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WORKING PAPER

Informational efficiency of the US SO_2 permit market

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ABSTRACT

We test the information efficiency of the market for SO_2 permits in the US. In order to do so, we perform a number of unit root tests and test if the changes in the SO_2 permit price are serially correlated. Furthermore, we test if it would have been possible to earn a profit based on knowledge on the SO_2 permit's price history. The evidence presented in this paper suggests that this market is efficient from an informational point of view. Although one could question this hypothesis from a statistical point of view, economic significance suggests that this market is indeed efficient.

KEYWORDS: SO_2, tradable permits market, market efficiency, unit roots, random walk hypothesis.

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1. Introduction

The sulphur dioxide (SO_2) regulation in the United States gradually evolved from a body of technical regulation with national air quality standards and New Source Performance Standards (NSPS) for new power plants into an innovative trading program in SO_2 emissions allowances. The trading program followed from Title IV of the 1990 Clean Air Amendments that set a goal in 2010 of reducing annual SO_2 emissions by 10 million tons from the 1980 level. Phase I of the trading scheme began in 1995 and affected 263 units at 110 mostly coal-burning electric utility plants located throughout 21 eastern and midwestern States. Phase II which began in 2000 further tightened annual emission restrictions on the larger, higher emitting Phase I plants and set emission restrictions on smaller, cleaner plants. Participation into the program has been strong and it is generally acknowledged that the flexibility of the program provided annual cost savings of approximately \$0.9 billion to \$1.8 billion compared to costs under a command-and-control regulatory alternative (Council of Economic Advisors, 2004). With the offered flexibility, emitters have the freedom to decide how, when and which measures will be taken to lower or not SO_2 emissions. From the point of view of environmental policy, efficiency of permit trading is a key issue that is assumed in most of the work on the way in which permits can be used to address environmental problems (Joskow et al (1998)). Market efficiency can be analysed from a number of different perspectives. First of all, one can look at the process of matching supply and demand. The question to be answered from this perspective is whether the SO_2 permit market resembles a competitive and frictionless market? Joskow et al. (1998) have used this approach to study the efficiency of the US SO_2 permit market. They argue that their analysis of the "*evolution of the sulphur dioxide allowance market indicates that a* *relative efficient private market developed in a few years time, by at least mid-1994*" (Joskow et al. (1998), p. 683).

An alternative approach is to analyse the efficiency from an informational point of view. In general, informational market efficiency has to do with the speed with which new information is reflected in the price of an asset such as a permit to emit SO_2. Markets are efficient if new information is immediately reflected in prices. Hence, as Malkiel puts it: "*if tomorrow's price change will reflect only tomorrows news and will be independent of price changes today*" (Malkiel (2003), p. 59).

In this paper we analyse the efficiency of the US SO_2 permit market from an information point of view. We focus on the price history itself and do not allow explicitly for other variables affecting the value of SO_2 permits. Examining the SO_2 permit price process, allows us to assess the efficiency of this market from an informational point of view: does the price of an SO_2 permit reflect all information that is available to market participants? If the SO_2 permit market is efficient in an informational sense, this would be evidence which supports the hypothesis that market participants have a good idea of the market-clearing price and the influence of new information on this price level. Indeed, if only today's news impact the market price and yesterday's news does not have an influence, market participants must have a good sense of its impact on market-clearing prices. If, on the other hand, the SO_2 market is found to be inefficient and yesterday's news has an impact on today's prices, this would support the hypothesis that participants have no good sense of the market-clearing price and they are either slow to react to new information or overreact.

As market efficiency is one of the key conditions for any permit scheme to work properly, an analysis of the US SO_2 permit market is a worthwhile exercise. The remainder of this paper is organized as follows: the second section discusses some theoretical issues. In the third section, we proceed with an empirical analysis of the US SO_2 permit market. The final section concludes.

