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INFLATION AND MEASUREMENT OF HISTORICAL RISK

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Abstract

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Previous studies of objective risk measures abstracted from effects of inflation. This study calculates the variance of farm milk prices for 1960-90, using feasible generalized least squares to detrend nominal and real price series. Results support the use of nominal data but indicate the need for additional research.

INFLATION AND MEASUREMENT OF HISTORICAL RISK

Objective risk measures are utilized extensively in agricultural economics to provide information about risk in historical production and market environments both to farm managers (Carter and Dean; Mathia; Walker and Lin) and policy makers (Hazell; Miranda). These measures are also important parameters in risk programming and other simulation models for farm management (Mapp and Helmers) and policy analysis (Miranda and Glauber). Objective risk measures are statistical measures such as variances, covariances, and mean absolute deviations that are calculated from historical data. Considerable agricultural economics literature exists on evaluation of methods used to calculate historical risk measures (Young, 1984). Choices of estimators and detrending methods has been the focus of much past research (Adams, Menkhaus, and Woolery; Fackler and Young; Kramer, McSweeney, and Stavros; Swinton and King; Young, 1984).

The effect of inflation on historical risk measurement has not been extensively investigated. Young (1980) found that both nominal and real data were used to calculate risk measures, but the issue of inflation was not evaluated. From a conceptual viewpoint, Brake, Levy and Sarnat, and White and Musser suggested that unexpected changes in inflation would increase risk. Thus, real data may have a source of risk removed that would not have been removed with trend analysis of nominal data. Using the variance of a detrended data series as a measure of risk, this view supports the hypothesis that the variance of real data will be smaller than the variance of nominal data. In addition, Barnett, Bessler, and Thompson suggested that the use of deflated price series will result in biased coefficient estimates. Estimates of variance using deflated data may also be biased.

The purpose of this paper is to investigate the use of nominal versus real data in calculating historical risk measures for quarterly U.S. milk prices received by farmers for the 1960-1990 period, using the variance of a detrended data series as a measure of risk. A perception of increasing price risk over time resulting from structural changes affecting the dairy sector during the 1960-90 period (Fraher; Hamm) presents an opportunity to evaluate the consequences of using nominal versus real data in quantifying this perception. Two hypotheses form the framework for this analysis: (1) the variance of the detrended real price series will be smaller than the variance of the detrended nominal price series, and (2) variance will be increasing over time. Following Fackler and Young, the analysis also considers the use of feasible generalized least squares methods to detrend historical data rather than the traditional approach of ordinary least squares.

Data and Methods

The milk price series used in this analysis is for the thirty-one-year period from 1960 to 1990. The original data are monthly observations of prices received by farmers measured in dollars per hundredweight (USDA, NASS). Quarterly prices were calculated as averages weighted by monthly milk production. Real prices (1990 dollars) were calculated with the Index of Prices Paid by Producers for Production Items, Interest, Taxes, and Wage Rates (USDA, NASS). The deflator was chosen so that real prices would reflect constant purchasing power for all farm inputs. The inflation rates for input and output prices may differ substantially, so increases in milk prices can result in decreased real profits if production costs are inflating more rapidly than output prices. Based on this logic, the Index of Prices Paid by

Producers for Production Items, Interest, Taxes, and Wage Rates has been used to deflate prices in econometric studies measuring crop supply response to risk (Lin; Traill).

The nominal and real price series are illustrated in Figures 1 and 2, respectively. The observed nominal milk prices for 1960-90 appear consistent with the hypothesis that price risk has been increasing over this time period. Based on the data patterns in Figure 1 and historical economic and policy environments, three sub-periods were delineated to examine the hypothesis of increasing risk: 1960-72, 1973-80, and 1981-90. The 1960-72 sub-period was characterized by relatively stable economic and policy conditions. Higher energy prices and increased export demand (Musser, Mapp, and Barry) increased the cost of producing milk in the early-1970's. Rapidly increasing inflation rates and changes in federal policy during the 1972-80 sub-period led to higher support price levels. The 1981-90 sub-period was characterized by significant changes in the structure of dairy policy during the early-1980's, later followed by increased export demand for dairy products and unusual crop weather conditions in 1988-89 which significantly affected the supply of milk (USDA, ERS).

The risk measure used in this analysis is the variance of the detrended milk price series. Following standard procedures, the deterministic component of the series is removed with least squares estimation techniques to isolate the remaining random variation. Figures 1 and 2 suggest that seasonal and trend components make up the deterministic portion of the price series. Thus, the nominal and real price series were regressed on a time variable and quarterly dummies with the first quarter as the base. A quadratic time variable was also included in the real price model since the deflated data appeared to contain a nonlinear trend (Figure 2).

