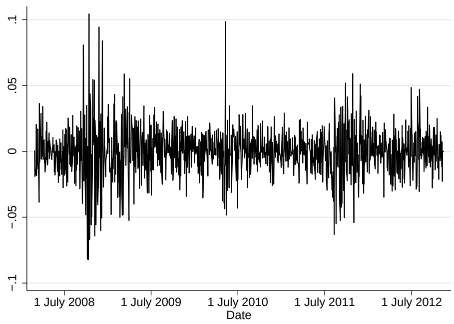
Appendices

A. Data description

Table 10: Descriptive statistics of the daily variations of the logarithmic price series

Variable	Nb. of Obs.	Mean	St. Dev.	Min.	Max.
EUA	1195	-0.00075	0.024	-0.093	0.193
CER	1182	-0.00234	0.031	-0.179	0.195
Gas	1195	0.0001495	0.03297	-0.1220	0.3600
Coal	1195	-0.0002768	0.02031	-0.2248	0.1631
Eurex	1195	-0.0003692	0.01823	-0.08208	0.1044

(a) Euro Stoxx 50 index



(b) Euro Stoxx 50 return

Figure 8: Euro Stoxx 50 index (in logarithm) and returns (first differences of the logarithmic series)

Table 11: Results of the Clemente Montañès and Reyes tests on EUA and CER permit prices (in logarithms)

Test	EUA future price				CER future price			
Procedure	IO		AO		IO		AO	
Series	Level	Variation	Level	Variation	Level	Variation	Level	Variation
DU_1	-0.016	0.002	-0.546	0.0036	-0.006	-0.005	-0.471	-0.021
	(-4.67)	(1.47)	(-49.46)	(1.955)	(-1.90)	(-0.669)	(-22.90)	(-2.79)
	{0.000}	{0.141}	{0.000}	{0.052}	{0.058}	{0.504}	{0.000}	{0.005}
DU_2	-0.016	0.0005	-0.606	0.0011	-0.006	-0.0003	-1.298	0.016
	(-4.82)	(0.287)	(-63.43)	(0.608)	(-1.39)	(-0.038)	(-72.74)	(2.08)
	{0.000}	{0.774}	{0.000}	{0.543}	{0.163}	{0.970}	{0.000}	{0.037}
ρ-1	-0.028	0.925	-0.034	-0.895	-0.005	-0.899	-0.014	-0.904
	(-5.36)	(-25.43)	(-4.67)	(-10.66)	(-1.427)	(-24.34)	(-2.473)	(-10.12)
	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-5.49]
Conclusion	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)
Dates of	13/10/08		03/11/08				21/11/08	23/11/11
Breaks	15/09/11		28/11/11				28/11/11	16/12/11

Note: The values in () and [] are respectively the t-statistics and the critical values at the 5% significance level tabulated by Clemente Montañès and Reyes. Values in {} are p-values.The null hypothesis of the unit root test is rejected when the t-statistic is smaller than the critical value.

Figure 9: Residuals of the regression of the CER price on the EUA price.

Table 12: Stationarity tests in the presence of breaks applied to the gas price series (in log)

Test	Clemente Montañès and Reyes test				Perron Vogelsang test			
Procedure	IO		AO		IO		AO	
Series	Level	Variation	Level	Variation	Level	Variation	Level	Variation
DU_1	-0.031***	-0.007**	-0.824***	-0.004	0.004**	-0.002	0.347***	10^{-5}
	(-5.37)	(-2.13)	(-65.4)	(-1.35)	(2.31)	(1.52)	(19.1)	(0.002)
DU_2	-0.026***	0.008***	0.665***	0.006***				
	(5.87)	(3.85)	(65.6)	(2.72)				
ρ-1	-0.034	-1.14	-0.036	-1.14	-0.006	-0.955	-0.057	-1.16
	(-5.74)	(-21.4)	(-3.89)	(-20.4)	(-2.13)	(-18.0)	(-1.77)	(-20.4)
	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-4.27]	[-4.27]	[-3.56]	[-3.56]
Conclusion	I(0)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)
Dates of	02/06/09	28/08/08	17/02/09	17/02/09	21/04/10		13/12/10	
Breaks	22/04/10	27/08/09	25/05/10					

Note: The values in () and [] are respectively the t-statistics and the critical values at the 5 % significance level, tabulated by Clemente Montañès and Reyes on the one hand, and by Perron and Vogelsang on the other. The null hypothesis of unit root is rejected when the t-statistic is smaller than the critical value. *, ** and *** respectively refer to the 10%, 5% and 1% significance levels of the estimated coefficients. The test procedure is sequential. We first apply the Clemente Montañès and Reyes test and check the significance of the estimated parameters of the dummy variables accounting for the structural breaks. If they are significant, we interpret the unit root test. Otherwise, we run the Perron-Vogelsang unit root test and check for the significance of the break before interpreting the unit root test. If the break date is also not significant, we apply standard unit root tests without breaks.

