Economics Bulletin

Volume 29, Issue 2

Investigating efficiency in the U.S sulfur dioxide permit market

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Abstract

This paper constitutes –to our best knowledge– the first empirical analysis investigating long-run and short-run efficiency in the U.S sulfur dioxide (SO2) permit futures market with respect to its ability to unbiasedly predict future spot prices, using cointegration and error-correction models. Empirical results show that the market is inefficient, suggesting the presence of profitable arbitrage opportunities among U.S SO2 permit prices. In light of our findings, we recommend that market actors consider warily the information incorporated in SO2 futures prices.

I would like to thank Bénoit Sévi for his help in collecting sulfur dioxide spot price data
Citation: Mohamed Amine Boutaba, (2009) "Investigating efficiency in the U.S sulfur dioxide permit market", *Economics Bulletin*, Vol. 29 no.2 pp. 1308-1319.
Submitted: Mar 04 2009. **Published:** June 05, 2009.

1. Introduction

The 1990 Clean Air Act Amendements (CAAA) introduced the first large-scale cap-and-trade program for air pollution. Title IV of the CAAA established a system of tradable permits for sulfur dioxide (SO_2) emissions among utilities in the U.S. The aim of the system was 10 million tons per year reduction in SO_2 emissions from 1980 levels by the year 2010. Phase I (1995-1999) of the permit market extracts emission reductions from the 263 most polluting coal-fired electricity generating units with an output capacity greater than 100 megawatts (MW), belonging to 110 power plants located in 21 eastern and mid-western states. These 263 units, also called "Table A units", are allocated a fixed number of permits each year sufficient for an average emission rate of 2.5 pounds SO_2 per million Btu of average 1985-1987 heat input. Power plants may select units not originally affected until phase II to enter the program early as substituting or compensating units to help fulfil the compliance obligations for "Table A units" targeted by phase I. In addition, industrial emission sources, such as refineries and smelters, may voluntarily enter the program if they feel they can make emission reductions at low cost (opt-in units). Phase II which began in the year 2000, covers the remaining generating units fired by coal, oil and gas with an output capacity greater than 25 MW. Units are allocated permits sufficient for a more stringent average rate of 1.2 pounds of SO_2 per million Btu of average 1985-1987 heat input. The SO_2 permit trading program has dramatically reduced emissions faster and at far lower costs than anticipated, yielding wide-ranging environmental and human health benefits. Thereby, the SO_2 program's successes have encouraged policy makers in many countries to establish emissions trading schemes for other pollutants such as greenhouse gas emissions.

Since the passage of the 1990 CAAA, several studies have attempted to assess the efficiency of the SO_2 permit market with mixed results. Joskow et al. (1998) assess the efficiency of the market for SO2 permits by comparing the price of permits auctioned by the Environmental Protection Agency[1] (EPA) between 1993 and 1997 with prices associated with private confidential trades. Joskow et al. (1998) discover that by late 1994 these prices were almost identical and thereby conclude that the private market for tradable permits was relatively efficient. Schmalensee et al. (1998) also conclude that the private market for tradable permits was relatively efficient by noting the growth in the level of the trading volume from 1995 to 1997. Ellerman et al. (2000, pp. 185-190) conclude that the flattening of the term structure after 1995 provides evidence of a relatively market efficiency. Carlson et al. (2000) find that the market failed to realize potential gains from trade in the first two years of phase I. Ellerman (2003) and Ellerman et al. (2003) conclude that banking has played an important role in improving the economic and environmental performance of SO_2 cap-and-trade program. Arimura (2003) uses the coal price data from 1985 to 1998 and estimates a hedonic model in order to investigate the link between sulfur premium in coal and the permit price. In the first two years of the program, he finds that the sulfur premium was higher than the permit price in the EPA auction for the range of sulfur from 0 to 0.6 pound per mmbtu. Arimura (2003) argues that the deviation is due to the rent exploited by coal mine companies in the west from the high sulfur coal. For 1997 and 1998, however, the estimation results show that the permit price is in the range of 95% confidence interval for sulfur premium from 0 to 1 pound per mmbtu, suggesting that the permit price in the auction reflects the sulfur

[1] Since 1993, an auction of approximately 2.8% of the total annual permits was conducted on behalf of the Environmental Protection Agency (EPA) by the Chicago Board of Trade (CBoT). This auction is supposed to increase the market liquidity, to provide a price signal for private trades and to be an assured source of permit supply. Beginning in March 2006, CBoT decided to stop administering the auctions, resulting in EPA now conducting them directly.

