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Modelling the Impact of Joining the WTO on the Volatility of Hog Prices in Taiwan

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Abstract : The hog industry, where prices are determined according to an auction system, is of vital importance to the agricultural industry in Taiwan by providing significant production and employment. In particular, there were significant impacts on daily hog prices in the periods before, during and after joining the WTO, which we will refer to as periods of anticipation, adjustment and settlement. The purpose of the paper is to model the growth rates and volatility in daily hog prices in Taiwan from 23 March 1999 to 30 June 2007, which enables an analysis of the effects of joining the WTO. The paper provides a novel application of financial volatility models to agricultural finance. The empirical results have significant implications for risk management and policy considerations in the agricultural industry in Taiwan, especially when significant structural changes, such as joining the WTO, are concerned. The three sub-samples relating to the period before, during and after joining the WTO display significantly different volatility persistence, namely symmetry, asymmetry but not leverage, and leverage, respectively, whereby negative shocks increase volatility but positive shocks of a similar magnitude decrease volatility.

1. Introduction

Time-varying volatility in agricultural commodity prices, such as hog prices, usually accompanies riskiness in the rates of growth (or returns). How to capture the pattern or characteristics of volatility is of concern to farmers. Under the World Trade Organization (WTO) regulations, direct price support programs of agricultural authorities have had to be progressively eliminated, so that farm prices are essentially determined by the market. Therefore, the volatility associated with prices imposes significant pressures on agricultural producers.

Price changes are associated with volatility and risk. If agricultural commodity prices have predictable time-varying volatility, they can be analysed using recently developed financial econometric methods that incorporate important aspects of optimal portfolio management. McAleer (2009) explains why time-varying volatility can be useful in areas such as environmental finance and tourism finance. Similar arguments can be used for applications in agricultural finance. Volatility from high frequency data can be aggregated, whereas aggregated data at low frequencies typically display no volatility, thereby enabling the prediction of risk associated with the imposition of agricultural taxes. Dynamic confidence intervals can also be computed. Moreover, modelling volatility permits an analysis of the asymmetric and leveraged responses of prices and associated commodity inflation rates to positive and negative shocks of equal magnitude. In this way, commodity prices behave like financial stock prices, so that the theory of finance can be applied directly to agricultural commodity prices.

In order to assist farmers to predict the volatility in prices, several related issues need to be investigated. This paper focuses on the asymmetric response of volatility to positive and negative shocks to prices and returns because the stochastic property of agricultural commodity prices might not be symmetric. The empirical models examined in this paper will evaluate hog prices, their returns and associated volatility in Taiwan. For further details, see Chang et al. (2009).

2. Data

The data set comprises daily hog prices in Taiwan from 23 March 1999 to 30 June 2007, giving a total of 2,024 observations. The data were obtained from the website of the National Animal Industry Foundation (NAIF) in Taiwan. There has been a large decrease in daily hog prices, as well as in the logarithm of daily hog prices, during the period 23 June 1999 to 27 December 2001, a large increase in prices and log prices during the period 28 December 2001 to 6 August 2004, and then a significant reduction in prices and log prices during the period 7 August 2004 to 30 June 2007.

As described above, there would seem to be significant increasing and decreasing trends in the daily hog price and logarithmic hog price throughout the sample period. These variations in daily hog prices are likely to have been caused by Taiwan's decision to join the World Trade Organization (WTO) in 2002, whereby trade liberalization led to strikes in the domestic hog market in Taiwan. Prominently, the three sub-samples described above correspond to the three stages in terms of Taiwan joining the WTO. For this reason, we will interpret sub-sample 1 as the "planning" stage of Taiwan joining the WTO, sub-sample 2 as the "adjustment" period immediately after Taiwan joined the WTO, and sub-sample 3 as the "settlement" stage.

3. Unit Root Tests

It is well known that the traditional ADF and PP unit root tests suffer from low power and size distortions. However, these shortcomings have been overcome by modifications to the testing procedures. The modified unit root tests given by $MADF^{GLS}$ and MPP^{GLS} were applied to the time series of daily hog prices in Taiwan. In essence, these tests use GLS de-trended data and the modified Akaike information criterion (MAIC) to select the optimal truncation lag. The asymptotic critical values for both tests are given in Ng and Perron (2001).

