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Abstract

In this paper we test the empirical implications of a simple pricing model for commodity futures for the marginal process of prices of Sugar futures. According to the pricing model, the futures price bias depends linearly on the conditional variance. We find significant coefficients, from monthly as well as daily data, if the conditional variance is modelled using the GARCH-M model put forward by Engle, Lillien and Robins (1987). These estimates imply contango in the futures market and a net hedging demand on the long side of it.

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1.Introduction

Many models consider relations between prices of futures contracts and corresponding spot prices (see e.g. Anderson and Danthine (1983), Hirschleifer (1989, 1990)). In this paper, we test whether a simple special case of these models, which is discussed at length e.g. in the recent textbook by Duffie (1989), is relevant for the futures market for Sugar. Because of data limitations the tests are directed towards the implications for the marginal process of the futures prices only. According to the model under consideration, the futures price bias depends linearly on the conditional variance. Three different estimates of the monthly conditional variance from january 1972 to june 1989 are considered. First, the monthly conditional variance is estimated from daily data in the previous month. Second, the GARCH model proposed by Bollerslev (1986) is used to derive estimates of the conditional variances and finally the GARCH-M model proposed in Engle, Lilien and Robins (1987) is used. In addition daily data since 1979 are used to estimate the dependence of the futures price bias on the daily conditional variance. The GARCH-M models yield significant coefficients for the conditional variance, which imply contango in the Sugar futures market and a net hedging demand on the long side of the market.

The plan of this paper is as follows. In section 2 we present the pricing model under consideration. The stylized facts in the data are described in section 3. The significance of the three different estimates of the monthly conditional variance is tested in sections 4, 5 and 6. Daily data are analyzed in section 7, while section 8 concludes.

2. A simple pricing model

Consider the demand for commodity futures contracts of agents with mean variance utility functions, who possibly have positions on next periods spot market for the commodity under consideration. Assume that the agents can trade the corresponding futures contract as well as a stock market portfolio. Let $q_{t+1}^{(i)}$ denotes agent i's next period's spot market position, while s_{t+1} denotes next periods spot price. If futhermore $f_t^{[k]}$ denotes the price of the corresponding futures contract k periods before expiration, $B_t^{(i)}$ is the amount invested in the stock market portfolio and r_{t+1} is the return on the stock market portfolio, the wealth of agent i in the subsequent period, $W_{t+1}^{(i)}$, can be expressed as

$$W_{t+1}^{(i)} = B_t^{(i)} (1+r_{t+1}) + q_{t+1}^{(i)} s_{t+1} + y_t^{(i)} (f_{t+1}^{[k-1]} - f_t^{[k]}), \tag{1}$$

where $y_t^{(i)}$ denotes the agent's futures position in period t. Note that it is assumed that are no margin requirements in cash. Moreover we assume that the returns on the stock market portfolio are uncorrelated with the spot and futures prices, which is not in conflict with the empirical evidence for commodity futures (Dusak (1973)). Every agent is assumed to choose his position in the futures market by maximizing a mean-variance utility function, i.e.

$$y_t^{(i)*} = \underset{y_t^{(i)}}{\text{argmax}} \; \{ E_t[W_{t+1}^{(i)}] - \rho^{(i)} \, Var_t[W_{t+1}^{(i)}] \}, \tag{2}$$

where $\rho^{(i)}$ is agent i's risk aversion coefficient and $E_t[.]$ and $Var_t[.]$ denote the conditional expectation and conditional variance respectively. Differentiation with respect to $y_t^{(i)}$ in (2) yields the first order condition

$$\begin{aligned} y_t^{(i)*} &= 0.5 \; \{ E_t[f_{t+1}^{[k-1]}] - f_t^{[k]} \} \; / \; \{ Var_t[f_{t+1}^{[k-1]}] * \rho^{(i)} \} \\ &\quad - q_{t+1}^{(i)} \, Cov_t[s_{t+1}, f_{t+1}^{[k-1]}] \; / \; Var_t[f_{t+1}^{[k-1]}]. \end{aligned} \tag{3}$$

