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Understanding volatility dynamics in the EU-ETS market: lessons *from* the future

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In this paper we study the short term price behavior of December 2008 future prices for EU emission allowances. We model returns and volatility dynamics of this price showing that a standard ARMA-GARCH framework is not adequate and that the gaussianity assumption is rejected due to the occurrence of a number of level and volatility outliers. To improve the fitness of the model, we combine the underlying price process with an additive stochastic jump process. The resulting distribution, a mixture of Gaussians, allows for endogenously determined jumps in the process governing the returns, while the mixing law determines the jump probability. The performance of the model is improved by introducing a time varying jump probability explained by two variables. The first one is the daily relative change in the volume of transactions and suggests that sharp increases in volume increase volatility even in the absence of changes in what recent literature considers as market fundamentals. The second one accounts for changes in the jump probability associated to the European Commission's announcements regarding the NAPs for Phase II. We find that announcements concerning the NAPs induce jumps in the process and tend to increase volatility. This result suggests authorities should advocate to increase stability in the regulatory environment which is crucial to allow traders to realize efficient trading strategies and informed investment decisions regarding pollution reduction.

Keywords: EUA market, EU-ETS, carbon emission trading, Garch model, normal mixture
JEL Classification: C16, C32, C51, C53, Q52, Q53

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1. INTRODUCTION

The countries that have ratified the Kyoto Protocol have committed to reduce greenhouse gas emissions, during the period 2008-2012, to an extent established by international negotiations. To enforce the CO_2 reduction, the European Union has chosen a "cap and trade" regulation. This approach establishes a global cap on emissions, which coincides with the reduction target, and creates a market for pollution permits, called EU Emission Trading Scheme (EU-ETS). In this market, agents can exchange their surplus or deficit of allowances (EUAs). The EU-ETS is implemented in two phases: Phase I from 2005 to 2007 and Phase II from 2008 to 2012. Phase II is the period of the actual implementation of the Protocol objectives.

Firm-level trading[1] started in January 2005. At the beginning of 2006, the volume of transactions had already increased by a factor of 10 (Ellerman and Joskow, 2008). The development of the EU-ETS market has been also due to the increasing market participation of intermediaries, i.e. risk managers, brokers and traders, which may trade on behalf of their clients or hold their own stock of EUAs. The market gains both in complexity and in flexibility as intermediaries introduce an increasing range of new instruments such as futures, forward contracts and other derivatives. In this regard, many recognize that the creation of the EU-ETS has been a success while others remain skeptical. In particular, the rules behind the price formation mechanism and the price dynamics are still unclear. While some authors support the argument that EUA price responds to market fundamentals which affect the production of CO_2 and thus demand and supply of EUAs, e.g. energy prices, extreme weather conditions and economic growth (see Bunn and Fezzi, 2007, Rickels et al., 2007, Mansanet-Bataller et al., 2007, and Alberola et al., 2008), others find no such evidence and favour a pure time-series approach (see Milunovich and Joyeux (2007), Paolella and Taschini, 2008, Benz and Trück, 2008, Chesney and Taschini, 2008, Seifert et al., 2008). In particular, Seifert et al. (2008) analyze the dynamics of EUA spot price using a stochastic equilibrium model finding that an adequate EUA price process does not necessarily exhibit a seasonal component, that it should possess the martingale property and a time dependent volatility dynamic. Paolella and Taschini (2006) discuss forecasting methods based on the analysis of supply and demand fundamentals and on the spot-future parity. They conclude that both approaches yield to implausible conclusions due to the complexity of the market and to the particular behavior of the emission allowances. They advocate the use of statistical models relying exclusively on historical price information and suggest to analyze the riskiness of the emission allowances by addressing the unconditional tail behavior and the conditional heteroskedasticity for the dynamics of the returns. Milunovich and Joyeux (2007) examine the issues of market efficiency and price discovery in the EU-ETS futures market. They find that futures with maturities 2006 and 2007 exhibit a stable long-run relationship with Phase I spot price, while 2008 futures do not form such a relationship. They attribute their finding to the unavailability of a relevant spot price for Phase II. Indeed, due to the non-bankability of the EUAs between Phase I and Phase II, futures with maturity 2008 are not expected to be cointegrated with Phase I spot price but they rather act as a vehicle of price discovery for Phase II spot price.

An adequate assessment of short term price and volatility dynamics in the EU-

[1]The ETS is different from the International Emissions Trading (IET), where countries can trade their surplus or deficit of allowances with respect to their national allocations. The IET became operative in the begining of 2008.

ETS is crucial since an accurate measurement and forecasting of market risk is a key factor for portfolio management and hedging, to realize efficient trading strategies and to take informed investment decisions in a market that is steadily gaining in complexity.

We analyze the short term price and volatility dynamics of December 2008 EUA future prices as they represent an adequate proxy for Phase II spot price. We model the conditional mean and variance of these price returns within an ARMA-GARCH framework. The standard approach based on the Gaussianity assumption is rejected due to the presence of a number of level and volatility outliers. Consequently, we rely on a Bernoulli mixture of Gaussian distributions (BMN) to allow for endogenously determined additive jumps in the price process. The individual distributions in the mixture can be interpreted as different regimes while the mixing law gives the probability of each regime (Alexander, 2004 and Alexander and Lazar, 2006). We find that a two-regime model based on a BMN proves adequate to fit the data.

Paolella and Taschini (2006) have adopted a similar modelling strategy. They propose a 3-component mixture which identifies two different GARCH-type volatility dynamics plus a constant variance component. Although their model does not account for an additive jump component, they provide solid arguments to support the use of a mixture of distributions such as the method's extreme flexibility, the fact that it induces time varying skewness and kurtosis (see also Hansen, 1994, Harvey and Siddique, 1999, Rokinger and Jondeau, 2002 and Brännäs and Nordman, 2003) and the accuracy of the out of sample VAR forecasts. For an extensive overview of the properties of the mixture of distributions see Alexander and Lazar (2006) and Haas et al. (2004), among others.