The above standard detrending models are quite naive; thus specification errors are likely to occur which result in biased coefficient estimates. If model specification error is manifested in autocorrelation and heteroskedasticity, coefficient estimates are also inefficient, and their standard errors are biased. While autocorrelation is recognized as a problem in time series data (Judge, et. al.), the hypothesis of increasing risk over time in this analysis also suggests by definition a violation of the usual assumption of homoskedasticity. Epps and Epps concluded that standard tests for heteroskedasticity are invalid in the presence of autocorrelation, whereas tests for autocorrelation are not appreciably affected by heteroskedasticity. Therefore, this research used a sequential process to consider these potential econometric problems. Following their analysis, parameters were estimated with ordinary least squares (OLS), and residuals tested for first-order autocorrelation using the Durbin-Watson statistic. If autocorrelation was present, models were re-estimated using the Yule-Walker method (Gallant and Goebel). Residuals from the Yule-Walker (YW) models were then checked for heteroskedasticity using the Glesjer test, where absolute values of the residuals were regressed on transformed dummy variables for the time periods hypothesized to have increasing variance. Unlike the Breusch-Pagan and Goldfeld-Quandt tests, the Glesjer test yields a specific estimate of the form of heteroskedasticity necessary for feasible generalized least squares estimation (Johnston). If the hypothesis of homoskedasticity was rejected, predicted values from the Glesjer models were used to re-estimate the model with weighted least squares (WLS). Detrended variances were calculated from the OLS, YW, and WLS residuals. For the WLS model, the residuals were rescaled with the predicted values from the Glesjer model in order to recover the unweighted estimated errors. The residuals were then used to estimate variances for 1960-90 and the three sub-periods.

Results

Variances of the original time series of nominal and real prices are presented in Table 1. Not surprisingly, the variance was highest during 1973-80 for nominal prices. Input prices, especially feed, fuel costs, and wage rates were quite volatile during this period due to the energy crisis, changes in the structure of U.S. feed grain markets, and an accelerating level of overall inflation. The effect of inflation on variance in the nominal series is illustrated by the magnitude of the real price variance: real variance was less than one third of the nominal variance in the second sub-period (1973-80). A substantial upward trend contributed to the large variance observed in the nominal data during the second sub-period. This variance would not usually be considered risk. Neither the variance of the real series nor the nominal series was the highest in the third sub-period (1981-90), and the variance of real prices was higher than the variance of nominal prices for 1960-72 and 1981-90. These results are inconsistent with the basic hypotheses of this research. The larger magnitudes of variances for the whole period, especially for the nominal data, are consistent with trends in the data and support the need for detrending to evaluate the research hypotheses.

Results for the trend analyses of nominal data are reported in Table 2. The Durbin-Watson statistic for both OLS models are smaller than the lower bound of significance at the five percent level, indicating positive first-order autocorrelation. The Yule-Walker estimates for nominal and real models have lower standard errors for the seasonal dummy variables than the OLS estimates, a result which is consistent with the presence of autocorrelation in the OLS models. However, standard errors are higher in both YW equations for the time trend variable. Preciseness of the Durbin-Watson values for the YW models may be somewhat affected by the presence of heteroskedasticity. The Durbin-Watson statistic for the YW model of real prices falls within the inconclusive region, while the statistic for the nominal price model is just below the lower bound at the one percent significance level.

Results of the Glesjer regressions are reported in Table 3. The final specification in Table 3 includes dummy variables for the second and third sub-periods. Since Judge, et. al. indicate that pretest error is not a problem with heteroskedasticity tests, this model was adopted after performing pretests of several specifications using variables from the OLS and YW models. The final specification provides some direct evidence on the hypothesis of increasing risk. The sub-period variables were significant in both Glesjer models, supporting the view that heteroskedasticity was present in the YW and OLS models. Positive coefficients on the sub-period variables suggest that risk increased in subsequent time periods. However, the larger coefficient in the second sub-period relative to the third for the real data does not indicate a further increase in risk in the third sub-period for this model. In contrast, the magnitudes of the coefficients for the nominal data are consistent with increasing risk over time.

Regression estimates for the WLS models are reported in Table 2. Standard errors for the quarterly dummy variables in both nominal and real models were slightly reduced compared to the YW models. Standard errors on the time trend variables actually increased in the nominal and real WLS models.

Variances calculated with the regression models are reported in Table 1. The process for correcting heteroskedasticity is *ad hoc*, and empirical results can be quite sensitive to assumptions made about the error term (Johnson, Johnson, and Buse). If heteroskedasticity were accommodated in the WLS model, the variance estimates would be expected to be constant across sub-periods. To further evaluate the hypothesis of increasing risk, the WLS residuals were adjusted with coefficient estimates on the time variables from the Glesjer model. This procedure was adapted from the standard method of adjusting data for time trends (Steel and Torrie):

$$\hat{u}_t^2 = \hat{e}_t^2 + \sum_{i=1}^{2} \hat{b}_i (D_{it} - \overline{D_i}) , \tag{1}$$

where $\hat{e}_t^2$ is the rescaled residual from the WLS model, $\hat{b}$ equals an estimated coefficient on the Yule-Walker transformed dummy variable in the Glesjer model, D_{it} is the dummy variable for time t, $\overline{D}_i$ equals the sample mean of D_{it}, and $\hat{u}_t^2$ is the adjusted residual.