The results of the unit root tests are obtained from the econometric software package EViews 5.0. The null hypothesis of a unit root is not rejected for the levels of daily hog prices in the models with a constant and with a constant and trend as the deterministic terms. A similar result holds for the logarithms of daily hog prices, where both the MADF^{GLS} and MPP^{GLS} tests do not reject the null hypothesis of a unit root for the models with a constant and with a constant and trend. However, for the series in log differences (or growth rates), the null hypothesis of a unit root is rejected for both specifications using both the MADF^{GLS} and MPP^{GLS} tests. Overall, the null hypothesis of a unit root is not rejected for the levels or logarithms of daily hog prices, but is rejected for the growth rate of daily hog prices. Similar results of the unit root tests are found in each of the three sub-samples.

4. Conditional Mean and Conditional Volatility Models

The alternative time series models to be estimated for the conditional means of the daily hog prices, as well as their respective conditional volatilities, are discussed below. Daily hog prices and the logarithm of daily hog prices do not show persistence in volatility, whereas the first differences (that is, the log difference or growth rate) of daily hog prices in Taiwan show periods of persistent high volatility from 23 June 1999 to 27 December 2001, followed by relatively low volatility from 23 June 1999 to 27 December 2001, and then by relatively high volatility from 28 December 2001 to 6 August 2004. One implication of this persistent time-varying volatility is that the assumption of conditionally homoskedastic residuals would seem to be inappropriate for sensible empirical analysis.

For a wide range of financial data series, time-varying conditional variances can be explained empirically through the autoregressive conditional heteroskedasticity (ARCH) model of Engle (1982). When the time-varying conditional variance has both autoregressive and moving average components, this leads to the generalized ARCH(*p*,*q*), or GARCH(*p*,*q*), model of Bollerslev (1986). The lag structure of the appropriate GARCH model can be chosen by information criteria, such as those of Akaike and Schwarz, although it is very common to impose the widely estimated GARCH(1,1) specification in advance.

In the selected conditional volatility model, the residual series should follow a white noise process. Li et al. (2002) provide an extensive review of recent theoretical results for univariate and multivariate time series models with conditional volatility errors. McAleer (2005) reviews a wide range of univariate and multivariate, conditional and stochastic, models of financial volatility. McAleer et al. (2007) discuss recent developments in modeling univariate asymmetric volatility, while McAleer et al. (2008) develop the regularity conditions and establish the asymptotic properties of a general model of time-varying conditional correlations.

Consider the stationary AR(1)-GARCH(1,1) model for daily hog prices in Taiwan (or their growth rates, as appropriate), y_t:

$$y_t = \phi_1 + \phi_2 y_{t-1} + \varepsilon_t, \qquad |\phi_2| < 1 \tag{1}$$

for $t = 1,...,n$, where the shocks (or movements in daily hog prices) are given by:

$$\begin{aligned} \varepsilon_t &= \eta_t \sqrt{h_t}, \quad \eta_t \sim iid(0,1) \\ h_t &= \omega + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1}, \end{aligned} \tag{2}$$

and $\omega > 0, \alpha \geq 0, \beta \geq 0$ are sufficient conditions to ensure that the conditional variance $h_t > 0$. The AR(1) model in equation (1) can easily be extended to univariate or multivariate ARMA(*p*,*q*) processes (for further details, see Ling and McAleer (2003a)). In equation (2), the ARCH (or α) effect indicates the short run persistence of shocks, while the GARCH (or β) effect indicates the contribution of shocks to long run persistence (namely, $\alpha + \beta$).

In equations (1) and (2), the parameters are typically estimated by the maximum likelihood method to obtain Quasi-Maximum Likelihood Estimators (QMLE) in the absence of normality of η_t, the conditional shocks (or standardized residuals).

As the GARCH process in equation (2) is a function of the unconditional shocks, the moments of ε_t need to be investigated. Ling and McAleer (2003) showed that the QMLE for GARCH(*p*,*q*) is consistent if the second moment of ε_t is finite. The well known necessary and sufficient condition for the existence of the second moment of ε_t for GARCH(1,1) is $\alpha + \beta < 1$.