2. SO_2 permit prices as random walks

We will start our analysis from the assumption that the SO_2 permit price process can be modelled as a random walk. The idea of a random walk and the efficient market hypothesis are closely related (Malkiel (2003)). If new information is unpredictable and if markets are efficient, price changes must be unpredictable. Consider the following model for the natural logarithm of the price of a SO_2 permit at time t (which we will denote with $p_t = \ln(P_t)$ and we will use 'SO_2 permit price' to refer to the natural logarithm)

$$p_t = p_{t-1} + \gamma + \varepsilon_t \qquad [1]$$

with γ the drift parameter and ε_t the random increment of the process with $\mathrm{E}[\varepsilon_t] = 0$ and $\mathrm{var}[\varepsilon_t] = \sigma^2$. The random increment can be seen as the impact on today's price of today's news. The random walk model in [1] has a number of properties which we will briefly discuss in order to assist the empirical part of this paper. If $\gamma = 0$, each realization of ε_t has a permanent effect on p_t as $p_t = p_0 + \sum_{i=1}^{t} \varepsilon_i$ (with p_0 some constant initial value) and the variance of p_t grows with t as $\mathrm{var}[p_t] = \sigma t$. If, on the other hand, $\gamma \neq 0$, the process of p_t is governed by a linear deterministic trend, γt, as well as a stochastic trend $\sum_{i=1}^{t} \varepsilon_i$ (Enders (1995)). These properties imply that a random walk is a specific type of non-stationary process. From [1], it can be seen that $\Delta p_t = \gamma + \varepsilon_t$ is stationary. Hence, a random walk is a first difference stationary process.

Campbell, Lo and MacKinlay (1997) distinguish three types of random walks based on the properties of the increments ε_t. If they are identically and independently distributed (IID-property), equation [1] is a random walk of type 1 (RW1). The IID-property implies that for two arbitrary functions f and g and scalar $k \neq 0$

$$\operatorname{cov}[f(\varepsilon_t), g(\varepsilon_{t-k})] = 0 \qquad [2].$$

The increments ε_t of a type 2 random walk (RW2) are independent but not identically distributed. Type 3 random walks (RW3) are characterized by dependent but uncorrelated increments. RW3 implies that the equality in [2] holds for all linear functions f and g but not for non-linear functions. From [1] it follows that the condition in [2] could also be written in terms of Δp_t.

If the SO_2 permit prices are random, the permit market is said to be (weakly) efficient as it is impossible to profit by trading on the information contained in the permit price history (Campbell et al. (1997)). However, even if this is not the case, the market may still be efficient as each transaction involves trading costs. Hence, the permit market would still be efficient as long as the information contained in the price history is insufficient to allow a market participant to earn a profit after transaction costs have been accounted for. As such one has to judge whether the results are significant from a statistical as well as an economic point of view (Malkiel (2003)).

3. Efficiency of the US SO_2 permit market

3.1. Data

To analyse the behaviour of the US SO_2 permit prices we have used monthly data from August 1994 to December 2001 (89 observations). Figure 1 shows the series based on price information by the US Environmental Protection Agency (EPA). EPA (2001) presents price data from Cantor Fitzgerald and Fieldston Publications. Figure 1 is based on data from Fieldston Publications. We have chosen not to include the year 2002 as the SO_2 and other permit programs in the US were severely affected by the Californian energy crisis. We have done so because it is not unreasonable to assume that the temporary halting of the NO_X permit market did not have an important impact on the market for SO_2 permits. Given the data that is available to us, including the Californian energy crisis could put to much weight on this a-typical period.

Table 1 reports the summary statistics for both p_t as well as its first difference Δp_t. Figure 1 shows the price history of p_t.

`[insert table 1 about here]`

`[insert figure 1 about here]`

3.2. Unit root tests

A random walk is a first difference stationary process. Hence, the first issue to be looked at is whether the SO_2 permit price series contains a unit root while the first difference of this series does not.

