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A Dynamic Analysis of the Effects of a Price Support Program on Price Dynamics and Price Volatility

Kwansoo Kim and Jean-Paul Chavas

This study presents an econometric analysis of the effects of a government price support program on price dynamics and price volatility. Price support programs, a common feature of agricultural policy, provide a lower-bound censoring of the distribution of market prices. An econometric model of market prices is developed using a dynamic Tobit specification under time-varying volatility. The model is applied to the U.S. non-fat dry milk market. It is used to investigate the impact of market liberalization on price dynamics and price volatility in the presence of private and public stocks. The econometric results show how the price support program and stocks (both private and public) affect expected price and price volatility.

Key words: censored regression, market liberalization, price dynamics, price volatility

Introduction

The importance of government intervention in agricultural markets is well documented (e.g., Gardner). Intervention has involved many policy instruments, including import quotas and price floors. Price floors (price support programs) have been a key feature of U.S. agricultural policies since the 1930s. They have been implemented as a means of stabilizing and increasing farm prices, and raising farm income (e.g., Shonkwiler and Maddala; Holt; Holt and Johnson). Price support programs involve government purchase of storable products. In particular, in the U.S. dairy sector, support prices are set for butter, non-fat dry milk, and American cheese. If the market price falls below the support price, then the government purchases dairy products, thus increasing public stocks.

Until the 1990s, U.S. government price support programs were active for major field crops and the dairy sector. The 1990s saw a shift in U.S. agricultural policy toward market liberalization, which has lowered agricultural price support levels for many commodities. The influence of this policy shift on the functioning of agricultural markets remains poorly understood. Given the empirical evidence that most farmers are risk averse (e.g., Lin, Dean, and Moore; Binswanger; Antle; Saha, Shumway, and Talpaz), an understanding of the effects of this recent policy change on price uncertainty should provide insights into the impact of market liberalization. Lowering a support price means reducing the role of government in stock holding, and thus increasing the importance of private stocks.

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Previous empirical research has documented the stock effects on price and price volatility (e.g., Shively). Here, we expand on previous work by considering the impacts of the changing role of private versus public stock holdings under market liberalization.

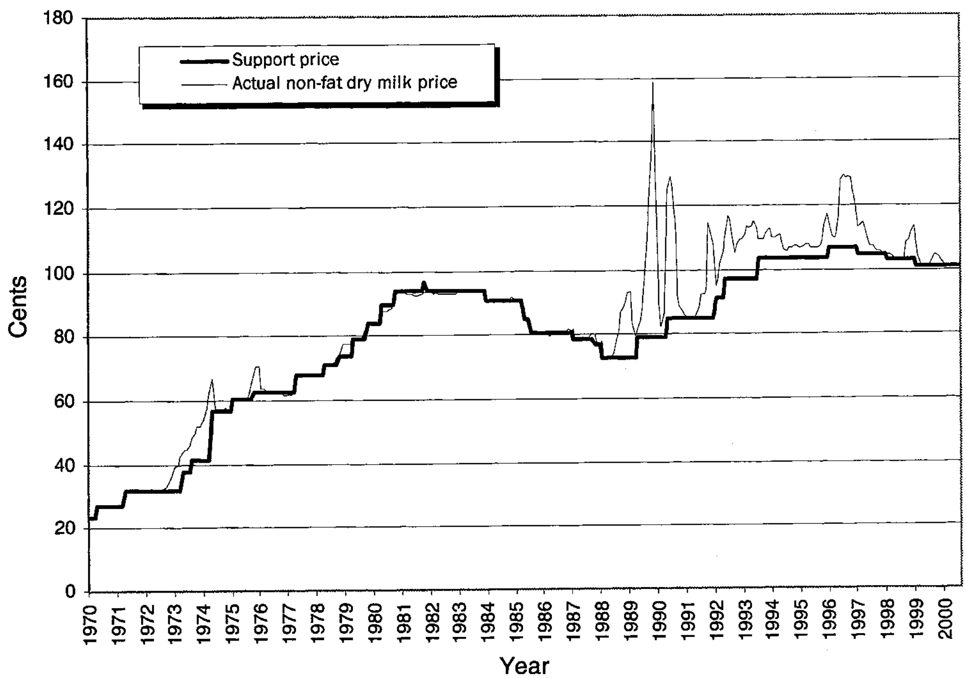
The objective of this study is to develop a model of price dynamics under market liberalization, with an analysis of the effects of price support and stocks on the mean and variance of prices. Methodologically, the framing of this analysis is innovative in several ways. First, a reduced-form model is specified representing price dynamics in the presence of a price support program. As investigated by Shonkwiler and Maddala, Holt and Johnson, and others, price support programs tend to increase expected price by censoring the price distribution at the price support level. This generates endogenous switching between a "market regime" (when the market price is higher than the support price) and a "government regime" (when government purchases take place to prevent the price from falling below the support price). Second, by introducing time-varying volatility in the model, we analyze the changing price volatility and its interaction with the price support program. Third, the effects of private and public stocks on price volatility are investigated.