As discussed in McAleer et al. (2007), Elie and Jeantheau (1995) and Jeantheau (1998) established that the log-moment condition was sufficient for consistency of the QMLE of a univariate GARCH(*p*,*q*) process. For GARCH(1,1), this is given as:

$$E(\log(\alpha\eta_t^2 + \beta)) < 0. \qquad (3)$$

The effects of positive shocks (or upward movements in daily hog prices) on the conditional variance, h_t, are assumed to be the same as negative shocks (or downward movements in daily hog prices) of a similar magnitude in the symmetric GARCH model. In order to accommodate asymmetric behaviour, Glosten, Jagannathan and Runkle (1992) proposed the GJR model, for which GJR(1,1) is defined as follows:

$$h_t = \omega + (\alpha + \gamma I(\eta_{t-1}))\varepsilon_{t-1}^2 + \beta h_{t-1}, \qquad (4)$$

where $\omega > 0, \alpha \ge 0, \alpha + \gamma \ge 0, \beta \ge 0$ are sufficient conditions for $h_t > 0,$ and $I(\eta_t)$ is an indicator variable that is defined by:

$$I(\eta_t) = \begin{cases} 1, & \varepsilon_t < 0 \\ 0, & \varepsilon_t \ge 0 \end{cases}$$

as η_t has the same sign as ε_t. The indicator variable differentiates between positive and negative shocks of equal magnitude, so that asymmetric effects in the data are captured by the coefficient γ. For financial data, it is typically expected that $\gamma \ge 0$ because negative shocks increase risk by increasing the debt to equity ratio, but this interpretation need not hold for hog price data in the absence of a similar interpretation in terms of risk. The asymmetric effect, γ, measures the contribution of shocks to both short run persistence, $\alpha + \frac{\gamma}{2}$, and to long run persistence, $\alpha + \beta + \frac{\gamma}{2}$. Ling and McAleer (2002a) showed that the regularity condition for the existence of the second moment for GJR(1,1) under symmetry of η_t is given by:

$$\alpha + \beta + \frac{1}{2}\gamma < 1, \qquad (5)$$

while McAleer et al. (2007) showed that the weaker log-moment condition for GJR(1,1) was given by:

$$E(\ln[(\alpha + \gamma I(\eta_t))\eta_t^2 + \beta]) < 0, \qquad (6)$$

which involves the expectation of a function of a random variable and unknown parameters.

An alternative model to capture asymmetric behaviour in the conditional variance is the Exponential GARCH (EGARCH(1,1)) model of Nelson (1991), namely:

$$\log h_t = \omega + \alpha |\eta_{t-1}| + \gamma \eta_{t-1} + \beta \log h_{t-1}, \qquad |\beta| < 1 \qquad (7)$$

where the parameters α, β and γ have different interpretations from those in the GARCH(1,1) and GJR(1,1) models (see also Ling and McAleer (2002b)).

5. Estimated Models

It is well known that the estimates of volatility will depend on the adequacy of the specification of the conditional mean equation, which yields the standardized residuals. A related issue is the effect of ignoring structural change in the conditional mean of the estimates of the conditional variance. The effects of misspecifying the conditional mean on the estimates of the conditional volatility will be analyzed below. Both the asymptotic standard errors, as well as the robust standard errors of Bollerslev and Wooldridge (1992), are presented. In virtually all cases, the asymptotic standard errors are smaller than their robust counterparts.

The estimated conditional mean and conditional volatility models are given in Table 1. As shown in the unit root tests, the levels and logarithms of daily hog prices are not stationary, but the log differences (or growth rates) are stationary. For this reason, only the growth rates and their associated volatility will be modelled for the full sample period, which is given in Table 1. Although not reported here for reasons of space limitation, the estimates for the three sub-samples are available on request from the authors.

As the second moment condition is less than unity in all cases, it follows that the weaker log-moment condition is less than zero in all cases (see Tables 2-4). Thus, the regularity conditions are satisfied, the QMLE are consistent and asymptotically normal, and inferences are valid. The EGARCH(1,1) model is based on the standardized residuals, so the regularity condition is satisfied if $|\beta| < 1$, and hence the QMLE would seem to be consistent and asymptotically normal (see, for example, McAleer at al. (2007)).