A number of alternative tests are available to analyze if a process contains a unit root. The Augmented Dickey Fuller (ADF) test uses the following regression

$$p_t = \alpha + \gamma t + \theta p_{t-1} + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \qquad [3]$$

-where α and γ are drift and deterministic trend components- to determine if $\theta = 1$ against the alternative that $\theta < 1$. The deterministic trend component is added to [3] to allow the process to be trend stationary. The latter would be the case if $\gamma \neq 0$ and $\theta = 0$. If that were the case, the permit price process could be made stationary be de-trending the data. The test-statistic which is used for this test is the t-statistic on θ . However, under the null of a unit root, this statistic does not have a standard distribution. Many authors have used Monte Carlo methods to determine the critical values of the distribution (see e.g. Hamilton (1994)).

The ADF-test requires that error terms ε_t are independent and homogeneous. The addition of lagged differences in [3] is meant to remove any autocorrelation from the error terms. The number of lags *k* that should be included to make the ε_t sequence white noise is, however, unknown. Too many lags will reduce the power of the test whereas if too few are included, the test may lead to seriously biased conclusions (Verbeek (2000)). Various alternatives have been explored to select the optimal lag length. Model specification tests use the Akaike information criterion (AIC) or the Bayesian information criterion (BIC). However, as Ng and Perron (1995) have shown, these criteria often select very small values of *k.* The general to simple procedure starts with a specified number of lags *k* and reduces that number to *k*-1 lags if the lag of length *k* is insignificant. As an alternative, one can use the Ljung-box test to test for serial correlation and add lags until the test fails to reject no serial correlation at a predefined level.

The test due to Phillips and Perron (1988) (PP) is a generalization of the ADF-test and is less demanding with respect to the error terms. The PP-test allows the error terms to be weakly dependent and heterogeneously distributed by including a weighting function and various lags of the error process to calculate a consistent estimate of the variance.

Determining if $\theta = 1$ using [3] including a trend and/or drift if it holds that $\gamma = 0$ or $\alpha = 0$ reduces the degrees of freedom and the power of the test and hence, one could conclude that the process for p_t contains a unit root when this is not the case. Furthermore, the distribution of the test statistics used to determine if $\theta = 1$ depend on the question if a deterministic trend and/or drift is included or not. On the other hand, not including the deterministic trend if it holds that $\gamma \neq 0$ results in an upward bias in the estimated value of θ (Enders (1995)).

Both the ADF and PP's null hypothesis is that a series contains a unit root. However, unit root tests often lack power. Kwaitkowski, Phillips, Schmidt and Shin (1992) (KPSS) propose a test whose null is stationarity. Their test is based on the residuals from

$$p_t = \alpha + \gamma t + \varepsilon_t \qquad [4].$$

The test statistic is given by $\sum_{t=1}^{T} S_t^2 \big/ \hat{\sigma}^2$ with $S_t = \sum_{s=1}^{t} \varepsilon_s$ and $\hat{\sigma}^2$ an estimator of the variance of the error terms. The 5% KPSS critical value for the null of trend stationarity equals 0.146. To test the null of stationarity, the trend is omitted from [4] and the 5% critical value equals 0.463 (Kwiatkowski et al. (1992)). In order to compute $\hat{\sigma}^2$, KPSS propose a procedure similar to PP and include a weighting function to correct for autocorrelation.

Perron (1997) proposes a test that includes the possibility of changes in the intercept and slope of the deterministic trend in [1]. Indeed, if there is a one-time increase in the intercept of a trend-stationary process, standard unit root tests are biased towards accepting the null of a unit root. Three models, all of which use OLS estimates, are used to test if $\theta = 1$ in the presence of a break at time $t = \tau + 1$. The test statistic in all three cases is the t-statistic for the test that $\theta = 1$. The distribution of this t-statistic is non-standard. Perron (1997) however provides critical values for various sample sizes and models.

The first model tests the null of a unit root with a one-time shift in the non-stationary process against the alternative of a trend-stationary process with a one-time shift in the intercept. The test uses the following regression

$$p_t = \alpha + \gamma t + \psi D_L + \theta p_{t-1} + \delta D_P + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \qquad [5]$$

with $D_L = \begin{cases} 0 \Leftrightarrow t \leq \tau \\ 1 \Leftrightarrow t > \tau \end{cases}$ and $D_P = \begin{cases} 0 \Leftrightarrow t \neq \tau + 1 \\ 1 \Leftrightarrow t = \tau + 1 \end{cases}$.