An alternative approach, based on a two regime Markov switching model, has been recently proposed by Benz and Trück (2009). They argue that the occurrence of spikes in EUA prices and volatility could be caused by changes in policy and the regulatory framework such as announcements concerning the National Allocation Plans (NAPs) or fluctuations in production levels due to unexpected changes in market fundamentals (such as fuel prices and weather conditions). However, this hypothesis cannot be directly tested in their case since they assume constant the probability that governs the switch between regimes yielding few economic insights.

The procedure based on the use of a GARCH-type model with mixed innovations to fit an underlying price process combined with an additive jump component has been proposed by Vlaar and Palm (1993), Vlaar (1994) and Beine and Laurent (2003). Their approach is appealing because it provides useful insights on the occurrence of the jumps and their economic interpretation. In this paper, the determinants and the occurrence of the jumps are further investigated by letting the probability associated to the jump component vary over time and depend on exogenous variables. In particular, we explicitly account for two drivers of the shifts between regimes: the daily relative change in the volume of transactions and the change in the regulatory environment induced by the European Commission's disclosure of Phase II NAPs for each member State.

Our result regarding the destabilizing effect of large incoming volumes which translates into large negative returns and sudden volatility movements is in line with Gabaix et al. (2006). They show that significant spikes in returns can be motivated by trades placed by large investors in relatively illiquid markets, even in the absence of important news about fundamentals. Indeed, the EU-ETS, due to its novelty and uniqueness, represents a market where long-term future contracts are traded in a relatively illiquid market (Milunovich and Joyeux, 2007) and where the concentration of the market among few leading players, the relatively low number of market transactions, the lack of transparency and therefore the discontinuous flow of information available, specially in its initial phase, play a dominant role (Benz and Hengelbrock, 2008).

The consequences of announcements concerning the NAPs on EUA short term price behavior is comparable to the effect of Central Banks intervention on the exchange rate market assessed by Beine and Laurent (2003) in the sense that announcements concerning the NAPs induce jumps and tend to increase volatility. The instability following the announcements for Phase II NAPs can be explained by the unexpected relative scarcity of EUA for the second phase. The adopted NAPs revealed to be sensibly more restrictive than the target proposed by each member State. In fact, the emission cap approved by the European Commission for Phase II (i.e. the sum of the national allocations) was less then 90% of the total emission target proposed by the member States[2].

The reminder of the paper is organized as follows. Section 2 briefly discusses the main features of the EU-ETS market and describes the data used for the empirical analysis. Section 3 presents the standard ARMA-GARCH model and a set of tests statistics used for it's validation. Section 4 introduces a procedure for outlier detection. Section 5 presents the Bernoulli mixture of normals and its extension that allows for a time varying jump probability. Section 6 concludes.

2. STYLIZED FACTS AND DATA DESCRIPTION

The EU-ETS covers up to 46% of European CO_2 emissions coming from 11.400 industrial installations over the European Union (Mission climat, 2006). These installations receive periodically a free amount of pollution allowances that can be traded in any of the Exchanges (e.g. Powernext, European Climate Exchange and Nordpool), over the counter (OTC) or trough a private transaction. Unused allowances with vintage belonging to the period 2005-2007, i.e. corresponding to Phase I, expire at the end of this phase and cannot be banked and used during Phase II. As a consequence, the EUA spot price at the end of Phase I could either converge to zero, as it was the case as the market was long, or reach the upper bound of 40 euros which represents the penalty, per ton of CO_2, established by the EU Commission to be paid by those installations which fail to cover their emission with allowances.

As shown in Figure 1, the starting spot price was about 17 euros and rose up to 30 euros in June 2005. The disclosure of 2005 verified emissions in April 2006 proved the NAPs for Phase I where too generous (or that firms had engaged in more pollution abatement than expected[3]). After the announcement, the market appeared to be about 4% long, provoking the EUA spot price to fall from 29.5 euros to less than 12 in a few days. Also the December 2008 future prices fell, stopping at 18.25 euros. The first disconnection between the spot and future price occurs. From April to September 2006, the spot price, as well as the December 2008 future price, remained stable between 15 and 20 euros. From October 2006, while EUA spot price started converging towards 0, December 2008 futures rose and settled well above 20 euros. Milunovich and Joyeux (2007) show that while futures with maturities 2006 and 2007 exhibit a stable long run relationship with

[2] The yearly cap during Phase I was 2.298 billion tons of CO2 while in Phase II it has been set to 2.081 billion tons.

[3] For a further discussion see Ellerman and Buchner (2007).

FIG. 1 EUA spot and Dec. 2008 future prices

Phase I spot price, the 2008 futures do not form such relationship but they rather act as a vehicle of price discovery for Phase II spot price due to the non-bankability of Phase I permits.

In this paper we consider the returns on daily December 2008 future prices (in euros per ton of CO_2) traded in the European Climate Exchange (ECE) between the first quotation in April 22, 2005 and December 31, 2007 (676 daily observations). The ECE represents the most liquid future market in the EU-ETS, with 75.6% of the futures exchange volume in 2006. We compute returns (r_{EUA}) as the first difference of the natural logarithm of the price series.

Following previous literature (Mansanet-Bataller et al., 2007, and Alberola et al., 2008) we consider, as possible market fundamentals, several fuel prices and weather indexes. In particular we consider: (i) daily Month Ahead Future Natural Gas price (in euros per therm) traded on the Zeebrugge Hub; (ii) daily coal Future Month Ahead price CIF ARA (in euros per ton); (iii) daily Future Month Ahead Base price for electricity (in euros per MWh) traded in the Powernext; and (iv) daily Brent Crude Future Month Ahead price negotiated on the Intercontinental Futures Exchange, expressed in euros per barrel using the European Central Bank exchange rate. Returns are denoted r_{gas}, r_{coal}, r_{elec}, r_{oil} respectively.

Further, we consider a weather index based on a weighted average of deviations from historical temperatures as in Alberola et al. (2008). Based on the index, we create two dummies: D_{tmp_lo} and D_{tmp_hi} that account for temperatures in the 5% lowest and highest percentile, respectively. Summary statistics are reported in Table 1.