The variances for all models of nominal prices have identical qualitative patterns. Variances are higher in 1973-80 than in 1960-72, and variances in 1981-90 are higher than in 1973-80 and 1960-90. All these estimates support the hypothesis that milk price risk has increased over time. However, the OLS and YW estimates are expected to be biased because of the presence of heteroskedasticity in both models and autocorrelation in the OLS model. In this application, the WLS and adjusted WLS variance estimates do not appear to differ greatly from the YW estimates, suggesting a limited effect of heteroskedasticity on the variance estimates. However, the OLS variances are larger for all periods than the YW and WLS variances, an indication that autocorrelation in the OLS model resulted in biased estimates of variance. The YW and WLS estimates of real variances have patterns similar to the nominal models: variances are consistent with increasing risk and show significant reductions in magnitude in comparison with OLS variances. However, the OLS and adjusted WLS estimates of real variances did not have the same qualitative pattern as observed for the nominal models. The OLS variance is larger for 1960-72 than for 1973-80, while the adjusted WLS variance is larger for 1973-80 than for 1981-90.

The variance results presented above do not support the other basic hypothesis of this study that deflation of data removes a source of risk. The YW and both WLS estimates of variance for detrended real data are greater than the respective estimates for detrended nominal data. While this issue warrants additional research, the simple reasoning behind this hypothesis may be incorrect and/or the index used to deflate the nominal data may be inappropriate for milk prices. Nevertheless, a case can be made that the nominal variance estimates are more consistent with common perceptions concerning evolution of dairy pricing environments than are the real estimates. The WLS variance estimate of nominal data for 1981-90 is approximately 30 times higher than that for 1960-72. In contrast, the 1981-90 WLS estimate of real data is only twice that of 1960-72. Given the stability of milk pricing policy during the 1960's, the variances obtained from the nominal data seem the most plausible.

Conclusions

This research focused on the impact of deflating price data before detrending to calculate variances of milk prices received by farmers. Contrary to expectations, the YW and WLS variances of detrended real prices were greater than for detrended nominal prices. While the conclusions of this analysis still support the use of nominal data, more research on the issue of inflation and historical risk measurement is warranted. The research does support the use of feasible generalized least squares to detrend data since OLS variance estimates are known to be biased under autocorrelation and/or heteroskedasticity. Variances estimated with OLS detrending models were much larger than variances estimated using models that accommodated autocorrelation and heteroskedasticity. The impact of autocorrelation as evidenced by large OLS variances appeared to be more of a problem than heteroskedasticity. However, one would expect to observe heteroskedastic processes if risk is changing over time. Thus, additional evaluation of feasible generalized least squares methods in detrending data is needed.

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Table 2. Regression Estimates of Trends in Quarterly Milk Prices.[a]

	Ordinary Least Squares		Yule-Walker		Weighted Least Squares	
	Nominal	Real	Nominal	Real	Nominal	Real
Intercept	2.80** (.287)	17.12** (.261)	3.53** (.732)	17.16** (.495)	0.18** (.069)	5.05** (.164)
Time	.100** (.003)	0.04** (.008)	0.09** (.010)	0.04* (.018)	0.10** (.012)	0.05** (.020)
Time2		-.001** (.0001)		-.001** (.0001)		-0.001** (.0002)
Quarter2	-0.54 (.309)	-1.03** (.212)	-0.50** (.061)	-1.02** (.108)	-0.47** (.056)	-1.06** (.103)
Quarter3	-0.35 (.309)	-0.38 (.212)	-0.29** (.070)	-0.36** (.124)	-0.25** (.065)	-0.35** (.118)
Quarter4	0.13 (.309)	0.65** (.212)	0.23** (.061)	0.69** (.109)	0.23** (.057)	0.73** (.103)
Lag1			-0.93** (.035)	-0.70** (.066)		
Durbin-Watson	0.10	0.59	1.32	1.62	1.37[b]	1.63[b]
R2	0.90	0.78	0.99	0.89	0.85	0.99

a Standard errors of estimated coefficients are in parentheses.
b Bounds for the DW statistic are reinterpreted for regressions without an intercept (Johnston).
* Denotes significance at the .05 level of confidence.
**Denotes significance at the .01 level of confidence.

Table 1. Milk Price Variances from Original and Detrended Series.