The first and second term in (3) are known as the pure speculative demand and the pure hedge demand respectively, as the second term vanishes for speculators ($q_t^{(i)} = 0$) while the first term vanishes if $\rho^{(i)} \to \infty$. Equilibrium on the futures market requires

$$\sum_{i=1}^{N} y_t^{(i)} = 0 \tag{4}$$

where N is the number of agents in the economy. From (3) and (4) one can easily derive the following expression for the futures price bias $f_t^{[k]} - E_t[f_{t+1}^{[k-1]}]$,

$$E_t[f_{t+1}^{[k-1]}] - f_t^{[k]} = 2\ \rho\ q_{t+1}\ \mathrm{Cov}_t[s_{t+1}, f_{t+1}^{[k-1]}], \tag{5}$$

where $\rho = (\ \Sigma_{i=1}^{N} (\rho^{i})^{-1}/N\ )^{-1}$ is the markets risk aversion and $q_{t+1} = \Sigma_{i=1}^{N} q_{t+1}^{(i)}/N$ is the net hedging pressure on the market. If we finally assume the absence of basis risk ($\mathrm{Cov}_t[s_{t+1}, f_{t+1}^{[k-1]}] = \mathrm{Var}_t[f_{t+1}^{[k-1]}]$) and a constant net hedging pressure ($q_{t+1} = q$), equation (5) yields testable implications on the marginal process for $\Delta f_{t+1} = f_{t+1}^{[k-1]} - f_t^{[k]}$,

$$\Delta f_{t+1} = \delta\ \mathrm{Var}_t[\Delta f_{t+1}] + \varepsilon_{t+1}, \tag{6}$$

where $\delta = 2\ \rho\ q$ and where ε_{t+1} is independent of past realizations of Δf_{t+1}. Carter et.al. (1983) have considered a model similar to (5) and avoided the assumption of a constant net hedging pressure by making use of data on the percentage of speculators who were net long. Carter et.al. (1983) however assumed conditional homoskedasticity.

The basic model that is outlined above has been refined by many authors (see e.g. Anderson and Danthine (1983) and Hirschleifer (1989, 1990)) taking into account e.g. production uncertainty and covariance of commodity prices with the market portfolio. In this paper we will restrict ourselves to testing the relevance of the basic model for Sugar futures.

3. Some stylized facts on Sugar futures.

Sugar-11 contracts which expire in January, March, May, July, September and October are traded on the New York Coffee, Sugar and Cocoa Exchange (NYCSCE). A data-series of daily observations on price changes in the contract which is closest to expiration (excluding last month observations to avoid potential delivery obligations) will be analyzed in sections 4 to 7 to test the pricing model that was presented in the previous section. The series of monthly changes which is constructed from this data-set is presented in figure 1. Throughout the paper all prices are in dollars per 10,000 lbs. Two periods of bad weather (1973/1974 and 1979/1980) caused high international sugar prices (see FAO (1985)), which are evidently reflected in the futures prices. Figure 1 also clearly shows that volatility of the Sugar futures prices is time varying and that the marginal distribution of the price changes is fat tailed. These are well known stylized facts, which hold true for many futures (compare e.g. Taylor (1986)). The kurtosis of the monthly series is estimated as 11.9, while the standard Lagrange Multiplier test statistic for sixth order ARCH yields the very significant value of 63.6. The LM test for sixth order autocorrelation yields a value of 17.6 which is close to the 1% critical value of a χ^2-distribution with six degrees of freedom. Note however that standard tests for serial correlation in the mean will be biased upward in the presence of conditional heteroskedasticity (see e.g. Diebold (1987)) and thus lead to over-rejection. In figure 2 the monthly price changes are plotted for the sub-sample from January 1982 to June 1989 during which period the Sugar prices were much more stable. For this sub-period the estimated kurtosis is 3.8 and the LM tests for ARCH and autocorrelation yield the insignificant values of 2.8 and 5.7.

The two daily series, which will be analyzed in section 7, are plotted in figures 3 and 4. The conditional heteroskedasticity is obviously even more pronounced for the full sample than it is for the less volatile sub-sample. Note that, according to the results on temporal aggregation of GARCH processes in Drost and Nijman (1990), the conditional heteroskedasticity in the daily data in the subsample is not conflicting with the apparent homoskedasticity of the monthly data.

4. Tests based on the use of daily data to estimate monthly conditional variances.

According to the pricing model discussed in section 2, the futures price bias depends linearly on the conditional variance. In sections 4, 5 and 6 this implication will be tested for three different estimates of the monthly conditional variance. In this section we derive estimates of the monthly conditional variance from daily data, while in sections 5 and 6 the variance is assumed to be generated by a GARCH and a GARCH-M model respectively.