Table 1: Summary statistics

Variable	min	mean	max	std dev	sk	ku
Estimation sample: 22/04/05-31/12/07 (676 observations)						
r_{EUA}	-28.82	0.041	18.65	3.046	-1.443	18.77
r_{vol}	-3.683	0.002	3.321	0.660	0.096	8.838
r_{gas}	-13.63	0.071	30.58	3.398	2.738	20.09
r_{coal}	-8.781	0.080	11.72	1.436	1.381	22.48
r_{elec}	-18.89	0.124	30.26	4.392	1.713	16.07
r_{oil}	-5.073	0.061	6.506	1.733	0.111	3.185
D_{tmp_lo}	0.000	0.050	1.000	0.218	4.115	17.93
D_{tmp_hi}	0.000	0.050	1.000	0.218	4.115	17.93

Table 1 shows that EUA returns (r_{EUA}) exhibit negative skewness and a large excess kurtosis. All the other variables are characterized by excess kurtosis and positive skewness. The only exception is the return on oil, for which both skewness and kurtosis coefficients are very close to the ones implied by the Gaussian distribution.

The presence of excess kurtosis in r_{EUA} means that extreme values for the returns (either positive or negative) occur with a frequency which is higher than the one implied by the Gaussian distribution. Indeed, the occurrence of outliers is primarily responsible for the rejection of the Gaussianity assumption for the EUA future returns.

As a determinant for the occurrence of jumps, we explicitly consider the relative change in the daily volume of future contracts traded in ECE (r_{vol}) and a binary variable (D_{NAP}) that accounts for the European Commission's announcement of member State's NAPs for Phase II (details are reported in Table 2). The presence of outliers can also be motivated by other specific events such as changes in abatement decisions due to a switch in the relative costs of coal and cleaner fossil fuels such as natural gas. Even if it may be difficult to identify specific dates for such changes, these are accounted for in our analysis since their impact is included in the constant of the jump probability.

Table 2: Announcements of NAPs for Phase II

Date ($D_{NAP} = 1$)	State	NAP Phase I	2005 emissions	NAP Phase II (% of proposed)
09/01/06	Publication of guidelines for NAP approval			
29/11/06	Germany	499	474	453.1 (94.0%)
	Greece	74.4	71.3	69.1 (91.5%)
	Malta	2.9	1.98	2.1 (71.0%)
	UK	245.3	242.4	246.2 (100%)
16/01/07	Belgium	62.1	55.58	58.5 (92.4%)
	Netherlands	95.3	80.35	85.8 (94.9%)
05/02/07	Slovenia	8.8	8.7	8.3 (100%)
26/02/07	Spain	174.4	182.9	152.3 (99.7%)
26/03/07	Czech Rep.	97.6	82.5	86.8 (85.2%)
	France	156.5	131.3	132.8 (100%)
	Poland	239.1	203.1	208.5 (73.3%)
02/04/07	Austria	33	33.4	30.7 (93.6%)
16/04/07	Hungary	31.3	26	26.9 (87.6%)
04/05/07	Estonia	19	12.62	12.72 (52.2%)
15/05/07	Italy	223.1	225.5	195.8 (93.7%)
04/06/07	Finland	45.5	33.1	37.6 (94.8%)
13/07/07	Ireland	22.3	22.4	22.3 (98.6%)
	Latvia	4.6	2.9	3.43 (44.5%)
	Lithuania	12.3	6.6	8.8 (53.0%)
	Luxembourg	3.4	2.6	2.5 (63.0%)
	Sweden	22.9	19.3	22.8 (90.5%)
18/07/07	Cyprus	5.7	5.1	5.48 (77.0%)
31/08/07	Denmark	33.5	26.5	24.5 (100%)
22/10/07	Portugal	38.9	36.4	34.8 (96.9%)
26/10/07	Bulgaria	42.3	40.6	42.3 (62.6%)
	Romania	74.8	70.8	75.9 (79.3%)
07/12/07	Slovakia	30.5	25.2	32.6 (78.9%)
Total		2298.5	2122.2	2082.7 (89.6%)

3. BENCHMARK APPROACH

The starting point for the investigation of the EUA price determinants of returns and volatility dynamics is the ARMA-GARCH framework. This model, widely used in the literature, allows for the presence of exogenous regressors and specifically account for conditional heteroskedasticity and serial dependence in the returns. To assess the relevance of the ARMA-GARCH setting and discriminate between competing specifications, we suggest a set of four diagnostic tests as detailed in the remainder of this Section.

3.1. ARMA-GARCH model

Consider the stochastic process $r_t = p_t - p_{t-1}$, where p_t is the natural logarithm of the EUA price. The conditional mean of the process is expressed as

$$\Phi(L)r_t = \beta X_{t-1} + \Psi(L)\varepsilon_t \tag{1}$$

$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, \sigma_t^2) \tag{2}$$

where $\Phi(L) = 1 - \sum_{i=1}^{m} \varphi_i L^i$ and $\Psi(L) = 1 + \sum_{i=1}^{n} \psi_i L^i$ are the usual AR and MA polynomials of order m an n respectively. Ω_t is the information set at time t and L is the lag operator such that $L^k x_t = x_{t-k}$ $(k > 0)$ and X_{t-1} is a matrix of lagged regressors (up to a constant). The normality assumption is justified by the fact that the Gaussian Quasi-Maximum Likelihood (QML) estimation delivers consistent estimates even when normality assumption is rejected, provided that mean and variance are correctly specified (see Weiss, 1986 and Bollerslev and Wooldridge, 1992 among others).

For the conditional variance we consider the $GARCH(p, q)$ specification (Bollerslev, 1986), that is

$$B(L)\sigma_t^2 = c + A(L)\varepsilon_t^2 \tag{3}$$

with characteristic polynomials $B(L) = 1 - \sum_{i=1}^{q} b_i L^i$ and $A(L) = 1 + \sum_{i=1}^{p} a_i L^i$.

The model is estimated by quasi maximum likelihood. The sample log likelihood is given by

$$LLF = -\frac{T}{2}\log(2\pi) - \frac{1}{2}\sum_{T}\log(\sigma_t^2) - \frac{1}{2}\sum_{T}\frac{\varepsilon_t^2}{\sigma_t^2} \tag{4}$$

where T is the sample size and which is maximized numerically for $(\beta, \varphi_k, \psi_l, c, a_i, b_j)$, $k = 1, ...m; l = 1, ..., n; i = 1, ..., p; j = 1, ..., q$.