	1960-90	1960-72	1973-80	1981-90
Nominal Prices	14.29	0.57	3.56	0.67
Real Prices	3.21	1.44	1.03	0.87
Detrended Nominal Prices				
Ordinary Least Squares	1.42	0.74	0.98	2.28
Yule-Walker	0.14	0.03	0.10	0.27
Weighted Least Squares	0.13	0.01	0.10	0.28
Adjusted Weighted Least Squares	0.15	0.01	0.14	0.29
Detrended Real Prices				
Ordinary Least Squares	0.63	0.65	0.41	0.73
Yule-Walker	0.42	0.27	0.38	0.63
Weighted Least Squares	0.33	0.23	0.39	0.41
Adjusted Weighted Least Squares	0.35	0.23	0.46	0.41

Table 3. Glesjer Model for Quarterly Milk Prices.[a]

	Nominal Prices	Real Prices
Intercept	0.192** (.028)	0.327** (.049)
Sub-Period2	0.784** (.260)	0.768** (.233)
Sub-Period3	1.104** (.349)	0.425* (.240)

Note: Dependent variable is the absolute value of YW residuals.
a Standard errors of estimated coefficients are in parentheses.
* Denotes significance at the .10 level of confidence.
** Denotes significance at the .01 level of confidence.

Figure 1. U.S. Milk Prices
Received by Farmers, 1960-90

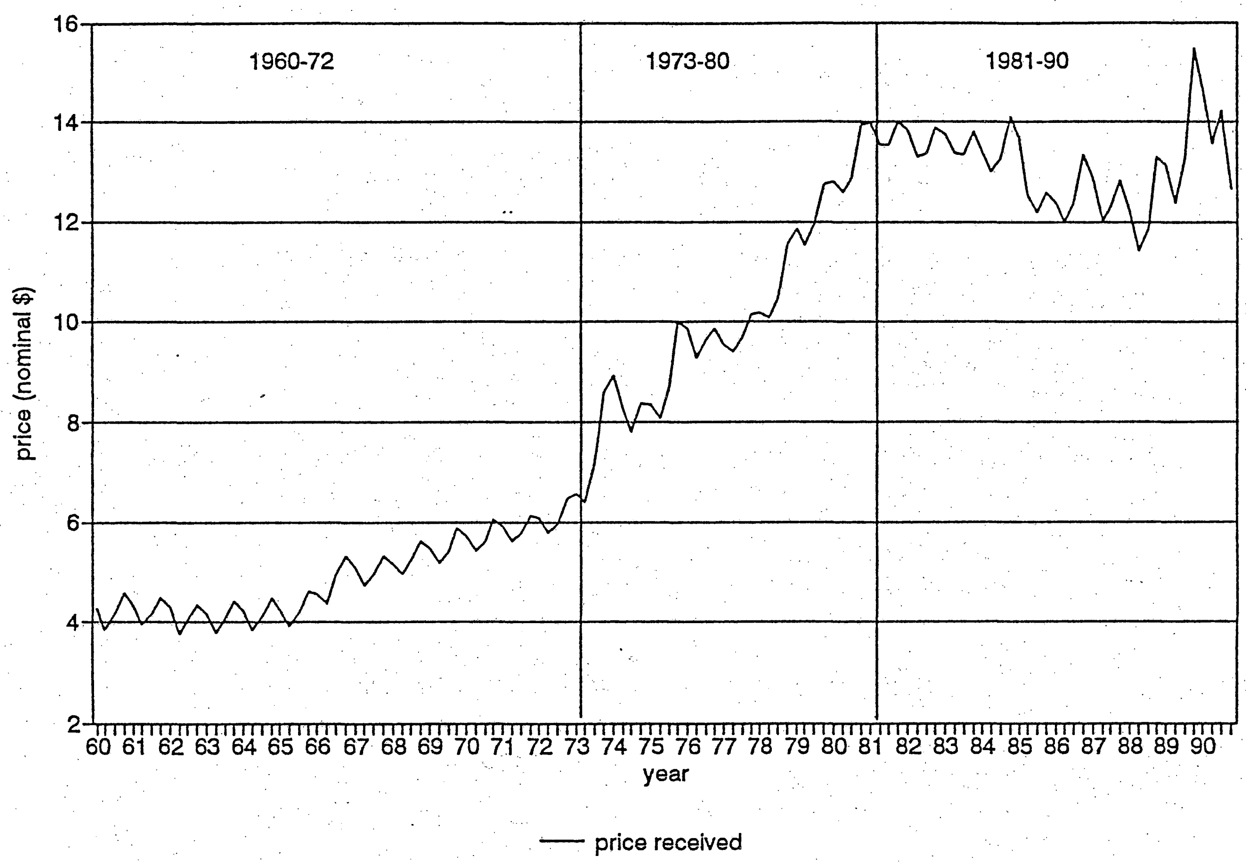

Figure 2. U.S. Real Milk Prices
Received by Farmers, 1960-90