The second model allows for a change in both the intercept and the slope of the deterministic trend and uses

$$p_t = \alpha + \gamma t + \psi D_L + \varphi D_L t + \theta p_{t-1} + \delta D_P + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \qquad [6].$$

The alternative hypothesis in [6] is a one-time change in the intercept and slope of a trend-stationary process. Finally, the third model allows for a change in the slope of a trend-stationary process but assumes that both segments of the trend are joined at the time of the break. The test uses

$$p_t = \alpha + \gamma t + \varphi D_L (t - \tau) + \theta p_{t-1} + \sum_{i=1}^{k} \beta_i \Delta p_{t-i} + \varepsilon_t \qquad [7]$$

to test if $\theta = 1$. Hence, the hypothesis of a unit root is tested against the alternative of a change in the slope of a trend-stationary process.

Perron (1997) proposes various alternatives to select the break data t endogenously. The first minimizes the t-statistics on $\theta = 1$. The second alternative minimizes the t-statistic on ψ (model 1) or on φ (models 2 and 3). The third alternative is similar to the second one but uses the absolute values of the t-statistics. To determine the lag length k Perron (1997) proposes the general to specific procedure.

We have performed unit root tests for the SO_2 permit price series (p_t) as well as the first difference of this series (Δp_t). Table 2 reports the results.

[insert table 2 about here]

First of all, we have used the ADF test with a general to specific procedure starting with 20 lags using the t-statistic on the last lag. We have estimated all models both with trend and constant, with constant and without trend or constant. However, neither the joint test of a unit root and no linear trend nor the joint test of a unit root and no constant was acceptable. For the test in levels, the test statistic for the former equalled 3.93 (10% critical value is 5.47) while the test statistic for the hypothesis of a unit root but no constant equalled 1.86 (10% critical value: 3.86). The test for the first differences was 4.27 (unit root but no trend) and 4.24 (unit root but no constant). Hence table 2 only reports ADF-tests for the model without trend and constant. Although the ADF test reveals an explosive process for p_t, the unit root hypothesis is rejected for $\Delta p_{t.}$

Our second test is the PP-test. The table only reports the results for the test without a trend but the conclusions are not affected if a trend is added. The PP-test clearly reveals that the series p_t is difference stationary. Thirdly, we have done a KPSS-test to confirm our findings from the ADF en PP tests. The null of stationary series is clearly rejected for p_t but the test is unable to reject this hypothesis for Δp_t.

Because the series possibly exhibits a break in 1998, we have used the Perron (1997) test to check if the series contain a unit root if this structural break is accounted for. We used the procedure that minimizes the t-statistic on θ to select the break date. With the exception of the estimates of equation [7] for Δp_t, all endogenously determined breaks are located between February 1998 and May 1998. The estimates for p_t further reveal positive significant values of ψ in [5] and φ in [7]. The estimates of δ in [5] and [6] are also positive and significant for Δp_t. The significance of these values notwithstanding, table 2 supports the hypothesis that the series p_t contains a unit root and the series Δp_t is stationary.

Based on the evidence from the ADF, PP and PP with endogenously determined time breaks, the hypothesis that the SO_2 permit price series contains a unit root can not be rejected. For the first differenced series on the other hand, the evidence clearly suggests that the hypothesis of a unit root should be rejected. The KPSS-test does not allow us to accept the null of stationarity for p_t while it fails to reject the null of stationarity for Δp_t. All in all, this suggests that p_t is a non-stationary process. However, this is not sufficient to conclude that the series is a random walk.