If subsequent changes in futures prices are approximately uncorrelated, which is one of the stylized facts discussed in the previous section, the sample variance in daily data in the current month times the number of trading days in the next month, appears to be a natural first estimator of the conditional variance in the next month. This estimator of the conditional variance can be motivated e.g. as an inefficient estimate of the conditional variance in an assumed integrated GARCH model (see Engle and Bollerslev (1986)) which is close to the model that will be estimated in section 7. The variance estimates which are generated from daily data are presented in figure 5, from which it is once more evident that the volatility is strongly time-varying.

Estimates of (6) are presented in Table 1 for the full sample period as well as for the "erratic" period 1972-2 to 1981-12 and for the less volatile period 1982-1 to 1989-5. The fourth column in the table presents the coefficient estimate, while the corresponding routinely estimated t-statistic is reported in colum five. This t-statistic will not be asymptotically standard normally distributed if the error terms in (6) are heteroskedastic, which is not excluded by the model. Consequently we report in column six heteroskedasticity consistent t-statistics, which have been computed along the lines of White (1980). Lagrange Multiplier tests for up to sixth order autocorrelation (A) and conditional heteroskedasticity of the ARCH form (CH) are presented in column seven and eight. Under the null hypothesis of no autocorrelation and no conditional heteroskedasticity both test statistics are asympototically centrally χ^2 distributed with 6 degrees of freedom. The critical values of this distribution at the 5% and 1% level are 12.6 and 16.8 respectively. In the final column of the table an estimate of the kurtosis of the distribution of the error terms is presented. Large values of the kurtosis parameter indicate fat tails. The kurtosis of the normal distribution equals 3.

The first row of Table 1 contains ordinary least squares (OLS) estimates for the full sample period. The estimate of δ is negative, which suggests a net hedging demand on the long side of the market. According to the t-statistics, however, this parameter is insignificantly different from zero. Moreover the test for ARCH indicates the presence of strong conditional heteroskedasticity and the kurtosis estimate implies a very fat tailed error distribution, which imply that the OLS estimator is very inefficient. A simple more efficient estimator is the generalized least squares (GLS) estimator. Because according to the model $E_t[\varepsilon^2_{t+1}] = Var_t[\Delta f_{t+1}]$ and $E\, \varepsilon_t \varepsilon_s = 0$ if $t \neq s$, the generalized least squares regression coincides with the ordinary least squares regression on the transformed equation

$$\Delta f_{t+1}/se_t[\Delta f_{t+1}] = \delta\, se_t[\Delta f_{t+1}] + \tilde{\varepsilon}_{t+1}, \qquad (7)$$

where $se_t[\Delta f_{t+1}] = (Var_t[\Delta f_{t+1}])^{1/2}$. According to the last three lines of Table 1 this transformation reduces the conditional heteroskedasticity problem, but does not solve it as it should if (6) and the estimate of the conditional variance are correctly specified. Note that the test statistics in the generalized least squares regression are based on the transformed residuals $\tilde{\varepsilon}_{t+1}$. Note also that after the correction for conditional heteroskedasticity the error distribution is still fat tailed, which is however a stylized fact for many financial series and does not invalidate the model. The coefficient estimates in the last three rows of Table 1 show that for none of the sample periods under consideration there is any evidence of a significant impact of the conditional variance on the futures price bias. Possible explanations for this empirical finding are misspecification of (6), misspecification of the estimate of the conditional variance that is used in this section or a net hedging pressure q_t in (5) which is very close to zero in the sample. In the subsequent section we will analyze to what extent misspecification of the conditional variance can be responsible for the empirical findings in Table 1.

5. Tests based on the use of a monthly GARCH model to estimate monthly conditional variances.

In the previous section we assumed that the conditional variance in the next month depends on the average (daily) sample variance in the current month only. Alternatively, one can assume that the conditional variance depends on a weighted average of past volatility. A popular way to model such a dependence is to use the GARCH model proposed by Bollerslev (1986). An excellent survey of the many applications of GARCH and related models to financial data is provided by Bollerslev et al. (1990). The assumption underlying the simplest GARCH model, GARCH(1,1), for the changes in futures prices are

$$\Delta f_t = \xi_t h_t \qquad \xi_t \sim \text{i.i.d., with } E\,\xi_t = 0, \quad E\,\xi_t^2 = 1, \tag{8}$$