3.2. Diagnostic tests

In order to correctly estimate the risk borne by an agent trading on EUA market, the choice of an adequate distribution is crucial. Therefore, to discriminate between model specifications and verify distributional assumptions, we devote particular attention to diagnostic tests.

Following Vlaar (1994) and Beine and Laurent (2003), we focus on the following set of statistics. First, we consider two test for the estimated skewness (b_3) and kurtosis (b_4) coefficients of standardized residuals (i.e. $\eta_t = \varepsilon_t/\sigma_t$) respectively. Second, we check the hypothesis of independent and identically distributed (*iid*) residuals based on the statistic proposed by Brock et al. (1996) (BDS test). The *iid* hypothesis will become crucial to interpret the results of the Pearson goodness-of-fit test. Indeed, the rejection of the *iid* hypothesis would make unclear the interpretation of the results.

The BDS test for independence is based on the estimation of correlation integrals at various dimensions (m) and for a given dimensional distance (ϵ). The idea is that if a random variable X is *iid* the joint probability of two observations as well as their two predecessors being within a distance ϵ is equal to the square of the probability of any two observation being within the distance ϵ. Formally this can be stated as

$$\begin{aligned} P_m &\equiv P(|X_i - X_j| < \epsilon, ..., P(|X_{i-m+1} - X_{j-m+1}| < \epsilon) \\ P_1 &\equiv P(|X_i - X_j| < \epsilon) \\ P_m &= P_1^m \Longleftrightarrow X \text{ is } i.i.d. \end{aligned} \tag{5}$$

The BDS statistic for embedding dimension m (BDS(m)) on a sample of size n, is given by

$$BDS(m) = \sqrt{n-m-1}\frac{c_{m,n} - c_{1,n-m-1n}^{m}}{\sigma_{m,n}}, \tag{6}$$

where $c_{m,n}$ is the correlation integral for dimension m and $\sigma_{m,n}$ is the variance of $c_{m,n} - c^m_{1,n-m-1}$.

Brock et al. (1996) show that this statistic asymptotically follows a standard normal distribution for any m and ϵ. About the choice of the embedding dimension m and the distance ϵ, we set m equal to 6 and, as suggested Kanzler (1999) and ϵ such that the first correlation integral is equal to 0.7.

The Pearson goodness-of-fit test compares the empirical distribution of η_t to the theoretical one. For a given number of cells g, the statistics is given by

$$P(g) = \sum_g \frac{(n_i - En_i)^2}{En_i}, \tag{7}$$

where n_i is the number of observation in cell i and En_i is the expected number of observations. For *iid* observation, under the null of a correct distribution the statistic is distributed as chi-square with $g-1$ degrees of freedom, i.e. $\chi^2(g-1)$. Accordingly to Palm and Vlaar (1997) the number of cells, to be chosen in proportion to the sample size, is set to 30.

Preliminary results (Model 1 in Table 4) based on the standard setting introduced so far suggest, according to the Pearson goodness-of-fit test, the rejection of the Gaussian distribution at standard significance levels. In fact the standardized residuals show excess skewness (b_3) and kurtosis (b_4) with respect to the normal distribution.

4. OUTLIERS DETECTION

The rejection of the benchmark model, i.e. the ARMA-GARCH model with normally distributed errors, is primarily due to the presence of a large number of extreme observations (outliers). Standardized residuals of Model 1 (Table 4) exhibit a kurtosis three times larger than the one of standard Gaussian distribution. The excess kurtosis, along with the skewness, may be related to a number of events that took place during the period of study and that the Gaussian density cannot take into account.

An interesting approach to identify level and variance shifts is the outliers detection procedure by Doornik and Ooms (2005). This procedure makes clear distinction between Additive Level Outliers (ALO), that only affect the level but leave the variance unaffected, and Additive Variance Outliers (AVO), which in turn also affect the conditional variance. The technique is based on a sequence of likelihood-ratio tests under the null of no outliers (at some unknown date), and is carried out in an iterative way which requires five sequential steps (see Doornik and Ooms, 2005 for details). This approach is appealing because it allows to identify the date at which the extreme observation appears, to discriminate between level and volatility outliers, to determine the outlier size and finally to accordingly adjust the data.

Table 3 reports the detected outliers for the ARMA(0,0)-GARCH(1,1) model for the EUA returns at a significance level $\alpha = 0.25$ which, given the approximation of the asymptotic distribution of the test provided by Doornik and Ooms, corresponds to a quintile of 13.95.

Table 3: Outliers

Date	Size	Stat	Type
16/06/05	3.997	14.442	ALO
17/06/05	6.240	16.939	ALO
12/07/05	-5.634	22.995	ALO
13/07/05	-15.284	25.697	AVO
19/07/05	-15.304	15.127	ALO
22/02/06	-8.804	27.442	ALO
19/04/06	-4.194	16.912	ALO
24/04/06	-4.655	19.129	ALO
25/04/06	-28.975	75.841	AVO
12/05/06	18.768	13.993	ALO
11/10/06	-4.329	18.383	ALO
16/02/07	-7.553	15.856	AVO

Interestingly, within this set of outliers we can clearly identify two clusters: June - July 2005 and end of April 2006. The presence of outliers in the period June - July 2005 could reflect a period of high variability of the natural gas future price. In fact, the two series were highly correlated between the beginning of Phase I and July 2005, while, from August of the same year, natural gas and EUA prices showed independent paths. In turn, it is well established that the large negative outliers in the end of April 2006 coincide with the release of the carbon emission reports from Belgium, Czech Republic, Estonia, France, Netherlands and Spain which showed that the market was longer than expected.

Behind the rejection of the Gaussian distribution is the relatively large number of outliers. An appealing solution to take into account extreme observations, is the introduction of endogenous level and variance shifts using a mixture of distributions. In the following section we introduce the Bernoulli Mixture of Normal distributions (BMN). Other than extreme observation, the BMN has the advantage to account for excess skewness and kurtosis (Vlaar, 1994 and Alexander and Lazar, 2006).