premium of coal for low sulfur coal. From these results, Arimura (2003) concludes that the market is becoming efficient in 1997 and 1998. Using an output distance function approach, Swinton (2002, 2004) calculates the shadow prices of emission reductions and finds that they diverge among some power plants, suggesting that these plants have not taken full advantage of the permit market during much of phase I. Burtraw et al. (2005) suggest that this divergence of marginal abatement costs among some plants is due to the effects of implementing of electricity restructuring in some states which provided incentives to reduce costs. Keohane (2006) uses a unit –level econometric model of technique choice, based on actual decision by nearly 1000 units from 1995 to 1999, to estimate what would have happened if prescriptive regulation has been employed in place of an emissions trading scheme. The results show that cost savings appear to have been lower than estimated by others, noting that under the most natural choices of counterfactual regulations, the cost savings from trading, relative to a uniform emissions standard, ranged from $148 to $268 million annually: a cost savings of 16% to 25%. Ellerman and Montero (2007) show that the aggregate behavior of the SO_2 bank indicates that most participants have made reasonably efficient abatement decisions during the period 1995-2002. Helfand et al. (2007) discover that although the SO_2 price path does not reflect the Hotteling rule, profit opportunities appear relatively small and quite risky. They suggest that the SO_2 permit market appears to have been relatively efficient during the period 1994-2003.

However, research on SO_2 market efficiency from a financial market perspective is rather sparse. This is unfortunate given the characteristics of these trading permits which are similar to those of financial assets and energy. Indeed, the trading permits are perfectly homogeneous, like the financial assets[2], and their transaction does not generate transport and storage fees. In addition, as in the markets of energy, supply and demand of utilities covered by emissions trading scheme are stochastic and react to the change of permit price fundamentals such as the weather risk, the technological innovation in reducing emissions and fuel switching (Beaumais et al., 2008). Albrecht et al. (2006) examine the efficiency of the U.S SO_2 permit market from an informational point of view. They find that the random walk hypothesis and the economic profitable predictability are rejected, suggesting that the SO_2 market is weakly efficient. This paper constitutes –to our best knowledge– the first empirical analysis investigating long-run and short-run efficiency in the U.S SO_2 permit futures market with respect to its ability to unbiasedly predict future spot prices, using cointegration and error-correction models. Empirical results show that the market is inefficient, suggesting the presence of profitable arbitrage opportunities among U.S SO_2 permit prices. In light of our findings, we recommend that market actors consider warily the information incorporated in SO_2 futures prices. The remainder of the paper is organized as follows. In section 2, the methodology is described while in section 3 the data are presented. Empirical results and related findings are reported in section 4 and section 5 concludes the paper.

2. Methodology

A market is called efficient if prices always fully reflect available information (Fama, 1970). Hence, the opportunity for any abnormal gain on the basis on the information contained in historical prices is eliminated.

If market actors are risk neutral, then the current futures price should equal the expected future spot price with contract maturity. This implies:

[2] Kosobud et al. (2005) demonstrate empirically that SO_2 permits have rates of return and yield distribution that make them an asset option for inclusion in a risk diversified portfolio.

$E_{t-1}S_t = F_{t-1}$ (1)

where $E_{t-1}S_t$ is the expectation of the future spot price formed in period t-1, and F_{t-1} is the futures price with contract maturity in period t. Assuming rational expectations, so that $S_t = E_{t-1}(S_t|\Omega_{t-1}) + \mu_t$, where Ω_{t-1} represents the information set available in period t-1, μ_t is a rational expectations error and μ_t is orthogonal to all element in Ω_{t-1}, including lagged forecast errors, the hypothesis of efficiency (unbiasedness) is then tested by the following model:

$S_t = \alpha + \beta F_{t-1} + \mu_t$ (2)

The null hypothesis to be tested is the regression coefficients of the constant term and the futures price should not be statistically different from zero and one, respectively $[(\alpha, \beta) = (0,1)]$. This cannot be tested using standard regression analysis as price time series exhibit a non –stationary behavior. To avoid spurious regression results, the notion of cointegration can be used. According to Engle and Granger (1987), a linear combination of two or more non-stationary series (with the same order of integration) may be stationary. If such a stationary linear combination exists, the series are considered to be cointegrated and long-run equilibrium relationships exist.