$$h_t^2 = \psi + \beta h_{t-1}^2 + \alpha (\Delta f_{t-1})^2, \tag{9}$$

where ψ, α, $\beta > 0$. Although this model is at variance with (6) unless $\delta = 0$, as it implies that $E_t[\Delta f_{t+1}] = 0$, one can hope that it yields sufficiently reliable variance estimates if δ is small in absolute value. If ξ_t is normal, the log likelihood function for the sample $\Delta f_1, \ldots, \Delta f_T$ becomes, apart from initial conditions and a normalizing constant,

$$\text{LogLik}(\psi,\alpha,\beta) = -0.5 \sum_{t=1}^{T} \log h_t^2 - 0.5 \sum_{t=1}^{T} (\Delta f_t)^2/h_t^2. \tag{10}$$

This likelihood can be maximized numerically, using the recursion in (9) to compute subsequent values of h_t^2. It has been shown in Weiss (1986) that the value of (ψ,α,β) which maximizes (10) is, under regularity conditions, a consistent estimator of the parameters in (9) even if the normality assumption is invalid. Moreover Weiss (1986) has shown how consistent estimates of the large sample variance of this pseudo maximum likelihood estimator can be obtained.

The numerical results for Sugar futures yield

$$h_t^2 = \underset{\substack{(3.24)\\ [3.18]}}{2323} + \underset{\substack{(8.09)\\ [9.80]}}{0.52}\, h_{t-1}^2 + \underset{\substack{(6.22)\\ [7.53]}}{0.44}\, (\Delta f_{t-1})^2 \tag{11}$$

where the numbers in parentheses are t-statistics derived from estimates of the large sample variance which are routinely computed from the Hessian of the log likelihood while the numbers in square brackets are the t-statistics proposed by Weiss (1986) which are robust to departures from normality in the rescaled innovations ξ_t. The estimated kurtosis of the rescaled innovations is $\kappa = 4.8$. The estimates in (11) suggest that the conditional variance depends on past volatility in a limited number of months, which is however clearly larger than one. The estimates of the conditional variance which are generated by (11) are presented in figure 5, together with the estimates derived from daily data which were used in the previous section. Although the two series are derived under different assumptions, the results are strikingly similar.

The variance estimates which are generated by (11) have been used to compute GLS estimates of equation (6), using the transformed regression in (7). As we know from figure 5 that the present proxy for the conditional variance is close to the one used in the previous section, it is not too surprising that once more we do not find evidence in Table 2 of a significant impact of the conditional variance on the futures price bias. Moreover the residuals in the present regression are still conditionally heteroskedastic, just as in the previous section.

6. Tests based on a monthly GARCH-M model

As stated in the previous section, the approach which is used there is valid if the impact of the non-zero conditional mean is negligible in the computation of the conditional variance only. Although the estimates presented above do not yield any evidence of a non-zero conditional mean, we evidently can not claim that in fact δ in (6) equals zero, as the impact of a non-zero mean on the properties of the estimates in Table 2 is unclear. The problem can be solved by estimating the mean and variance parameters simultaneously, as proposed by Engle, Lillien and Robins (1987) who refer to this model as the GARCH-M (GARCH in mean) model. The complete model reads as

$$\Delta f_t = \delta h_t^2 + \xi_t h_t \tag{8'}$$

$$h_t^2 = \psi + \beta h_{t-1}^2 + \alpha \xi_{t-1}^2 h_{t-1}^2, \tag{9'}$$

where the properties of the ξ_t's are as in (8) above. The parameters in the GARCH-M model can be estimated by (pseudo) maximum likehood, using a slight generalization of (10).

Estimates of (8')-(9') from monthly data are presented in Table 3. The order of magnitude of the estimates of δ does not differ from earlier results, but the important point to note here is that for both the full sample (January 1972 - June 1989) and the high volatility subsample (January 1972 - December 1981) the impact of the conditional variance on the futures price bias appears to be significant. However signs of significant autocorrelation in the disturbances are present which might result in biased t-ratio's. Although the variance parameters in the GARCH and GARCH-M model almost coincide, the implied conditional variance series differ as in the first model the price changes themselves, without correction for the non-zero mean, drive the variance process. Attempts to estimate GARCH and GARCH-M models for January 1982 - June 1989 failed as the assumption of no persistence in variance ($\alpha + \beta < 1$) which is imposed by the software was violated, which probably has to do with erratic observations in the beginning of the sample and a small sample size. If the sample is chosen as January 1984 - June 1986 convergence is achieved. As in this case the estimated variance parameters are insignificant, it is not surprising to find an insignificant estimate of the mean parameter as well. In order to show that the evidence on the time varying risk premium is not caused by the presence of a constant risk premium and absence of a constant in (9'), we have also estimated the model including the possibility of a time-invariant risk premium, denoted by μ. As shown in the final row of Table 3 this parameter is insignificant and does not affect the significance of the remaining parameters.