3.3. Tests of the random walk hypothesis

Although the unit root tests have clearly shown that permit prices contain a unit root, this is not sufficient to adopt the random walk hypothesis (Campbell et al. (1997)). Indeed, the various random walk hypotheses impose restrictions on the error process that have not been analysed so far. We will start with the restriction imposed by the RW3-model. For the RW3-model, condition [2] requires that all autocorrelations between Δp_t and Δp_{t-k}, $\rho(k)$, equal zero for all values of $k>0$. Table 3 provides estimates of the autocorrelation coefficients for the level (p_t), Δp_t and $(\Delta p_t)^2$. The Ljung-Box Q(k) statistic allows to assess the significance of these coefficients. The results for p_t reveal a typical pattern for a non-stationary series: the autocorrelation coefficient for k=1 is close to unity and dies out slowly (Enders (1995)). With respect to Δp_t, the results suggest that $E\left[\Delta p_t \ \Delta p_{t-k}\right] \neq 0$ for various lag lengths. However, the correlation coefficients seem to be small and are only significant at the 5% level until k reaches 5. The significance of the autocorrelation coefficients is evidence against the hypothesis that the series p_t is a random walk of type 3. The evidence with respect to the autocorrelation coefficients for $(\Delta p_t)^2$

suggests that the series for Δp_t exhibits volatility clustering as $E\left[(\Delta p_t)^2(\Delta p_{t-k})^2\right] \neq 0$ implies that $E\left[\sigma_t^2, \sigma_{t-k}^2\right] \neq 0$. The estimated coefficients are, however, small and are no longer significant at a 5% level for values of $k>5$. Furthermore, one should be very careful with respect to the Q(k)statistic as it is a joint test on all autocorrelations up to a certain level k. Hence significance (say at $k=5$) could be due to one strongly significant autocorrelation (for instance at $k=2$).

[insert table 3 about here]

A second test on Δp_t is based on the ratio of the variances at two different frequencies. Let's assume that we compare Δp_t's at a monthly interval and a larger interval of q months. If we were to set $q = 3$ for instance, this would mean that we would compare monthly and quarterly changes in the permit price. The former equals Δp_t while the return at an interval equal to q equals $\sum_{i=0}^{q-1} \Delta p_{t-i}$. Because of [2], the variance of the latter equals q times the variance of the former. Campbell et al. (1997) derive a test statistic under RW3 that allows testing if the variance of the return at the q-interval is equal to q times the return at the monthly interval. The test statistic is derived from the sample autocorrelations and is given by

$$\psi^*(q) = \frac{\sqrt{T-1}\,(VR(q)-1)}{\sqrt{\hat{\theta}}} \sim N(0,1) \qquad [8]$$

with

$VR(q) = 1 + 2\sum_{k=1}^{q-1}\left(1 - \frac{k}{q}\right)\hat{\rho}(k)$ and $\hat{\theta}(q)$ a heteroskedasticity-consistent estimator of the variance of $VR(q)$ (for details, see Campbell et al. (1997), p. 55).

[insert table 4 about here]

The evidence presented in table 4 seems to reinforce the conclusions from table 3. In line with the evidence presented in table 3, the fact that $VR(q) > 1$ implies that the autocorrelations are positive. Furthermore, the variance ratio test rejects the hypothesis that the permit price series is a random walk of type 3 for levels of $q < 6$ at a 5% level of significance. However, for levels of $q \geq 6$, which compare monthly returns to for instance, yearly returns, the test fails to reject the null of no significant autocorrelation among the returns at the 5% level.

Both the evidence from the autocorrelation coefficients as well as the variance ratio tests offers some support for the hypothesis that the permit price process is not a random walk of type 3. The evidence presented here is not all that different from the evidence for financial markets. Lo and MacKinlay (1999) for instance find that autocorrelations are not all zero. The evidence against the hypothesis that the series is a random walk of type 3 is, however, not overwhelming. The autocorrelation coefficients presented in table 3 for instance are small and the variance ratio fails to reject the RW3 hypothesis for levels of $q \geq 6$. The size of the autocorrelations coefficients would suggest that one can question whether the significance in a statistical sense extends to significance in an economic sense.