5. BERNOULLI MIXTURE OF NORMALS

The high number of extreme returns relative to the sample size, requires the introduction of some alternative approach which allows to model level and variance shifts. Further, EUA future returns are not symmetric and therefore, a symmetric distribution as the Gaussian is unlikely to give appropriate results. We combine Gaussian distributions and an additive stochastic jump process. The resulting mixture of distribution has the advantage to account for excess skewness and kurtosis (Vlaar, 1994, Alexander and Lazar, 2006). Several parametrization have been suggested for the mixing law (see, for instance, Vlaar, 1994). We focus on the Bernoulli mixture of Gaussians. Its basic assumption implies that the mixing law for the returns densities is Bernoulli. The advantage of this parametrization is its intuitive interpretation: the individual distributions in the mixture represent different regimes while the mixing law gives the probabilities of each regime (Alexander, 2004 and Alexander and Lazar, 2006).

Given the stochastic process $r_t = p_t - p_{t-1}$, with conditional mean $\mu_t = E(r_t \mid \Omega_{t-1})$ and time varying conditional variance $\sigma_t^2 = E(r_t^2 \mid \Omega_{t-1})$, the mixture process can be defined as

$$\begin{array}{llll} r_t = \mu_t + \sigma_t z_t & & & \text{with probability } 1-\lambda \\ r_t = \underbrace{\mu_t + \sigma_t z_t}_{\textit{continuous comp.}} & + & \underbrace{\tau + \delta z_t^*}_{\textit{additive jump comp.}} & \text{with probability } \lambda \end{array} \tag{8}$$

where z_t and z_t^* are *iid* $N(0,1)$ and λ is the probability of having a level and variance shift and represents the parameter of the mixing law. Finally, τ and δ^2 are the mean and variance of the jump distribution respectively. It is worth noting that $1/\lambda$ represents the average interval between two consecutive jumps.

The model can be rewritten as:

$$\begin{aligned} r_t \quad &= \quad \mu_t + \lambda\tau + \varepsilon_t; \\ \varepsilon_t \quad &| \quad \Omega_{t-1} \sim (1-\lambda)N(-\lambda\tau, \sigma_t^2) + \lambda N(\tau - \lambda\tau, \sigma_t^2 + \delta^2). \end{aligned} \tag{9}$$

To ensure the condition $0 < \lambda < 1\ \forall t$, we use the following transformation

$$\lambda = 1 - (1 + \exp(\gamma_0))^{-1}. \tag{10}$$

Since a linear combination of normally distributed random variables is also normal, this combination results in a discrete mixture of normals.

Given the ARMA-GARCH setting detailed in Section 4, the conditional mean and variance of the process r_t can be expressed as

$$\begin{aligned} r_t \quad &= \quad \beta X_{t-1} + \sum_{i=1}^{m} \varphi_i r_{t-i} + \sum_{i=1}^{n} \psi_i \varepsilon_{t-i} + \lambda\tau + \varepsilon_t; \\ \sigma_t^2 \quad &= \quad c + \sum_{i=1}^{p} a_i \varepsilon_{t-i}^2 + \sum_{j=1}^{q} b_j \sigma_{t-j}^2; \\ \varepsilon_t \quad &\sim \quad N(0, \sigma_t^2). \end{aligned} \tag{11}$$

We consider lagged values of the exogenous variables (X_{t-1}) so that the pair (r_t, σ_t^2) are measurable with respect to the information available at time $t-1$, ensuring that the model is completely forecastable.

The log likelihood of this distribution is given by

$$\begin{aligned} LLF \quad = \quad & -\frac{T}{2}\ln(2\pi) + \sum_{t=1}^{T} \log\left[\frac{1-\lambda}{\sqrt{\sigma_t^2}} \exp\left(-\frac{(r_t - \mu_t)^2}{2\sigma_t^2}\right)\right. \\ & \left. + \frac{\lambda}{\sqrt{\sigma_t^2 + \delta^2}} \exp\left(-\frac{(r_t - \mu_t - \tau)^2}{2(\sigma_t^2 + \delta^2)}\right)\right]. \end{aligned} \tag{12}$$

As pointed out by Vlaar (1994), in such framework the Pearson goodness-of fit test cannot be applied on standardized residuals because the *iid* assumption is no longer satisfied. Palm and Vlaar (1997) redefine the sorting mechanism of the residuals and suggest the use of normalized residuals defined as

$$z_t = F^{-1}\left[(1-\lambda)F(\frac{r_t - \mu_t}{\sigma_t}) + \lambda F(\frac{r_t - \mu_t - \tau}{\sigma_t + \delta})\right], \tag{13}$$

where $F^{-1}()$ and $F()$ are respectively the quintile function and the cumulative distribution function of the standard normal density.

Table 4 reports estimation results for different specifications of (11). Following Beine and Laurent (2003), the ARMA and GARCH orders are selected by relying on the Schwarz Bayesian Information Criterion which is known to lead to a parsimonious specification. Following this criterion the specification selected is ARMA(1,0)-GARCH(1,1). Results are not reported to save space, but are available upon request.