The analysis is applied to the U.S. non-fat dry milk market, motivated in large part by the extensive government intervention in this market. As illustrated in figure 1, the non-fat dry milk price was at the support price level most of the time during the 1970s and 1980s. However, in the 1990s, figure 1 shows the non-fat dry milk market became somewhat liberalized, with the market price often being higher than the support price. These changes in government intervention in the U.S. non-fat dry milk market enable us to examine the impact of market liberalization on price dynamics and volatility in the presence of private stock as well as public stock.

Two of the important empirical questions to be addressed in this study are: How do agricultural policy changes affect price dynamics and price volatility? and How are those effects associated with changes in private versus public stock holding? A dynamic Tobit model under time-varying volatility shows how the price support program and stock holding affect both expected prices and the volatility of U.S. non-fat dry milk prices. Our findings show that the long-term censoring effects of the non-fat dry milk price support program can significantly increase expected price even if the price support is set below the current market price.

The Model

In this section, the process of market price determination is investigated in the presence of a government price support program. In the absence of stocks, prices can fluctuate over time in response to changes in supply and demand shifters (e.g., weather, consumer income, etc.). If such changes are unanticipated, they contribute to price uncertainty. However, in the presence of stocks, there is an incentive to reduce inventory when prices are high, and to increase inventory when prices are low. For example, an active risk-neutral storage firm would choose inventory such that the discounted expected next-period price is equal to the current price plus marginal storage cost (e.g., Williams and Wright; Deaton and Laroque 1992, 1996). Consequently, storage incentives are expected to affect price dynamics and reduce price volatility as long as stocks are positive. Then, the market price is determined by the interactions among supply, demand, and storage behavior.



Note: The prices are nominal.

Figure 1. Actual and support prices of non-fat dry milk

Let y_t^* be the market price for a commodity at time t in the absence of government intervention. Denote by $S(y_t^*, ss_t)$ the supply function and by $M(y_t^*, ds_t)$ the demand function (including demand for stocks) for that commodity at time t, where ss_t are supply shifters, ds_t are demand shifters at time t, $\partial S/\partial y_t^* > 0$, and $\partial M/\partial y_t^* < 0$. Then, the market equilibrium price y_t^* satisfies

$$S(y_t^*, ss_t) = M(y_t^*, ds_t). \tag{1a}$$

Solving this market equilibrium condition for y_t^* gives the reduced-form equation

$$y_t^* = f(\mathbf{X}_t, \beta) + e_t, \tag{1b}$$

where $\mathbf{X}_t$ is a vector of explanatory variables, β is a vector of parameters to be estimated, and e_t is an error term distributed as $N(0, \sigma_t^2)$.

Next, a government price support program is introduced in this market. Let y_t denote the observed market price at time t. The price support program involves a floor price s_t reflecting government policy at time t. When $y_t > s_t$, the price support is inactive. However, if the market price were to fall below s_t, then a government agency intervenes in the market and buys (and usually stores) the commodity at a price s_t. Thus, a perfectly elastic demand is effectively created at price s_t, thereby preventing any decrease in the market price below s_t. The observed market price y_t is then determined according to the reduced-form model:[1]

[1] The corresponding supply-demand structural forms have been analyzed by Shonkwiler and Maddala, and by Holt and Johnson.

(2a) $$y_t = \max\{y_t^*, s_t\},$$

(2b) $$y_t^* = f(\mathbf{X}_t, \beta) + e_t,$$

Equations (2a)–(2b) constitute a Tobit or censored regression model (Tobin; Amemiya), where the dependent variable y_t is censored at s_t at time t. Let $D_t = 1$ if $y_t^* > s_t$, and $D_t = 0$ otherwise. From (2a), the latent variable y_t^* is observed only if $D_t = 1$. This form corresponds to the "market regime" where the latent price is the market price ($y_t = y_t^*$) and the government price support program is inactive. Alternatively, y_t^* is censored and unobserved if $D_t = 0$, and corresponds to the "government regime" where the price support program determines the market price (with $y_t = s_t$). Equations (2a)–(2b) thus provide a generic model of price determination in the presence of a price support program, allowing for endogenous regime switching between the "market regime" and the "government regime."