7. Tests based on a daily GARCH-M model

In the previous section the GARCH-M model, which was motivated by the pricing model presented in section 2, was estimated from monthly data. It is not clear from the pricing model, however, how long the appropriate time periods are. If the agents solve a simple single period optimization problem, as they do in section 2, the use of monthly data already appears to imply a short planning horizon. On the other hand, however, the agents obviously have access to high frequency information on the underlying variables and have incentives to use this information. This argument suggests the use of high frequency, e.g. daily, data. Moreover, the behavior of agents who in fact maximize a multiperiod additively separable mean variance criterion function, can probably be closely mimicked by the behaviour of myopic agents if subsequent changes in asset prices are uncorrelated, which is roughly the case for the Sugar futures (compare also Ingersoll (1987), p.255-258). This implies that the GARCH-M model might well be a valid specification at the daily frequency as well.

Estimates of the daily GARCH-M model are presented in Table 4. Because of limitations of our software, daily observations have been analyzed from January 1, 1978 onwards only. The estimates for the full sample imply (once more) a significant negative impact of the conditional variance on the futures price bias. Moreover the estimate of δ that is obtained is close to the estimates from monthly data. As suggested by the results on temporal aggregation of GARCH processes in Drost and Nijman (1990) the daily model is close to integration in variance and the estimate of β from daily data exceeds the estimates from monthly data, while the opposite is true for the estimates of α. For the sample from 820101 - 890601 similar results are obtained, although the estimate of δ is somewhat small. For the full sample the problem of multicollinearity between a time-variant and a time-invariant component of the risk premium reappears. For the subsample, however, the significance of the conditional variance term is not affected by the presence of a constant in (9').

8. Concluding remarks

In this paper we have tested the empirical implications for the marginal process of prices of Sugar futures of a simple pricing model. No impact of the conditional variance on the change in futures prices was detected from ad hoc estimates of the conditional variance from daily data or from a GARCH model. Possible explanations of this empirical finding are misspecification of the simple pricing model, misspecification of the conditional variance estimates and a net hedging pressure which is close to zero in the sample. Misspecification of the conditional variance appears to be the correct explanation as significant parameters are found in monthly as well as daily GARCH-M models. These estimates imply contango in the futures markets and a net hedging demand on the long side of the futures market. Moreover our results suggest that the simple pricing model points at at least one important aspect of the pricing of Sugar futures. Tests of the more detailed pricing models which have been proposed in the literature are left for future research.

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Table 1: Dependence of the futures price bias on the estimate of the monthly conditional variance from daily data

sample period		estimation method	cond. variance coeff.	t	h.c.-t	LM_A	LM_{CH}	κ
72-2	89-5	OLS	-0.00023	-1.61	-0.53	15.9	73.8	12.2
72-2	89-5	GLS	-0.00009	-0.21	-0.25	13.5	43.2	5.2
72-2	81-12	GLS	0.00003	0.07	0.08	12.9	31.4	4.5
82-1	89-5	GLS	-0.00120	-0.93	-1.33	5.0	13.2	7.3

Table 2: Dependence of the futures price bias on the estimate of monthly conditional variance from the GARCH(1,1) model

sample period		estimation method	cond. variance coeff.	t	h.c.-t	LM_A	LM_{CH}	κ
72-2	89-5	GLS	-0.00006	-0.20	-0.23	11.7	56.9	7.8
72-2	81-12	GLS	0.00002	0.06	0.08	10.5	33.9	6.9
82-1	89-5	GLS	-0.00089	-1.13	-1.31	6.8	19.8	4.3