Table 4: Benchmark model and BMN with constant jump probability

Param.	Model 1	Model 2	Model 3	Model 4
γ_0	–	**−3.272**	**−2.782**	**−1.921**
		(0.453)	(0.379)	(0.310)
τ	–	**−3.330**	**−2.183**	**−1.131**
		(2.680)	(1.852)	(0.870)
δ^2	–	**62.063**	**44.260**	**23.985**
		(19.397)	(10.383)	(5.484)
β_0	**0.183**	**0.271**	**0.302**	**0.356**
	(0.116)	(0.155)	(0.166)	(0.178)
β_1	**0.119**	**0.101**	**0.111**	**0.097**
	(0.050)	(0.044)	(0.042)	(0.041)
β_{gas}	−0.001	−0.006	–	–
	(0.027)	(0.024)		
β_{coal}	−0.061	−0.013	–	–
	(0.069)	(0.080)		
β_{elec}	0.025	0.012	–	–
	(0.023)	(0.021)		
β_{oil}	**0.066**	**0.087**		–
	(0.041)	(0.049)		
β_{tmp_lo}	0.249	0.285	–	–
	(0.443)	(0.369)		
β_{tmp_hi}	0.149	−0.074	–	–
	(0.436)	(0.378)		
$\beta_{04/2006}$	–	–	–	**−6.093**
				(1.662)
c	**0.827**	**0.516**	**0.378**	**0.231**
	(0.176)	(0.161)	(0.130)	(0.101)
a	**0.342**	**0.120**	**0.120**	**0.108**
	(0.036)	(0.029)	(0.029)	(0.027)
b	**0.630**	**0.760**	**0.774**	**0.786**
	(0.042)	(0.051)	(0.046)	(0.045)
λ	–	0.036	0.058	0.127
Freq.	–	27.36	17.15	7.82
# Jumps	–	24	39	86
LLF	−1625.0	−1586.6	−1588.4	−1582.1

	Stat	p-val	Stat	p-val	Stat	p-val	Stat	p-val
b_3	−0.881	0.00	−0.053	0.56	−0.091	0.33	−0.140	0.13
b_4	6.441	0.00	0.465	0.01	0.527	0.00	0.598	0.00
$BDS(6)$	−1.717	0.08	0.547	0.58	0.376	0.71	−0.021	0.98
$P(30)$	80.12	0.00	32.90	0.28	38.32	0.11	33.81	0.29

Notes: Parameters standard errors in parentheses. Parameters significant at 10% in bold.

As it appears from Table 4, we do not find evidence to support the assumption that EUA returns dynamics can be explained by fundamentals. Model 1 represents a standard GARCH model with Gaussian innovation which is obtained by imposing the parameter restriction $\lambda = 0$ to (11). Interestingly, none of the explanatory variables is significant at standard levels. The only exception is represented by the oil price which in Model 1 is weakly significant. Benz and Trück (2008) suggest that, though the EUA prices may show phases of specific price behavior due to fluctuations in production levels induced by shocks in fuel prices and extreme weather conditions, these sources of uncertainty have a rather short-term impact and thus they could induce price and volatility jumps rather than exhibit a strong relation with the underlying return process. In fact we will further investigate this in what follows.

In Model 2, we consider the same specification used in Model 1 but in the BMN framework. Although the BMN is not rejected, all explanatory variables are again insignificant.

For ease of comparison, Model 3 replicates the same specification of Model 2, but excluding all explanatory variables. The advantage of using the BMN is striking. All parameters of the BMN are significant with the exception of the jump size (τ). As we have seen from Table 3, level outliers, though characterized by a large size in absolute value, show opposite signs. Since τ represents the average size of level outliers, the result suggests that a single parameter might not be sufficient to capture the sign of the level outliers because they tend to compensate, in average. The skewness parameter (b_3) associated to this specification is close to zero and insignificant, though the normalized residuals still exhibit some excess kurtosis (b_4). It is worth noting that even if the null of no excess kurtosis is rejected, the use of the BMN induces an important decrease in the excess kurtosis of the residuals compared with the results of the standard GARCH model with normally distributed errors. The coefficient b_4 reduces from 6.44 (Model 1) to 0.53 (Model 3). The Pearson test does not reject the BMN in Model 3 at 5% nominal level.

As noted from Table 3, and according to Alberola et al. (2008) and Milunovich and Joyeux (2007), a specific cluster of outliers coincides with the week before the compliance break on April 25, 2006. In Model 4, we account for this information by adding a dummy variable which refers to the period between April 19 and April 25, 2006 ($D_{04/2006}$). By explicitly modelling the compliance break period, characterized by large negative returns, we improve the accuracy of the model. We find a larger number of jumps (86) with respect to Model 3 but characterized by a smaller intensity (i.e. both parameters of the jump distribution τ and δ^2 become smaller). More precisely, the jump probability of Model 4 is twice the probability of Model 3, but the jump intensity reduces to half. The Pearson test indicates that the BMN is supported for Model 4 at 5% significance level.

Although the BMN in (11) outperforms the standard GARCH-normal approach, assuming a constant jump probability specification yields few economic insights. In fact, according to recent literature we could argue that these jumps may be due to changes in production due to shocks in fuel prices or by changes in the regulatory environment but we could test explicitly these events' impact in the jump probability. Furthermore, despite the use of the BMN, normalized residuals still exhibit excess kurtosis. In fact, in all specifications the statistics for excess kurtosis are found to be significant at 5%. In this regard, Beine and Laurent (2003) show that a BMN may fail in specifically account for excess kurtosis and they call for either a better specification of the conditional mean and variance, or for further flexibility in the distribution.

We will pursue the latter option through the introduction of a time varying jump probability as a function of a constant and a set of exogenous variables (x_{it-1}) related to the jump occurrence. This allows us to specifically test which are the determinants of outliers. The specification for λ_t follows a logistic functional form:

$$\lambda_t = 1 - (1 + \exp(\gamma_0 + \Sigma_i \gamma_i x_{it-1}))^{-1} \tag{14}$$

which ensures $0 < \lambda_t < 1 \; \forall t$.

Table 5: Dynamic jump probability

Param.	Model 5		Model 6		Model 7	
γ_0		**−2.871**		**−2.603**		**−2.394**
		(0.398)		(0.369)		(0.361)
γ_{vol}		**0.795**		**0.846**		**0.802**
		(0.391)		(0.399)		(0.394)
γ_{NAPs}		–		**1.770**		**1.781**
		–		(0.982)		(0.968)
τ		−1.561		−0.856		−0.730
		(1.502)		(0.950)		(0.813)
δ^2		**43.07**		**34.14**		**29.12**
		(9.083)		(5.655)		(6.024)
β_0		**0.261**		**0.235**		**0.251**
		(0.144)		(0.132)		(0.134)
β_1		**0.109**		**0.107**		**0.099**
		(0.042)		(0.420)		(0.041)
$\beta_{04/2006}$		–		–		**−4.125**
						(1.316)
c		**0.367**		**0.294**		**0.268**
		(0.126)		(0.107)		(0.104)
a		**0.122**		**0.127**		**0.120**
		(0.030)		(0.030)		(0.029)
b		**0.770**		**0.771**		**0.777**
		(0.046)		(0.045)		(0.045)
LLF		−1585.8		−1584.8		−1580.9
	Stat	p-val	Stat	p-val	Stat	p-val
b_3	−0.109	0.24	−0.173	0.06	−0.157	0.10
b_4	0.495	0.01	0.650	0.00	0.571	0.00
$BDS(6)$	0.319	0.75	−0.113	0.91	−0.250	0.80
$P(30)$	31.84	0.33	33.17	0.27	31.04	0.36

Notes: See Table 4.