Dynamic components are formally introduced in the model. Let $\mathbf{X}_t = (\mathbf{Y}_{t-1}, \mathbf{x}_t)$, where $\mathbf{Y}_{t-1} = (y_{t-1}, y_{t-2}, \ldots, y_{t-m})$ is a vector of m lagged market prices, and $\mathbf{x}_t$ denotes other explanatory variables (including previous stocks).[2] This specification gives a convenient and flexible representation of dynamics in the presence of censoring (e.g., Pesaran and Samiei 1992a, b). As noted by Zellner and Palm, there exist alternative dynamic specifications of (1b) that are consistent with the structural specification (1a). This will provide some flexibility in empirical application (see the discussion in the next section). In addition, to examine possible changes in price volatility, we allow for a time-varying standard deviation σ_t. This approach establishes a heteroskedastistic Tobit model.

The implications of the censored model (2a)–(2b) for the distribution of prices are of interest. In particular, the expected value of y_t is written as (Maddala):

(3a) $$\begin{aligned} E(y_t) &= \mathrm{Prob}(D_t = 1) * \left[f(\mathbf{X}_t, \beta) + E\left(e_t \mid e_t > s_t - f(\mathbf{X}_t, \beta)\right)\right] + \mathrm{Prob}(D_t = 0) * s_t \\ &= \left[1 - \Phi(h_t)\right] * \left[f(\mathbf{X}_t, \beta) + \sigma_t\left(\frac{\phi(h_t)}{1 - \Phi(h_t)}\right)\right] + \Phi(h_t) * s_t, \end{aligned}$$

where $h_t = [s_t - f(\mathbf{X}_t, \beta)]/\sigma_t$, and $\phi(\cdot)$ and $\Phi(\cdot)$ are, respectively, the density and distribution function for the standard normal random variable. The variance of y_t is given as (see the proof in the appendix):

(3b) $$V(y_t) = \sigma_t^2 * \left[1 - \Phi(h_t) + h_t * \phi(h_t) + h_t^2 * \Phi(h_t) - [h_t * \Phi(h_t) + \phi(h_t)]^2\right],$$

where the probability that the censored variable y_t^* is unobserved is $\mathrm{Prob}(D_t = 0) = \mathrm{Prob}[e_t < s_t - f(\mathbf{X}_t, \beta)] = \Phi(h_t)$.

Expression (3a) states that expected price $E(y_t)$ is a weighted average of the support price s_t and of the expected market price conditional on $D_t = 1$, and the weights involve the probability of censoring, $\Phi(h_t)$, i.e., the probability of facing the government regime

[2] An alternative dynamic Tobit specification is $\mathbf{X}_t = (\mathbf{Y}_{t-1}^*, \mathbf{x}_t)$, where $\mathbf{Y}_{t-1}^* = (y_{t-1}^*, y_{t-2}^*, \ldots)$ is a vector of lagged latent variables, and $\mathbf{x}_t$ denotes other explanatory variables (Lee; Wei). As noted by Lee, this includes as a special case the Tobit model under autocorrelated error terms (Zeger and Brookmeyer). We did not rely on this specification for two reasons: (a) using lagged latent variables means that the likelihood function involves multiple integrals (which requires switching from the standard maximum-likelihood method to simulated estimation methods); and (b) estimating time-varying σ_t becomes more difficult in this context (see Lee).

at time t. Equation (3b) indicates that the relative variance $[V(y_t)/\sigma_t^2]$ equals $[1 - \Phi(h_t) + h_t * \phi(h_t) + h_t^2 * \Phi(h_t) - [h_t * \Phi(h_t) + \phi(h_t)]^2]$. Therefore, this expression measures the impact of censoring from the price support program on price volatility. For example, in the absence of censoring, the relative variance would equal 1. Alternatively, under censoring (i.e., under the government regime), the relative variance $[V(y_t)/\sigma_t^2]$ is reduced, indicating how a price support program would decrease price volatility.

Finally, with $\mathbf{Y}_{t-1}$ involving lagged actual prices $[\mathbf{Y}_{t-1} = (y_{t-1}, y_{t-2}, \ldots, y_{t-m})]$ and the error terms e_t being independently distributed, the likelihood function of sample information can be evaluated using single integrals (Maddala, chapter 6). Therefore, model (2a)–(2b) can be estimated by standard maximum-likelihood estimation, thereby allowing more complex dynamics involving a larger number of lags m (compared to alternative specifications involving lagged latent prices).