Prior to proceeding with the cointegration test, we determine the order of integration of the variables and ensure that it is equal to one (I (1)) for each of the futures and spot price series. The conventional unit root tests namely Augmented Dickey–Fuller (1979, 1981) (ADF), Phillips–Perron (1987) (PP) and Kwiatkowski–Phillips–Schmidt–Shin (1992) (KPSS) are used to test for the stationarity of the series. However, Perron (1989) shows that usual unit root tests are subject to misspecification bias and size distortion when the series involved has undergone structural breaks leading to spurious acceptance of the unit root hypothesis. We overcome this limitation by also using Zivot and Andrews (1992) procedure to endogenously determine a break point and test for the presence of a unit root when the series have a broken trend.

In order to assess the dynamics of the U.S SO_2 futures markets and the corresponding spot markets, the Engle and Granger (1987) cointegration approach is implemented. The Engle-Granger method consists in estimating the cointegrating regression (eq.3) by ordinary least square, obtaining residuals and applying ADF tests for the residuals. The ADF test is as follows:

$\Delta \hat{u}_t = \phi \hat{u}_{t-1} + \sum_{i=1}^{p} \phi_i \Delta \hat{u}_{t-i} + \varepsilon_t$ (3)

where $\Delta \hat{u}_t$ includes the μ_t sequence and with the null hypothesis of $H_0 : \phi = 0$. The value of optimal lag length p is selected by the smallest Akaike information criterion or Schwartz Bayesian criterion. Since the residual series are calculated from a cointegrating equation, an intercept and time trend are not considered in the equation. The test statistics obtained is then compared against critical values in the table generated by Engle and Yoo (1987). If the variables are found to be cointegrated then some linear combination of them will be stationary. This means that there exists a long-run relationship among them.

If S_t and F_{t-1} are cointegrated and the joint restrictions hold in eq. (2), this implies that long-run unbiasedness and hence efficiency is substantiated. However, in the short-run it is possible that deviations exist from the long-run equilibrium relationship. Such short-run deviations would lead to both market inefficiency and speculative profit opportunities for arbitrageurs. Short-run efficiency can be tested using an error correction model (ECM) which captures the short-run dynamics of spot and futures prices. In our case, the ECM takes the following form:

$$\Delta S_t = \omega + \gamma(S_t - F_{t-1})_{-1} + \sum_{i=0}^{k} \delta_i \Delta F_{t-1,t-i} + \sum_{j=1}^{l} \varphi_j \Delta S_{t-j} + \eta_t \qquad (4)$$

where Δ is the difference operator, k and l are the numbers of lags, η_t is the serially uncorrelated error terms, and $(S_t - F_{t-1})_{-1}$ is the lagged error correction term (ECT), which is derived from the cointegration relationship and represents the deviations from the long-run equilibrium for the two prices. Thus, deviations in this period's price vary in relation to past disequilibria. Short term efficiency requires the following conditions to be satisfied:

i. $\omega = 0$

ii. $\delta_0 = 1$

iii. $\gamma = -1$

iv. all other δ and φ = 0.

If the four conditions are met, then markets are efficient and futures prices provide unbiased estimates of future spot prices both in the long-and short-run.

3. Data

The data used in this analysis are daily spot and futures prices for U.S SO_2 trading permits and are sourced from the Chicago Climate Futures Exchange (CCFE)[3]. Standards contracts are 25 tons of SO_2 emission permits. The spot closing prices for permits are collected on the OTC market and are calculated as an average of bids to buy and offers to sell for current vintages of permits. Nine futures contracts for delivery at maturities from December 2006 to 2014 are considered in the present study and the futures prices are matched with the corresponding spot prices. Sample lengths are December 10, 2004 – December 29, 2006 (SO_2 Dec 06 contract, 536 observations), December 10, 2004 – December 31, 2007 (SO_2 Dec 07 contract, 797 observations), December 10, 2004 – August 29, 2008 (SO_2 Dec 08 contract, 971 observations) and May 17, 2006 – August 29, 2008 (SO_2 Dec 09, Dec10, Dec 11, Dec 12, Dec 13 and Dec 14 contracts; 598 observations). As is customary in this type of analysis, all variables are used in their natural logarithms.