We identify two potential candidates as drivers of the shifts between regimes: the daily relative change in the volume of transactions and changes in the regulatory environment due to the European Commission's disclosure of Phase II NAPs for each member State. Table 5 reports results for the specification (11)-(14). In Model 5, we let the jump probability depend on past realizations of the log-differential of the daily trading volume of December 2008 future contracts. The trading volume is found to significantly affect the jump probability and, in particular, positive variations increase the probability of being in the high volatility regime. Recent contributions on the relationship between volume of trade and jumps were made by Gabaix et al. (2006), Benz and Hengelbrock (2008) and Milunovich and Joyeux

(2007), among others. In particular, Gabaix et al. (2006) shows that significant spikes in returns can be motivated by trades placed by large investors in relatively illiquid markets, even in the absence of important news about market fundamentals. Large incoming volumes, indeed, destabilize the market and lead to large price and volatility movements. Benz and Hengelbrock (2008) emphasize that in the EUA market, particularly in its initial phases, the concentration of the market among few leading players, the relatively low number of market transactions, the lack of transparency and therefore the discontinuous flow of information available, play a dominant role. Furthermore, the EU-ETS, due to its novelty and uniqueness, represents a market where long-term future contracts are traded in a relatively illiquid market (Milunovich and Joyeux, 2007). In Model 5, the probability of being in the high volatility regime varies between 0.44 and 0, with an average probability of 0.06. This value is consistent with the constant jump probability associated to Model 3 ($\lambda = 0.058$) in Table 4. Figure 2(a) depicts the evolution of the probability associated to the high volatility regime as a function of time for Model 5, while Figure 2(b) shows the dynamics for r_{vol}. The introduction of a time varying probability as a function of the volume of contracts traded improves the fit of the model. As in Models 3 and 4, there is no evidence of significant asymmetry but the excess kurtosis of the normalized residuals, though still significantly different from zero, is sensibly reduced.

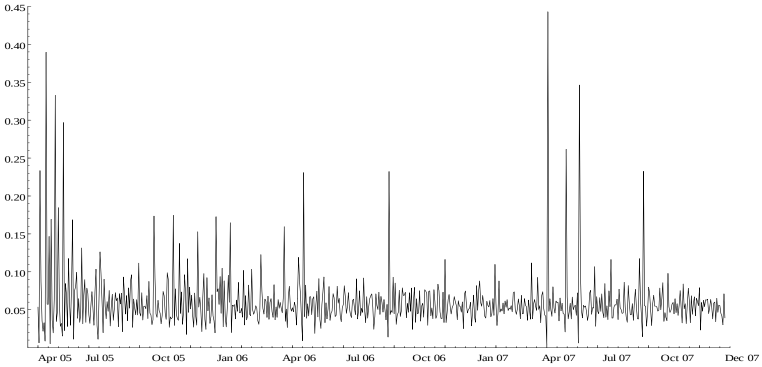

FIG. 2(a) Time varying jump probability (λ_t)

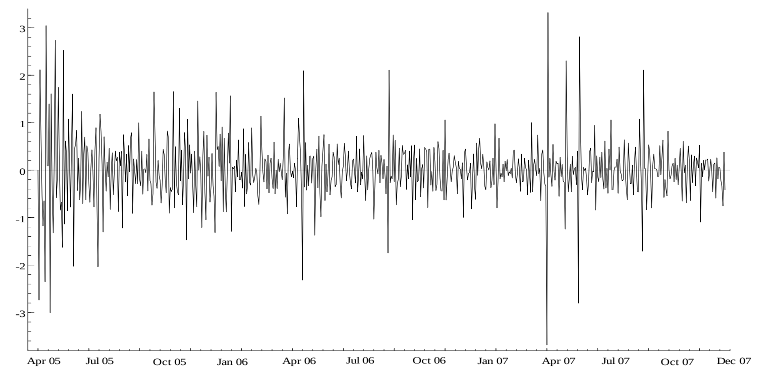

FIG. 2(b) Relative change in volume (r_{vol})

Similar considerations apply when we consider the impact on the jump probability of the announcements made by the European Commission about the approval of each member State's NAP for Phase II (Model 6). We find that announcements concerning the revision and approval of the NAPs for Phase II have substantial consequences on short-term price dynamics and volatility of EUAs since they significantly affect the jump probability. Being the adopted NAPs sensibly more restrictive than the national emission targets originally proposed by the member States (see Table 2 for details), the unexpected relative scarcity of EUAs for Phase II explains the large increase in the jump probability induced by the announcements. The average marginal contribution to the jump probability of the announcement is 24% which is well above the average probability observed when no announcement occurs (7.8%). Figure 3 shows the marginal contribution to the jump probability for each announcement in the case of Model 6 (i.e. $\lambda_t(\gamma_0+\gamma_1 D_{NAP,t}+\gamma_1 r_{vol,t-1})-\lambda_t(\gamma_0+\gamma_1 r_{vol,t-1})$).