An Application to the U.S. Non-fat Dry Milk Market

In this section, the dynamics of U.S. non-fat dry milk prices are analyzed. The determinants of non-fat dry milk price and its volatility are investigated, with a special focus on the role of the government price support program and the effects of private and public stocks. This analysis is conducted in the context of a heteroskedastic Tobit model that allows for endogenous regime switching and time-varying volatility, where commercial and government stocks affect both the mean and the variance of prices.

The empirical analysis is based on monthly data for the period January 1970–July 2000. Monthly non-fat dry milk stock data were obtained from the U.S. Department of Agriculture's (USDA's) National Agricultural Statistics Service and Agricultural Stabilization and Conservation Service. This stock series is measured in thousand pounds at the beginning of every month. Monthly non-fat dry milk prices (measured in cents/pound) were obtained from various issues of *Dairy Market News* (USDA/Agricultural Marketing Service).[3]

Figure 1 (introduced earlier) shows actual non-fat dry milk price and the corresponding support price. Two extreme periods of government involvement can be identified: the early 1980s when the market price was always at the support price, and the mid-1990s when the market price was always above the support price. In the former period, Congress set the support price at a high and constant level, implying the consistent presence of the government regime. In the latter period, the support price was typically lower than the market price, implying the consistent presence of the market regime. Other periods exhibited some changes between the market regime (when the price support is inactive) and the government regime (when the price support is active).[4]

The model specification (1)–(2) is general, and there remains the issue of choosing between the structural form (1a) and the reduced form (1b). The structural approach has the advantage of providing direct information on the supply-demand conditions (e.g., Holt and Johnson; Shonkwiler and Maddala). However, it requires information on supply-demand shifters that are crucial to identify the structural parameters. Because

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[3] We use wholesale price of non-fat dry milk for human food.

[4] Except for the period of the early 1980s, the Secretary of Agriculture had discretion in making some adjustments in the support price depending on market conditions and government stocks.

milk production is a continuous process, measuring supply-demand shifters on a *monthly* basis can be difficult. In addition, we can expect significant dairy market adjustments which take place over many months. Unfortunately, we do not have good a priori information about the dynamic effects of dairy supply-demand shifters from one month to the next. As a result, applying (1)–(2) to monthly non-fat dry milk prices using a structural approach [based on (1a)] is problematic. For this reason, we pursue below the reduced-form approach [based on (1b)]. While the reduced-form approach has the disadvantage of not providing direct information on the supply-demand conditions, it does not raise difficult identification issues for the estimated parameters. Moreover, it is well suited to investigate dynamic issues. As we illustrate below, this design can provide useful information on the interactions between government price support program and price dynamics.

As noted by Zellner and Palm, there exist alternative specifications corresponding to (1b). In the context of linear models, these alternative specifications are obtained by premultiplying (1b) by various polynomial lag matrices. This procedure can generate the "transfer function" specification (where each dependent variable depends on its lagged values and on exogenous variables) and the "final equation" specification (where each dependent variable depends only on its lagged values). Zellner and Palm stress that these alternative specifications are consistent with the structural specification (1a).[5]

Our empirical investigation utilizes the Tobit specification summarized in (2a) and (2b), where $f(\cdot) = \beta_0 + \Sigma_{j=1}^{m} \beta_j y_{t-j} + \mathbf{x}_t \bar{\beta} + e_t$,[6] and $\sigma_t = \exp[\gamma_0 + \mathbf{z}_t \bar{\gamma}]$.[7] The error term e_t is assumed to be distributed $N(0, \sigma_t^2)$ and serially uncorrelated. The parameters β_0, β_j, $\bar{\beta}$, γ_0, and $\bar{\gamma}$ are to be estimated. Finally, $\mathbf{z}_t$ is a vector of explanatory variables affecting σ_t. Note, in the case where $\bar{\gamma} \neq 0$, this allows for heteroskedasticity, where $\mathbf{z}_t$ affects the volatility of prices.

First, in order to investigate the effects of stocks on the conditional mean and variance of non-fat dry milk price, we introduce lagged non-fat dry milk stocks in $\mathbf{x}_t$ and $\mathbf{z}_t$. We allow the stock effects to differ between private stocks and public stocks. As shown in figure 2, private and public stocks exhibit different patterns over the sample period. As expected, government stocks are high (low) when the price support and government purchases are active (inactive) in the market. Note that private stock and public stock perform different market functions: the former is motivated by anticipated price increases (e.g., Williams and Wright), while the latter is the key policy instrument used in implementing the government price support program (which prevents price decreases). Because this pattern suggests private stocks and public stocks may have different effects on prices, we include separately lagged commercial stocks (CS_{t-1}) and lagged government stocks (GS_{t-1}) in $\mathbf{x}_t$ and $\mathbf{z}_t$. This combination provides a framework to investigate possible

[5] Zellner and Palm also discussed the strengths and weaknesses of each approach.