4. Empirical results

4.1. Testing for non-stationarity

An important first step in the analysis is to test the stationarity of the spot and futures price series. To ascertain the order of integration, we first used the conventional unit root tests. As shown in Table 1, the results of ADF, PP and KPSS unit root tests for levels and first differences show that none of the estimated variables are stationary while their differences are I (0). The results support the contention that futures and spot prices for the U.S SO_2 permits are I (1). Next the Zivot and Andrews unit root test is used in the analysis, which treats endogenously the presence of a structural break in the series. Table 4 reports the minimum t-statistics from testing the stationarity assuming a break in mean for the first differences of each futures and spot price series. The results confirm those from the conventional unit root tests that all series are I (1). The estimated breakpoints for futures prices of the contract for Dec 06 delivery and their corresponding spot prices are 26/01/2006 and 25/01/2006, respectively. The timings of these breakpoints are related to the downward adjustments of the

[3] New Mercantile Exchange (NYMEX) also provides SO_2 futures contracts. The use of data from the CCFE is justified by the high degree of liquidity in this trading platform.

Table 1. Results of conventional unit root tests

Series	ADF				PP		KPSS	
	Level		First difference		Level	First difference	Level	First difference
	Lag	Test statistic	Lag	Test statistic	Test statistic	Test statistic	Test statistic	Test statistic
Dec 06	1	-0.840 (1)	0	-21.522** (1)	-0.662 (1)	-22.498** (1)	0.824** (2)	0.353 (2)
Spot	2	-0.690 (1)	1	-13.522** (1)	-0.613 (1)	-18.985** (1)	0.834** (2)	0.347 (2)
Dec 07	1	-0.572 (1)	0	-27.071** (1)	-0.476 (1)	-27.480** (1)	1.683** (2)	0.141 (2)
Spot	2	-0.454 (1)	1	-17.632** (1)	-0.413 (1)	-24.648** (1)	1.698** (2)	0.137 (2)
Dec 08	2	-1.124 (1)	1	-19.071** (1)	-1.101(1)	-28.703** (1)	2.390** (2)	0.223 (2)
Spot	1	-1.232 (1)	0	-30.327** (1)	-1.108 (1)	-30.713** (1)	2.462** (2)	0.234 (2)
Dec 09	2	-1.009 (1)	1	-14.991** (1)	-0.972 (1)	-21.318** (1)	1.724** (2)	0.208 (2)
Dec 10	1	-1.217 (1)	0	-23.317** (1)	-1.137 (1)	-23.020** (1)	1.851** (2)	0.259 (2)
Dec 11	1	-1.046 (1)	0	-30.368** (1)	-0.871 (1)	-30.396** (1)	1.776** (2)	0.226 (2)
Dec 12	1	-1.159 (1)	0	-23.875** (1)	-0.945 (1)	-23.887** (1)	1.750** (2)	0.364 (2)
Dec 13	2	-1.130 (1)	1	-16.291** (1)	-0.876 (1)	-21.962** (1)	1.794** (2)	0.371 (2)
Dec 14	2	-1.052 (1)	1	-16.545** (1)	-0.776 (1)	-22.298** (1)	1.717** (2)	0.403 (2)
Spot	1	-1.203 (1)	0	-24.391** (1)	-1.107 (1)	-24.525** (1)	1.626** (2)	0.191 (2)

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

Table 2. Zivot-Andrews minimum *t*-statistics

Series	*t*-statistic	Period
Dec 06	-2.714 (4)	26/01/2006
Spot	-2.838 (4)	25/01/2006
Dec 07	-2.947 (4)	28/03/2006
Spot	-3.134 (5)	01/02/2006
Dec 08	-2.595 (5)	01/02/2008
Spot	-2.530 (5)	04/02/2008
Dec 09	-3.812 (3)	23/04/2008
Dec 10	-3.555 (3)	13/04/2007
Dec 11	-3.589 (3)	13/04/2007
Dec 12	-3.232 (3)	09/04/2007
Dec 13	-3.354 (3)	09/04/2007
Dec 14	-3.379 (3)	09/04/2007
Spot	-3.801 (3)	24/04/2008
ΔDec 06	-8.669*** (4)	09/12/2005
ΔSpot	-9.629*** (3)	08/12/2005
ΔDec 07	-12.799*** (3)	24/01/2006
ΔSpot	-12.456*** (3)	08/12/2005
ΔDec 08	-13.102*** (4)	08/01/2007
ΔSpot	-13.209*** (4)	08/01/2007
ΔDec 09	-10.953*** (4)	08/01/2007
ΔDec 10	-13.432*** (2)	08/01/2007
ΔDec 11	-22.930*** (0)	08/01/2007
ΔDec 12	-24.120*** (0)	08/01/2007