Table 3 Monthly GARCH-M models

sample	ψ	α	β	δ	μ	LM_A	κ
72-1 89-5	2324	0.43	0.52	-0.00055		16.8	4.5
	(3.21)	(5.73)	(7.44)	(-3.62)			
	[3.37]	[7.17]	[10.05]	[-3.63]			
72-1 81-12	2358	0.44	0.51	-0.00054		18.9	6.0
	(2.38)	(5.88)	(7.24)	(-3.13)			
	[2.12]	[8.55]	[10.41]	[-2.85]			
82-1 89-5	2789	0.12	0.60	-0.00064		5.0	3.6
	(0.64)	(0.55)	(1.01)	(-0.51)			
	[0.37]	[(0.25]	[0.53]	[-0.48]			
72-1 89-5	2320	0.43	0.52	-0.00053	-0.0043	17.0	4.5
	(3.22)	(5.74)	(7.45)	(-3.04)	(-0.50)		
	[3.26]	[6.99]	[9.90]	[-2.87]	[-0.45]		

Table 4 Daily GARCH-M models

sample	ψ	α	β	δ	μ	LM_A	κ
780101	6.14	0.084	0.906	-0.0008		8.6	4.5
--	(4.22)	(9.75)	(96.9)	(-2.10)			
890601	[3.02]	[6.60]	[65.1]	[-2.02]			
820101	10.41	0.081	0.900	-0.0025		3.6	4.7
--	(3.36)	(7.27)	(63.5)	(-2.70)			
890601	[2.21]	[4.60]	[41.0]	[-2.53]			
780101	6.07	0.085	0.906	-0.0005	-0.72	8.5	4.5
--	(4.20)	(9.75)	(96.8)	(-1.01)	(-1.74)		
890601	[3.00]	[6.63]	[65.3]	[-0.91]	[-1.70]		
820101	10.44	0.081	0.900	-0.0026	0.059	3.6	4.7
--	(3.30)	(7.25)	(62.8)	(-2.12)	(0.14)		
890601	[2.23]	[4.53]	[39.8]	[-2.08]	[0.21]		