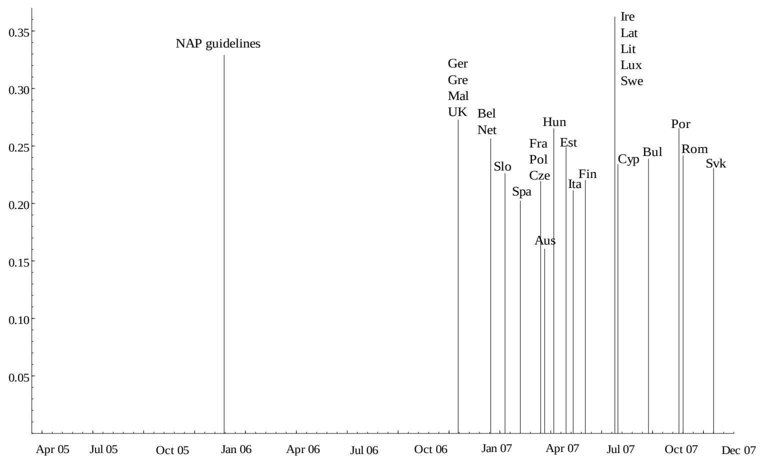

FIG. 3 Anouncements' marginal contribution ($D_{NAPs} * \gamma_{NAP}$)

The last specification, Model 7, accounts additionally for the compliance break period ($D_{04/2006}$) that, as mentioned, is characterized by large negative returns. The relevance of the mixture stands, though only the likelihood shows a significant improvement. The LR test indicates that Model 7 significantly outperforms all previous specifications.

Figure 4(a) reports the probability associated to the high volatility regime as a function of time while Figures 4(b) shows the marginal contribution to the jump probability (λ_t) of the NAPs announcements for Model 7. The probability of being in the high volatility regime varies between 0.57 and 0.005 with an average jump probability of 0.09. The crucial role of changes in the regulatory framework or other policy issues appears clear when comparing Figure 2(a) to Figure 4(a). Changes in policy directives or in the regulatory environment may affect the short-run dynamics of EUA price and volatility, and in particular, our result suggests that announcements concerning the NAPs induce jumps and increase volatility.

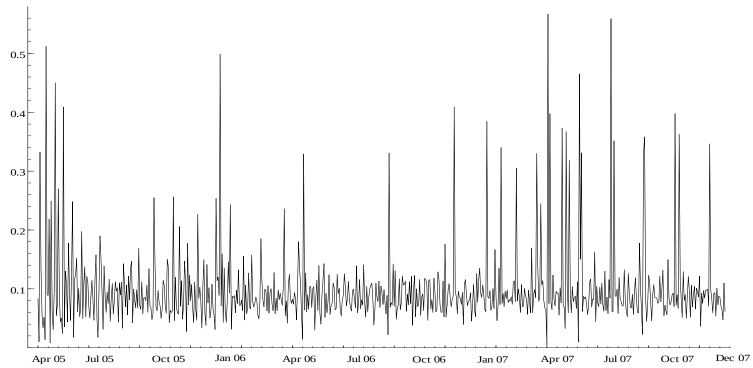

FIG. 4(a) Time varying jump probability (λ_t)

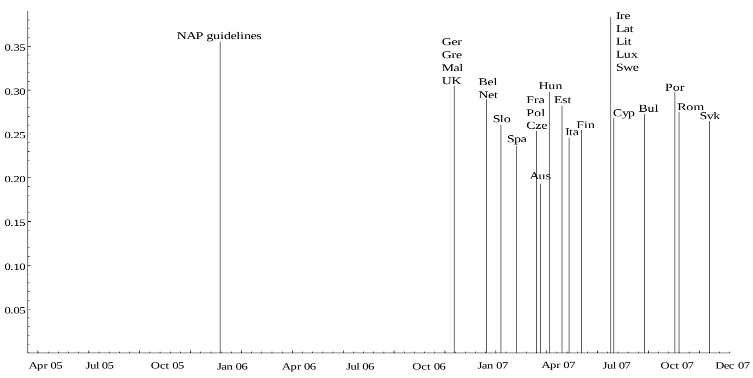

FIG. 4(b) Anouncements' marginal contribution ($D_{NAPs} * \gamma_{NAP}$)

6. CONCLUSION

In this paper we examine the price and volatility dynamics of December 2008 EUA future contracts traded in the ECE. An adequate assessment of short term price and volatility dynamics represents a crucial issue in the EU-ETS since accurate measurement and evaluation of market risk is crucial for portfolio management and hedging in a market that gains in complexity. We first model returns within a standard ARMA-GARCH framework with normally distributed errors. Contrary to previous literature, we find no evidence of the existence of market fundamentals such as energy prices or extreme weather conditions. We show that the standard approach is not adequate because the distributional assumption is unable to properly account for excess skewness and kurtosis in the returns. Furthermore, we detect a number of level and volatility outliers which are the primary cause of the rejection of the Gaussianity assumption. To account for the presence of outliers, we combine the underlying price process with an additive stochastic jump component. The resulting distribution, a mixture of Gaussians, allows for endogenously determined jumps in the return process. The mixing law that we select for the mixture of densities is Bernoulli. The individual distributions in the mixture can then be interpreted as different regimes, e.g. "continuous component" and "jump", while the mixing law gives the probabilities of each regime. We find that a two-regime model based on a Bernoulli Mixture of Normals proves adequate to fit the data and clearly outperforms the standard approach.

The determinants and the occurrence of the jumps are further investigated by introducing a time varying jump probability explained by the daily relative change in the volume of transactions and the change in the regulatory environment induced by the European Commission's approval of Phase II NAPs for each member State. This approach is appealing because it provides useful insights on the occurrence of the jumps and their economic interpretation.

We find that large incoming volumes have a destabilizing effect and translate into sudden and large volatility movements. This result is in line with the finding of recent literature showing that the EU-ETS represents a relatively illiquid market concentrated among few leading players, where the lack of transparency and the discontinuous flow of information available plays a dominant role.

Announcements concerning the revision and approval of the NAPs for Phase II have substantial consequences on short-run dynamics of EUA price and volatility. In particular, our result suggests that announcements concerning the NAPs induce jumps in the process and increase volatility. The instability following the announcements for Phase II NAPs can be explained by the unexpected relative scarcity of EUAs for the second phase. The adopted NAPs were, in fact, sensibly more restrictive than the initial targets proposed by each member State.

This result suggests that the regulatory environment, and thus the mechanism of release of new information, plays a dominant role for market stability. A highly volatile market would fail to give the agents the right incentives for environmental innovation and pollution reduction.

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