Table 13: Stationarity tests in the presence of breaks applied to the coal price series (in log)

Test	Clemente Montañès and Reyes test				Perron Vogelsang test			
Procedure	IO		AO		IO		AO	
Series	Level	Variation	Level	Variation	Level	Variation	Level	Variation
DU_1	-0.017***	-0.004	-0.700***	-0.006**	-0.006***	-0.002*	-0.458***	0.002*
	(-5.36)	(-1.55)	(-60.3)	(-2.37)	(-2.71)	(1.88)	(-25.1)	(1.72)
DU_2	0.008***	0.004**	0.391***	0.006***				
	(5.00)	(2.06)	(48.3)	(2.92)				
ρ-1	-0.019	-0.910	-0.029	-0.860	-0.007	-0.910	-0.005	-0.880
	(-5.56)	(-11.1)	(-3.85)	(-9.71)	(-2.87)	(-10.9)	(-1.88)	(-10.0)
	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-4.27]	[-4.27]	[-3.56]	[-3.56]
Conclusion	I(0)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)
Dates of	10/09/08	27/02/09	31/10/08	28/10/08	05/09/08		22/10/08	
Breaks	15/04/10		28/04/10	26/02/09				

Note: The values in () and [] are respectively the t-statistics and the critical values at the 5 % significance level, tabulated by Clemente Montañès and Reyes on the one hand, and by Perron and Vogelsang on the other. The null hypothesis of unit root is rejected when the t-statistic is smaller than the critical value. *, ** and *** respectively refer to the 10%, 5% and 1% significance levels of the estimated coefficients. The test procedure is sequential. We first apply the Clemente Montañès and Reyes test and check the significance of the estimated parameters of the dummy variables accounting for the structural breaks. If they are significant, we interpret the unit root test. Otherwise, we run the Perron-Vogelsang unit root test and check for the significance of the break before interpreting the unit root test. If the break date is also not significant, we apply standard unit root tests without breaks.

Table 14: Stationarity tests in the presence of breaks applied to the equity index series (in log)

Test	Clemente Montañès and Reyes test				Perron Vogelsang test			
Procedure	IO		AO		IO		AO	
Series	Level	Variation	Level	Variation	Level	Variation	Level	Variation
DU_1	0.002*	-0.046***	-0.021**	-0.046***	-0.002	0.002	-0.297***	0.002
	(1.81)	(-5.08)	(2.49)	(-4.43)	(-1.64)	(1.58)	(-32.1)	(1.44)
DU_2	-0.003**	0.047***	-0.183***	0.048***				
	(-2.24)	(5.31)	(-21.1)	(4.60)				
ρ-1	-0.013	-1.31	-0.013	-1.26	-0.012	-1.28	-0.013	-1.22
	(-3.05)	(-13.5)	(-2.84)	(-10.3)	(-2.93)	(-13.1)	(-2.35)	(-11.6)
	[-5.49]	[-5.49]	[-5.49]	[-5.49]	[-4.27]	[-4.27]	[-3.56]	[-3.56]
Conclusion	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)	I(0)
Dates of	06/07/11	02/10/08	13/08/09	03/10/08			13/10/08	
Breaks		09/10/08	19/07/11	08/10/08				

Note: The values in () and [] are respectively the t-statistics and the critical values at the 5 % significance level, tabulated by Clemente Montañès and Reyes on the one hand, and by Perron and Vogelsang on the other. The null hypothesis of unit root is rejected when the t-statistic is smaller than the critical value. *, ** and *** respectively refer to the 10%, 5% and 1% significance levels of the estimated coefficients. The test procedure is sequential. We first apply the Clemente Montañès and Reyes test and check the significance of the estimated parameters of the dummy variables accounting for the structural breaks. If they are significant, we interpret the unit root test. Otherwise, we run the Perron-Vogelsang unit root test and check for the significance of the break before interpreting the unit root test. If the break date is also not significant, we apply standard unit root tests without breaks.

Table 15: Unit root tests for the explanatory variables

	Augmented Dickey-Fuller (ADF)		Philipps-Perron (PP)		Kwiatkowski-Phillips -Schmidt-Shin (KPSS)	
Series (in logarithm)	Level	Variation	Level	Variation	Level	Variation
Gas	0.060(1)	-20.140(1)***	0.013(1)	-33.762(1)***	1.220(2)$^{\$\$\$}$	0.175(2)
Coal	-0.625(1)	-14.828(1)***	-0.653(1)	-33.401(1)***	0.646(2)$^{\$\$}$	0.154(2)
Eurex	-2.419(2)	-17.195(1)***	-2.753(2)*	-35.348(1)***	1.290(2)$^{\$\$\$}$	0.142(2)

Note: (1) model without constant or trend; (2) model with constant; (3) model with constant and trend. For ADF and PP tests, *, ** and *** respectively represent the rejection of the null hypothesis of a unit root at the 10, 5 and 1% significance levels. For the KPSS test, the null hypothesis of stationarity is rejected if the test statistic is above the critical value. With Barlett kernel and automatic selection of the number of lags, the critical values are 0.463 for 5% and 0.739 for 1% in the case without trend, 0.146 for 5% and 0.216 for 1% in the case with trend. For this test, $^{\$\$}$ and $^{\$\$\$}$ respectively represent the rejection of the null hypothesis of stationarity at the 5% and 1% significance levels. The model choice in ADF and PP tests is made according to a strategy of sequential tests from the most general to the most restricted one.

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