FIGURE 1: changes in monthly Sugar future prices

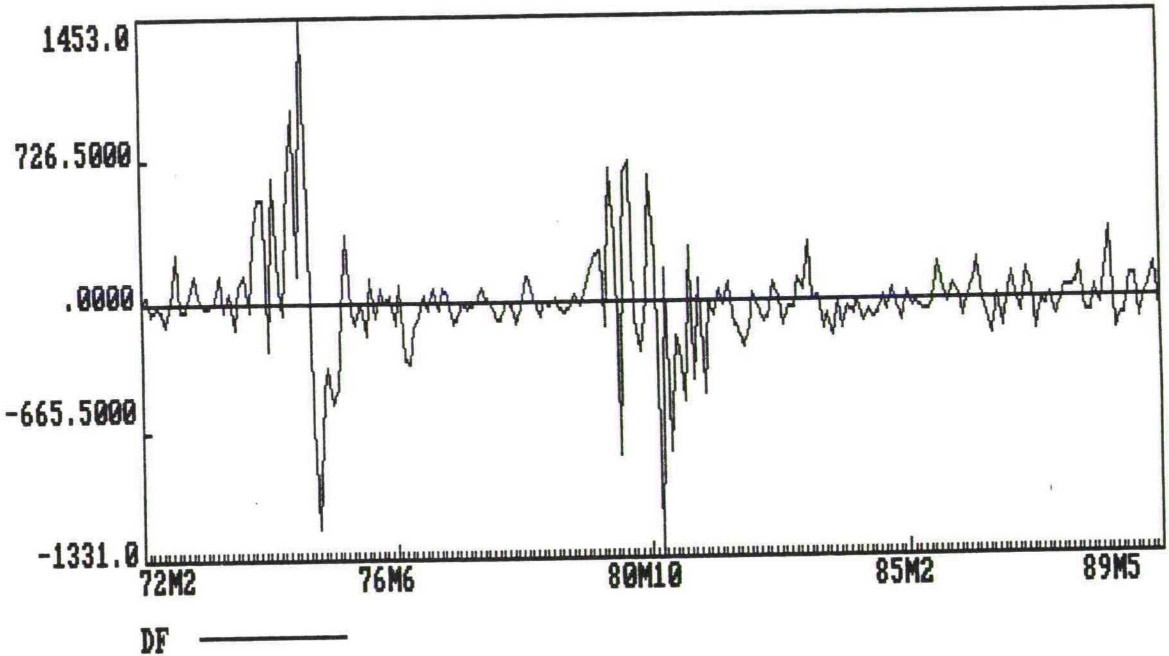
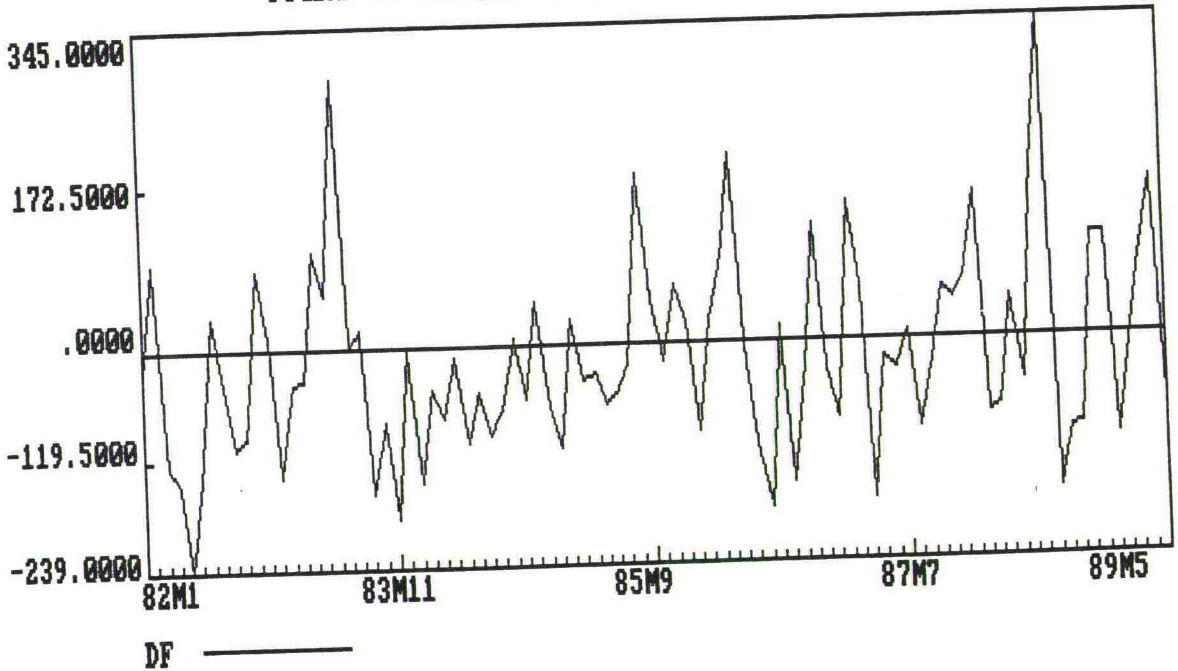
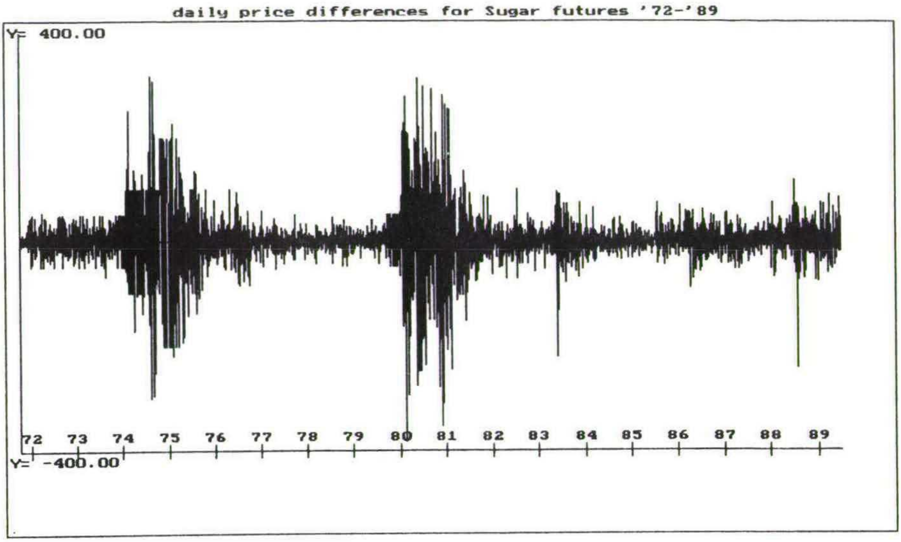


FIGURE 2: changes in monthly Sugar futures prices




daily price differences for Sugar futures '72-'89
Y= 400.00
Y= -400.00
72 73 74 75 76 77 78 79 80 81 82 83 84 85 86 87 88 89