3.4. Predictability

The question that emerges from the previous paragraph is whether the significant autocorrelations can be exploited from an economic point of view. If these significant autocorrelations can be exploited to earn a profit, one can not argue that SO_2 permit markets are efficient. New information which arrives and has an impact on the value of an SO_2 permit is not immediately reflected in its price. Hence, from an informational efficiency point of view, this would be evidence against efficient markets. If one can not earn such a profit it follows that all information that affects the value of SO_2 permits is included in the permit price. However, given the statistical significance of autocorrelations, new information is only reflected in prices up to such a level where it is possible to profit from the price history. Hence, from an economic point of view, exploiting the significant autocorrelations fully is not rational and one can argue that permit prices reflect all information which is significant from an economic point of view.

It follows that the issue that needs to be addressed is whether the SO_2 permit price history can be used to earn a profit. Obviously, there are various ways to test this hypothesis. We have chosen to estimate a model using the permit price history and see whether it could have been used to predict permit prices with a relative high level of certainty. We have estimated an AR(2)-GARCH(1,1) model $\Delta p_t = \delta_1 \Delta p_{t-1} + \delta_2 \Delta p_{t-2} + \xi_t$ with $\xi_t = \eta_t \sqrt{\sigma_t}$, $\eta_t \sim IID(0,1)$ and $\sigma_t^2 = \omega + \alpha \xi_{t-1}^2 + \beta_1 \sigma_{t-1}^2$. The GARCH(1,1) model for the random increments guarantees that the error process is white noise. The model requires that $|\delta_i| < 1$, $i = 1, 2$; $\omega > 0$, $\alpha \geq 0$ and $\beta_i \geq 0$ (Bollerslev (1986)). Table 5 reports the results.

The model was obtained after some experimentation. It was chosen because its residuals do not exhibit any autocorrelation. Secondly, the Q(k) statistics for the squared residuals reveal that the model removes most of the autocorrelation as only the Q(3) statistic is significant an only at the 10% level of significance.

The results in table 5 suggest that it would be hard to make a profit based on the past information. Although the estimates are significant, the R^2 is very low which is indicative of the fact that the model is not able to predict future SO_2 permit prices with much certainty. Hence, one can question if it would be possible to profit from knowledge of price history on the SO_2 permit market. As was the case with the non-zero autocorrelations coefficients, this feature is also present in financial markets (Malkiel (2003)). The estimates of the GARCH-terms confirm that the variance in SO_2 permit markets clusters.

This suggests that it would have been impossible to exploit the significant autocorrelations from an economic point of view. Market participants react fast to new information and continue to do so up until the point where it is no longer possible for them to use the information in their market behavior. This clearly suggests that US the SO_2 permit market is efficient and that SO_2 permit prices reflect all information which is significant from an economic point of view. This clearly suggests that market participants have a good understanding of the price process. If this were not the case, returns would be predictable. Assume for instance that market participants tend to overreact and that news arrives which causes the market price to jump upwards. To the extent that market participants overreact today, we would expect a correction (negative return) in the future. Hence, permit prices would be predictable: a spike would be followed by a correction. If, on the other hand market participants are slow to adjust prices, permit prices would exhibit a series of positive or negative returns. Hence, again, they would be predictable.

4. Discussion and conclusion

The evidence presented in this paper suggests that the market for SO_2 permits in the US is not all that different from financial markets. For financial markets, the random walk hypothesis (RW3) is also often rejected. However, as is the case for the SO_2 permits market; economic profitable predictability is mostly rejected as well. Hence, although one cannot reject the hypothesis that this market is weakly efficient from a statistical point of view, the economic significance of the predictability is very limited if not nonexistent.

The evidence presented in this paper suggests that new information is reflected in the permit price fast. The SO_2 permit market is basically as efficient as financial markets. This is clearly important as it is indicative of the fact that the value of SO_2 permits reflects all relevant information. This suggests that market participants have a good understanding of the price process and have a good understanding of the way in which new information affects the market-clearing price. The evidence presented in this paper supports the conclusion in Joskow et al. (1998) that is would be a

> *"hard to argue that bidders in the 1993 auctions had a good idea of a single market-clearing price. It would be a good deal easier to make this argument for the 1994 auctions."* (p. 681).