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Table 2. Continued

ΔDec 13	-11.859*** (3)	08/01/2007
ΔDec 14	-12.015*** (3)	08/01/2007
ΔSpot	-10.838*** (4)	08/01/2007

Notes: All *t*-statistics estimated from a break in intercept model. Values in parentheses are lag lengths used in the test for each series. Critical values are -5.43 (1%) and -4.80 (5%). *** denotes significance at 1% significance level.

expected marginal cost of reducing SO_2 emissions under the Clean Air Interstate Rule[4] (CAIR) as buyers and sellers more completely assess market fundamentals and to the dramatically increase of temperature on January 2006, which was the warmest January on record. The estimated breakpoints for futures prices of the contract for Dec 07 delivery and their corresponding spot prices are 28/03/2006 and 01/02/2006, respectively. The timings of these structural breaks are explained by their proximity to the dramatically rise of temperature on January 2006 and to the downward adjustments by the market actors of the expected marginal cost of reducing SO_2 emissions under CAIR. The estimated breakpoints for futures prices of the contract for Dec 08 delivery and their corresponding spot prices are 01/02/2008 and 04/02/2008, respectively. The timings of these breakpoints are explained by their proximity to the decision of the United States Court of Appeals for the district of Columbia Circuit vacating the Clean Air Mercury Rule on February 8, 2008. The estimated breakpoints for futures prices of the contracts for Dec 09, Dec10, Dec11, Dec12, Dec13 and Dec14 deliveries and their corresponding spot prices are 23/04/2008, 13/04/2007, 13/04/2007, 09/04/2007, 09/04/2007, 0904/2007, and 24/04/2008, respectively. While the timings of the estimated breakpoints of futures prices of the contract for Dec 09 delivery and their corresponding spot prices are explained by their proximity to the decision of the United States Court of Appeals for the district of Columbia Circuit invalidating the Clean Air Mercury Rule[5] (CAMR), the timings of the estimated breakpoint of futures prices of contracts for delivery at matunities from Dec 10 to Dec 14 and their corresponding spot prices are explained by their proximity to the supreme court decision on April 2, 2007, which named carbon dioxide and other greenhouse gases as air pollutants as defined in the Clean Air Act and therefore authorized the EPA to regulate greenhouse emissions from new automobiles and trucks.

4.2. Testing for cointegration

Having established that all of the futures and spot prices are I (1), we now proceed to the cointegration analysis. For each model, we include dummy variables in order to take into account possible structural changes and therefore to filter outliers in the time series. Three dummy variables taking respectively the value of 1 in November, 2005, December, 2005 and January, 2006 and zero otherwise are included in Dec 06 model. These dummy variables account for the downward adjustments by the market actors of the marginal cost of reducing SO_2 emissions under CAIR. A dummy variable taking the value of 1 from January 24, 2006 to February 28, 2006 and 0 otherwise is included in Dec 07 model and accounts for the

[4] On March 10, 2005, following the success of the CAAA of 1990, the EPA promulgated the CAIR that would dramatically reduce the SO_2 emissions that move across state boundaries in the 28 Eastern states and the District of Columbia in 2010 by over 70 percent in 2015 from 2003 levels.

[5] The Clean Air Mercury Rule, finalized on May 18, 2005 builds upon the CAIR to permanently cap and significantly reduce mercury emissions from coal-fired power plants, the largest remaining sources of mercury emissions in the USA, by nearly 70 percent from 1999 emission levels.

downward adjustments by the market actors of the marginal cost of reducing SO_2 emissions under CAIR. Two dummy variables taking respectively the value of 1 in February, 2008 and from July 11, 2008 to August 11, 2008 and 0 otherwise are included in the rest of models. These dummy variables account for the vacation of the Clean Air Mercury Rule and the Clean Air Interstate Rule, respectively. The test results are reported in Table 3. In all models, we clearly reject the null of no cointegration at 5% level. Thus each of the futures and the corresponding spot price series for each of the time spreads are cointegrated at the 5% significance level.

Table 3. Engle and Granger cointegration test

Models	ADF	p
Dec 06 futures and spot price series	-6.195	3
Dec 07 futures and spot price series	-4.379	3
Dec 08 futures and spot price series	-3.970	4
Dec 09 futures and spot price series	-3.446	2
Dec 10 futures and spot price series	-4.840	1
Dec 11 futures and spot price series	-4.366	2
Dec 12 futures and spot price series	-3.803	1
Dec 13 futures and spot price series	-3.589	1
Dec 14 futures and spot price series	-5.237	0

Note: The 5% critical values are -3.37 (p=0) and -3.25 (p=4) (Engle and yoo, 1987). Lag lengths p are selected by the smallest Akaike Information criterion.