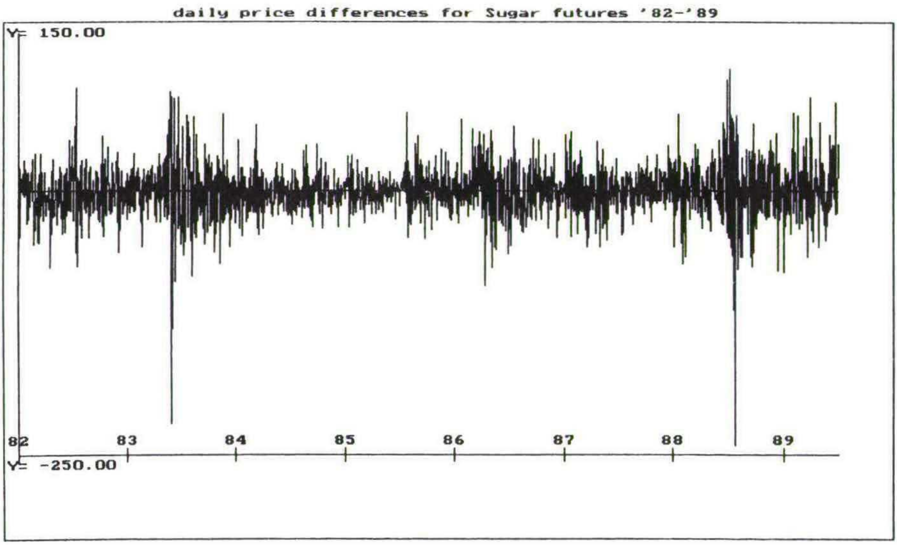
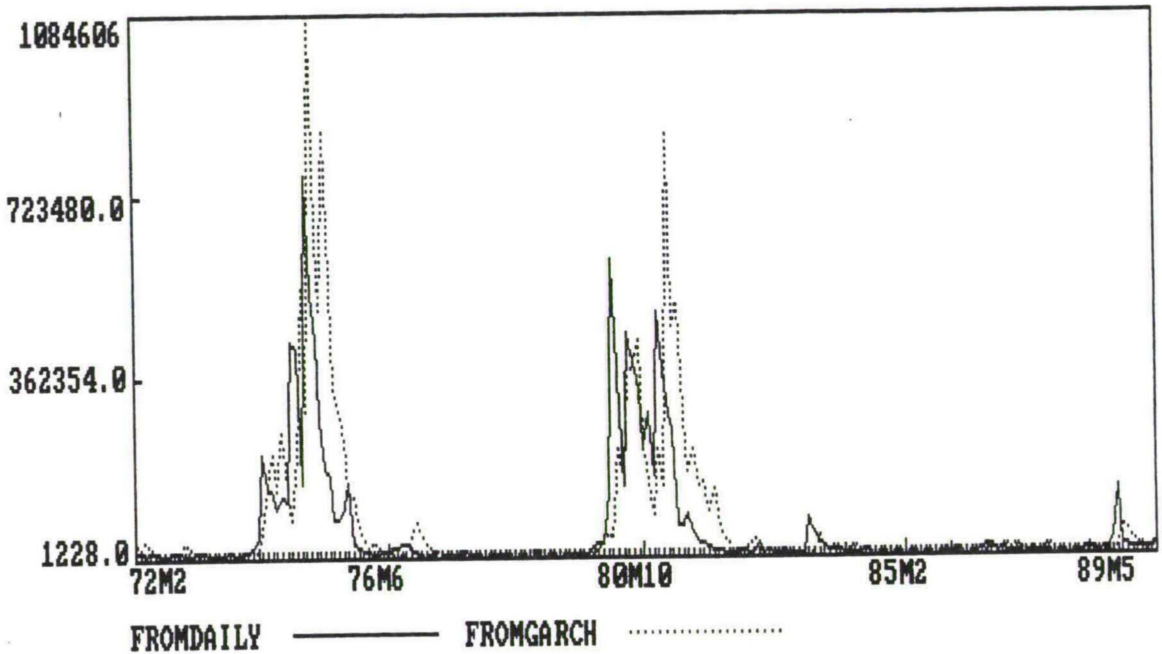

daily price differences for Sugar futures '82-'89
Y= 150.00
Y= -250.00
82 83 84 85 86 87 88 89

FIGURE 5: Variance estimates from daily data and from GARCH model



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