Based on our analysis of the history of SO_2 permit prices, we reach the same conclusion from a different perspective.

New information that increases the variance has the tendency to cluster. If an event increases uncertainty, this uncertainty does not return back to its previous level in one period. The significant GARCH-effects suggest that it takes time before it settles down.

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Table 1: summary statistics for p_t and Δp_t

	p_t	Δp_t
Mean	4.9177	0.0034
S.e. of mean	0.0317	0.0010
t-Statistic	154.9263	3.2140
Variance	0.2994	0.0004
Skewness	-0.1806	1.7069
Kurtosis	-1.1389	11.4473
Jarque-Bera	5.2943	2,063.1517
Observations	89	88

Table 2: Unit root tests for p_t and Δp_t[(1)]

Variable	Test	Test-statistic	Lags[(2)]
p_t	ADF	0.2006	9
p_t	PP (constant)	-1.4337	4
p_t	KPSS[(3)] (no trend)	0.1739$^{(***)}$	4
p_t	KPSS[(3)] (trend)	0.8389$^{(**)}$	4
p_t	PP – eq. [5]	-4.4863	7
p_t	PP – eq. [6]	-4.6832	11
p_t	PP – eq. [7]	-3.2258	7
Δp_t	ADF	-2.9194$^{(***)}$	8
Δp_t	PP	-7.9048$^{(***)}$	4
Δp_t	KPSS[(3)] (no trend)	0.1315	4
Δp_t	KPSS[(3)] (trend)	0.0935	4
Δp_t	PP – eq. [5]	-9.0228$^{(***)}$	0
Δp_t	PP – eq. [6]	-9.0366$^{(***)}$	0
Δp_t	PP – eq. [7]	-3.3899	8

[(1)]*, ** and *** refer to significance at the 10%, 5% and 1% level

[(2)] Lags refer to the number of lags obtained following the general to specific procedure for the ADF and endogenous break (PP-models) and to the number of lags used for the KPSS and PP tests.

[(3)]KPSS 1% critical value without (with) trend equals 0.739 (0.216), the 5% critical value without (with) trend equals 0.463 (0.146)

Table 3: autocorrelations for p_t, Δp_t and $(\Delta p_t)^{2(1)}$

Lag k	p_t		Δp_t		$(\Delta p_t)^2$	
	Autocor.	Q(k)	Autocor.	Q(k)	Autocor.	Q(k)
1	0.967	86.20	0.180	$2.96^{(*)}$	-0.011	0.01
2	0.924	165.67	0.268	$9.60^{(***)}$	0.344	$10.91^{(***)}$
3	0.862	235.76	-0.046	$9.81^{(**)}$	0.075	$11.44^{(***)}$
4	0.803	297.33	-0.003	$9.81^{(**)}$	-0.056	$11.75^{(**)}$
5	0.744	350.82	-0.050	$10.05^{(*)}$	0.012	$11.76^{(**)}$
6	0.688	397.07	-0.023	10.10	-0.034	$11.88^{(*)}$
7	0.633	436.68	0.181	$13.31^{(*)}$	-0.049	$12.12^{(*)}$
8	0.566	468.71	0.054	$13.61^{(*)}$	-0.045	12.32
9	0.494	493.49	-0.081	14.27	-0.079	12.95
10	0.428	512.32	-0.047	14.50	-0.073	13.50
11	0.365	526.16	-0.072	15.03	-0.034	13.62
12	0.306	536.03	-0.054	15.35	-0.071	14.16
13	0.251	542.76	-0.121	16.91	0.095	15.13
14	0.204	547.27	-0.064	17.35	-0.076	15.76
15	0.161	550.12	0.014	17.37	0.041	15.95
16	0.117	551.66	0.018	17.41	0.004	15.95
17	0.073	552.26	-0.027	17.49	-0.059	16.34
18	0.029	552.37	-0.133	19.52	-0.059	16.74