4.3. Testing the unbiased expectations hypothesis

The nature of the long-run relationships from the cointegrating vectors is examined by testing the restrictions placed on eq. (2); that is the null hypothesis is $\alpha = 0$ and $\beta = 1$. Table 4 reports the estimated coefficients of the cointegrating vectors and test statistics of the unbiased expectations hypothesis.

The null hypothesis of unbiasedness is rejected at the 5% significance level for all contracts. This result supports the contention that SO_2 futures contracts for delivery at maturities from December 2006 to 2014 are biased predictors of the subsequent spot prices. This is consistent with a long-run inefficiency and a risk-premium paid to speculators.

Table 4. Wald test of parameter restrictions on the cointegrating vectors

	$\hat{\alpha}$	$\hat{\beta}$	Testing $\alpha = 0$ and $\beta = 1$		Testing $\alpha = 0$		Testing $\beta = 1$	
			F-statistic	p-value	F-statistic	p-value	F-statistic	p-value
Dec 06	0.153	0.974	129.99	0.000	24.702	0.000	29.741	0.000
Dec 07	-0.103	1.011	414.448	0.000	20.745	0.000	10.304	0.001
Dec 08	-0.015	0.995	663.786	0.000	0.547	0.460	2.464	0.117
Dec 09	0.172	0.958	1790.013	0.000	27.327	0.000	65.292	0.000
Dec 10	-0.435	1.160	17515.35	0.000	53.228	0.000	234.631	0.000
Dec 11	-0.482	1.171	11129.24	0.000	37.517	0.000	153.051	0.000
Dec 12	-0.129	1.110	8578.030	0.000	2.189	0.000	5264.99	0.000
Dec 13	-0.121	1.110	7632.054	0.000	1.653	0.199	44.142	0.000
Dec 14	-0.089	1.105	6785.151	0.000	0.772	0.379	34.896	0.000

We test individually the hypothesis that $\alpha = 0$ and $\beta = 1$ in order to determine whether the rejection of the joint hypothesis for these contracts was driven by the presence of a risk premium or a bias in the futures prices. Results in table 4 show that the hypothesis that $\alpha = 0$ and $\beta = 1$ is rejected separately for all contracts, except for contact for delivery in Dec 08. These results imply complex interaction between expectations and a possible risk premium, which occur in the process of the futures prices formation for these time spreads. In the light of these findings, we suggest that speculative opportunities seem to exist for these futures contracts.

In general, it would not be expected that a market which was not exhibiting long-run efficiency would also not be short-run efficient. Hence, it would be assumed that all contracts would also not be short-run efficient. Eq. (4) was estimated for the nine contracts and the coefficients are tested to verify if they are consistent with the criteria noted in section 2. The results of these estimations are reported in Table 5[6]. On the basis of F-statistics, the restrictions that $\omega = 0$, $\delta_0 = \beta = 1$ and $\gamma = -1$ were tested separately. The estimated values of ω are insignificantly different from zero for all contracts. The p-values of Wald statistics for the restriction $\delta_0 = 1$ are such that the restriction can be rejected at the 5% level for contracts for Dec06, Dec07, Dec08, Dec11, Dec12 and Dec 14 deliveries. The estimated values of γ are significantly different from -1 for all contracts. A joint test of all three restrictions is also performed using Wald test. The results confirm the general pattern of those from analyzing the restrictions individually. These results strongly suggest that there are short-run deviations from the long-run efficiency conditions and therefore the existence of a short-run inefficiency.

Table 5. Wald test of parameter restrictions on the ECMs

$$\Delta S_t = \omega + \gamma(S_t - F_{t-1})_{-1} + \sum_{i=0}^{k} \delta_i \Delta F_{t-1,t-i} + \sum_{j=1}^{l} \varphi_j \Delta S_{t-j} + \eta_t$$