The trends and volatilities in daily hog prices (as well as in their logarithms) seem to have experienced two noticeable structural changes during the sample period. As has already been mentioned, these changes would seem to have arisen from Taiwan's decision to join the World Trade Organization (WTO) in 2002, when the national government delegated efforts to protect domestic hog producers against lower imported hog prices. These actions may have altered the trends in the hog prices, as well as their associated volatility, in those periods. In order to investigate the effects of joining the WTO, alternative models of the growth rates and their associated volatility were estimated for three sub-samples before, during and after Taiwan joined the WTO, namely sub-sample 1 from 23 June 1999 to 27 December 2001, sub-sample 2 from 28 December 2001 to 6 August 2004, and sub-sample 3 from 7 August 2004 to 30 June 2007, respectively. The estimates for the three sub-sample periods are given separately in Tables 2-4.

Although not reported here, in a comparison of the estimates for the three sub-samples, it may be concluded that the long run persistence of shocks in the planning period (that is, sub-sample 1) suggests a more significant impact on the conditional variance than in the other two sub-samples. In addition, the short run persistence of shocks in the settlement period (that is, sub-sample 3) suggests the weakest impact on the conditional variance among the three sub-samples. Overall, the asymmetric effect is found to be significant for the GJR (1,1) model in two of the three sub-samples, while the effect of negative shocks has tended to increase over the full sample period. However, in the asymmetric EGARCH(1,1) model, a leverage effect is observed only in the settlement period (that is, sub-sample 3), whereas it is not significant in sub-period 1 and is significant, but does not suggest the existence of leverage effects, in sub-sample 2. In summary, the adjustment period (that is, sub-sample 2) implies a dominant contributing role for estimating the impact of the short run and long run persistence of shocks for the full sample period.

In general, the QMLE for the GARCH(1,1), GJR(1,1) and EGARCH(1,1) models for the log differences (or growth rate) in daily hog prices in Taiwan are statistically adequate and have sensible interpretations. For the full sample period, in which any structural changes are ignored, there is asymmetry in volatility for the GJR(1,1) and EGARCH(1,1) models, but there is no presence of leverage effects, whereby negative shocks increase volatility but positive shocks of a similar magnitude decrease volatility. The three sub-samples exhibit different types of symmetry or asymmetry, with the period prior to joining the WTO showing symmetry, the period of joining displaying asymmetry but not leverage, and the period after joining indicating leverage. This enables an empirical analysis of the effects on the prices of the hog production industry of joining the WTO by Taiwan, whereby hog prices behave very much like financial commodity prices.

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Table 1. Conditional Mean and Volatility Models for Log Difference in Prices for the Full Sample, 23 March 1999 – 30 June 2007

Parameters	Dependent variable: DLY		
	GARCH	GJR	EGARCH
ϕ_1	0.0004	-0.0004	-0.0005
	(0.0004)	(0.0004)	(0.0004)
	[0.0005]	[0.0004]	[0.0004]
ϕ_2	-0.143***	-0.103***	-0.101***
	(0.028)	(0.026)	(0.025)
	[0.039]	[0.039]	[0.036]
ω	0.0001***	0.0002***	-2.907***
	(0.00001)	(0.00001)	(0.258)
	[0.00003]	[0.00003]	[0.628]
GARCH/GJR α	0.274***	0.104***	
	(0.016)	(0.019)	--
	[0.106]	[0.068]	
GARCH/GJR β	0.425***	0.365***	
	(0.034)	(0.030)	--
	[0.102]	[0.101]	
GJR γ		0.434***	
	--	(0.038)	--
		[0.200]**	
EGARCH α			0.423***
	--	--	(0.023)
			(0.077)
EGARCH γ			-0.148***
	--	--	(0.016)
			[0.070]**
EGARCH β			0.668***
	--	--	(0.032)
			[0.078]
Diagnostics			
Second moment	0.699	0.686	--
Log-moment	-0.231	-0.264	--

Notes:
DLY is Log Difference in Hog Price (New Taiwan Dollars per kilogram)
Numbers in parentheses are asymptotic standard errors, while numbers in brackets are the Bollerslev and Wooldridge robust standard errors.
** and *** denote significance at the 5% and 1% level, respectively.