(1)*, ** and *** refer to significance at the 10%, 5% and 1% level

Table 4: Variance ratio test[1]

q	$VR(q)$	$\psi^*(q)$	Sig. $\psi^*(q)$
3	1.4188	2.2401$^{(**)}$	0.0251
4	1.5137	2.0299$^{(**)}$	0.0424
5	1.5703	1.8931$^{(*)}$	0.0583
6	1.5913	1.7470$^{(*)}$	0.0806
7	1.5998	1.6251	0.1041
8	1.6501	1.6457$^{(*)}$	0.0998
9	1.7004	1.6773$^{(*)}$	0.0935
10	1.7261	1.6600$^{(*)}$	0.0969
11	1.7384	1.6221	0.1048
12	1.7375	1.5645	0.1177

[1]*, ** and *** refer to significance at the 10%, 5% and 1% level

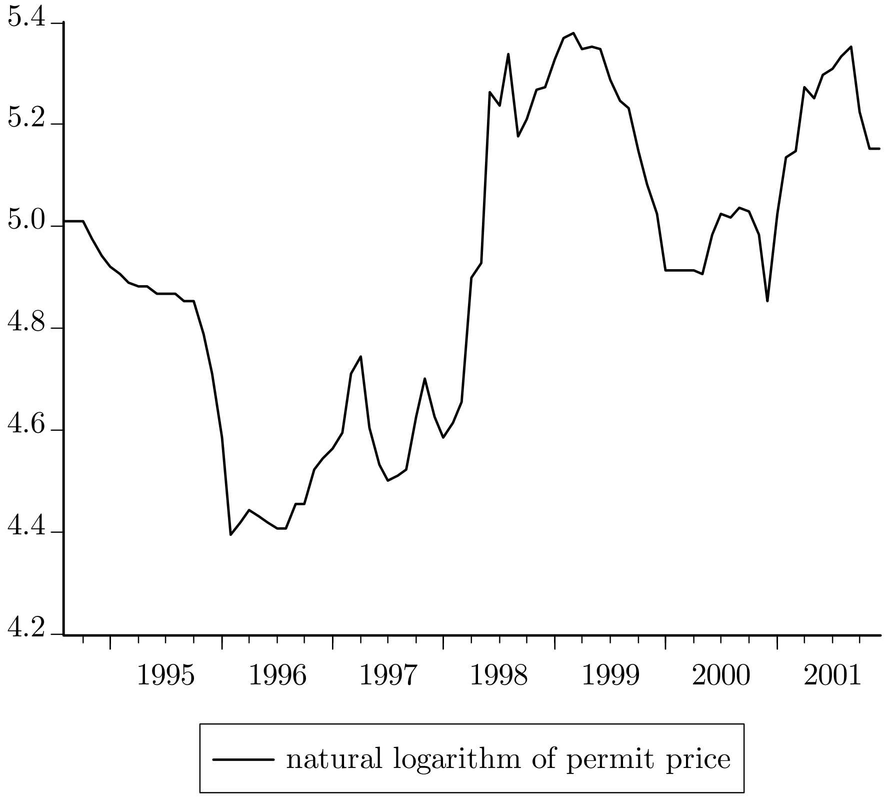
Table 5: AR(2)-GARCH(1,1) estimates[(1)]

	Estimate	Stand. Error[(2)]
AR(2)		
Δp_{t-1}	0.212831$^{(*)}$	0.122596
Δp_{t-2}	0.218871	0.168118
GARCH(1,1)		
C	0.000714$^{(**)}$	0.000356
ξ^2_{t-1}	0.357813$^{(*)}$	0.193965
σ^2_{t-1}	0.587051$^{(***)}$	0.154951
Uncentred R^2	0.0846	
Residuals		
Q(3)	0.6913	
Q(4)	0.8000	
Q(5)	1.7584	
Squared. Res.		
Q(3)	3.6226$^{(*)}$	
Q(4)	4.2237	
Q(5)	4.3018	

[(1)]*, ** and *** refer to significance at the 10%, 5% and 1% level

[(2)]Standard Errors are Heteroskedasticity-consistent (Bollerslev-Wooldridge)

Figure 1: natural logarithm of the SO_2 permit price



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