	$\omega = 0$	$\delta_0 = \beta = 1$	$\gamma = -1$	F-statistic
Dec 06	0.00074	0.634	-0.256	141.515
	(0.275)	(0.000)	(0.000)	(0.000)
Dec 07	0.00038	0.650	-0.104	727.710
	(0.488)	(0.000)	(0.000)	(0.000)
Dec 08	0.00017	0.891	-0.134	968.518
	(0.780)	(0.000)	(0.000)	(0.000)
Dec 09	-0.000005	0.988	-0.125	682.548
	(0.945)	(0.439)	(0.000)	(0.000)
Dec 10	0.0000078	1.009	-0.119	510.007
	(0.954)	(0.817)	(0.000)	(0.000)
Dec 11	-0.0004	0.522	-0.119	519.831
	(0.794)	(0.000)	(0.000)	(0.000)

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[6] For each model, we include dummy variables in order to take into account possible structural changes and therefore to filter outliers in the time series. Two dummy variables taking respectively the value of 1 from 1/24/2006 to 2/28/2006 and from 4/24/2006 to 05/26/2006 and zero otherwise are included in Dec 06 and Dec 07 models in order to account for the downward adjustments by the market actors of the marginal cost of reducing SO_2 emissions under CAIR. Two dummy variables taking the value of 1 from 1/24/2006 to 28/2/2006 and from 7/3/2006 to 7/31/2006 and zero otherwise are included in Dec 08 and Dec 09 models, respectively. These dummy variables account for the downward adjustments by the market actors of the marginal cost of reducing SO_2 emissions under CAIR. A dummy variable taking the value of 1 from 07/11/2008 to 07/31/2008 and 0 otherwise is included in Dec 10, Dec 11, Dec 12, Dec 13 and Dec 14 models in order to take into account the court decision striking down the Clean Air Interstate Rule.

Table 5. Continued

Dec 12	-0.00039	0.912	-0.062	1112.467
	(0.786)	(0.029)	(0.000)	(0.000)
Dec 13	-0.00048	0.981	-0.051	1391.897
	(0.731)	(0.604)	(0.000)	(0.000)
Dec 14	-0.00067	0.906	-0.049	1476.49
	(0.649)	(0.021)	(0.000)	(0.000)

Notes: numbers in parentheses are p-values. The F-statistic is for the restriction $\omega = 0$, $\delta_0 = 1$ and $\gamma = -1$.

5. Conclusion

This paper extends the literature investigating the efficiency in the U.S. SO_2 permit market by testing the hypothesis of unbiasedness of futures prices in both the long-and short-run using the cointegration approach. We find that U.S SO_2 permits futures prices are cointegrated with subsequent spot prices for nine contracts for delivery at maturities from December 2006 to 2014. Spot and futures prices are determined by the same fundamentals and so cointegration implies the existence of one long-run relationship between futures price series and the corresponding spot prices across all of these time spreads.

While the presence of cointegration between spot and futures prices by itself fulfils a necessary condition for market efficiency but not sufficient condition for the unbiasedness of futures prices. Hence the unbiased expectations hypotheses are tested for all contracts and the results reveal that markets are inefficient and futures prices provide biased estimates of future spot prices in the long-run and the short-run. These findings support the existence of a risk premium and speculative opportunities for arbitrageurs and can be explained by the Keynesian theory of normal backwardation. According to Keynes (1930), futures prices are unreliable estimates of expected future spot prices. This implies that even if future spot price is expected to remain the same as the current spot price, the futures price will be below the expected spot price by an amount equal to a risk premium. This risk premium is paid by hedgers because this reduces risk to the hedger, just as it adds to the risk for speculator. Keynes' suggestion is based upon the argument that the long (short) speculator realizes the premium by refusing to purchase a contract from the short (long) hedger except at a price below (above) that which the futures price is expected to approach.

References

Albrecht, J., and M. De Clercq (2006) "Informational Efficiency of the U.S. SO_2 Permit Market", *Environmental Modeling & Software* **21**, 1471–1478.

Arimura T. H. (2003) "Pollution-Permit Market Efficiency: Hedonic Prices of Sulfur in Coal under the U.S. SO_2 Allowance Market", *Asahigarasu Zaidan Jyosei Kenkyu Seika Hokoku* (web only).

Beaumais, O., Boutaba, M.A., and S. Lardic (2008) "Permit Price Dynamics in the U.S SO_2 Trading Scheme: A Cointegration Approach", *Thematic Meeting of the French Economic Association 2008: Frontiers in Environmental Economics and Natural Resources Management*, Toulouse, France, 9-11 June, 2008.

Burtraw D., Evans D. A., Krupnick A., Palmer K., and R. Toth (2005) “Economics of Pollution Trading for SO_2 and NO_x”, *Annual Review of Environmental and Resources* **30**, 253–289.

Carlson C. P., Burtraw D., Cropper M. and K. Palmer (2000) “SO_2 Control by Electric Utilities: What are the Gains from Trade?” *Journal of Political Economy* **108**, 1292–1326.

Dickey D.A., and W.A. Fuller (1979) “Distribution of the Estimators for Autoregressive Time Series with a Unit Root”, *Journal of the American Statistical Association* **74**, 427–431.

Dickey D.A., and W.A. Fuller (1981) “Likelihood Ration Statistics for Autoregressive Time Series with a Unit Root”, *Econometrica* **49**, 1057–1072.

Ellerman A. D., Montero J. P., Joskow P. L., Schmalensee R., and E. M. Bailey (2000) “*Markets for Clean Air: The US Acid Rain Program*”, Cambridge University Press: New York.

Ellerman A. D. (2003) “Ex Post Evaluation of Tradable Permits: the U.S SO_2 Cap-and-trade Program”, MIT Center for Energy and Environmental Policy Research Working Paper 03-003.

Ellerman, A.D., Joskow, P.L., and D. Harrison (2003) “Emissions Trading: Experience, Lessons, and Considerations for Greenhouse Gases”, Pew Center for Global Climate Change: Washington, DC.

Ellerman A. D., and J. P. Montero (2007) “The Efficiency and Robustness of Allowance Banking in the U.S. Acid Rain Program”, *The Energy Journal* **28**, 205 –233.

Engle R.F., and C.W.J. Granger (1987) “Cointegration and Error Correction: Representation, Estimation and Testing”, *Econometrica* **55**, 251–276.

Fama, E. F. (1970) “Efficient Capital Markets: a Review of Theory and Empirical Work”, *Journal of Finance* **25**, 383– 417.

Helfand G.E., Moore M.R., and Liu Y. (2007) “The Intertemporal Market for Sulfur Dioxide Allowances”, School of Natural Resources and Environment, University of Michigan, November.

Joskow P. L., Schmalensee R., and E.M. Bailey (1998) “The Market for Sulfur Dioxide Emissions”, *American Economic Review* **88**, 669–685.

Keohane, N. (2006) “Cost Savings from Allowance Trading in the 1990 Clean Air Act: Estimates from a Choice-Based Model”, in *Moving to Markets in Environmental Regulations: Lessons from Thirty Years of Experience* by J. Freeman and C. D. Kolstad, Eds., Oxford University Press: New York.

Keynes, J.M. (1930) “*A Treatise on Money*”, vols. I-II, Harcourt, Brace and Co.: New York; also in *The Collected Writings of John Maynard Keynes*, vols. 5-6, by D.E. Moggridge, Eds., Macmillan: London, 1971.

Kosobud R.F., Stokes H.H., Tallarico C.R., and B.L. Scott (2005)”Valuing Tradable Private Rights to Pollute the Public’s Air”, *Review of Accounting and Finance* **4**, 50–71.

Kwiatkowski D., Phillips P.C.B., Schmidt P., and Y. Shin (1992) “Testing the Null Hypothesis of Stationary against the Alternative of a Unit Root. How Sure are we that Economic Time series Have a Unit Root?”, *Journal of Econometrics*, **54**, 159–178.

MacKinnon J.G., 1991, “Critical Values for Cointegration Tests”, in *Long-run Economic Relationships* by R.F. Engle and C.W.J. Granger, Eds., Oxford University Press, 267–276.

Perron, P. (1989) “The Great Crash, the Oil Price Shock and the Unit Root Hypothesis”, *Econometrica* **57**, 1361–1401.

Phillips P.C.B., and P. Perron (1988) “Testing for a Unit Root in a Time Series Regression”, *Biometrica* **75**, 335–346.

Schmalensee R, Joskow PL, Ellerman AD, Montero JP, and E.M Bailey (1998) “An Interim Evaluation of Sulfur Dioxide Emission Trading”, *Journal of Economic Perspectives* **12**, 53–68.

Swinton J. R. (2004) “Phase I Completed: An Empirical Assessment of the 1990 CAAA”, *Environmental and Resource Economics* **27**, 227–246.

Swinton J. R. (2002) “The Potential for Cost Savings in the Sulfur Dioxide Allowance Market: Empirical Evidence from Florida”, *Land Economics* **78**, 390–404.

Zivot, E., and D. Andrews (1992) “Further Evidence of the Great Crash, the Oil-Price Shock and the Unit Root Hypothesis, *Journal of Business and Economic Statistics* **10**, 251–270.