[6] This is a linear version of the reduced-form equation for latent price y_t^*. Note, Deaton and Laroque (1992, 1996), and Ng have argued that private storage generates nonlinear price dynamics with regime switching between stockout and speculative stockholding. Ng reports empirical evidence of strong persistence in the stockout regime which is inconsistent with the theory. Such findings suggest a "convenience yield" may smooth out the differences across regimes, making it unclear what nonlinearities arise in price dynamics. In the absence of strong a priori information about nonlinearities, a linear specification is convenient and parsimonious for our purpose. Exploring nonlinearity issues is a good topic for further investigation.

[7] Alternative specifications were explored for the standard deviation σ_t. While we explored an autoregressive conditional heteroskedastic (ARCH) specification, we did not pursue it for two reasons. First, the increased variability toward the end of the sample period (see figure 1) meant the dynamic volatility appeared nonstationary. Second, this specification involves multiple integrals and requires the use of simulated maximum-likelihood methods which become more complex to estimate under ARCH specification (see Lee).

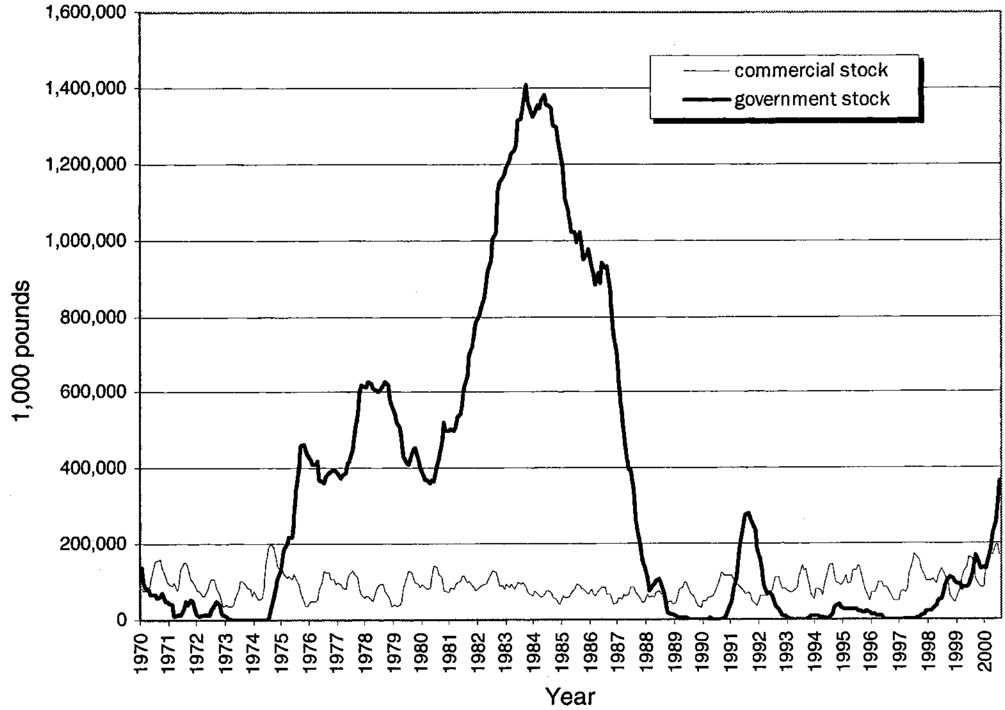

Figure 2. Commercial and government stocks of non-fat dry milk

differences between private and public stock impacts on price levels and price volatility (discussed below).

From the economic literature on storage (e.g., Williams and Wright; Deaton and Laroque 1992, 1996), we expect higher (lower) stocks at time $t - 1$ will tend to reduce (increase) the market price at time t. Also, larger (smaller) stocks are expected to generate lower (higher) price volatility. Second, from competitive storage theory, a higher (lower) interest rate provides a disincentive (incentive) to hold private stocks, which is expected to affect both price level and price volatility. On that basis, the interest rate (R_t)[8] is included in $\mathbf{x}_t$ and $\mathbf{z}_t$ to capture its effects on the conditional mean and variance of non-fat dry milk price.

Third, we include in $\mathbf{x}_t$ a time trend TT and monthly dummy variables ($M_i = 1$ for the ith month, zero otherwise). The time trend accounts for the long-term impacts of inflation and technological progress on prices. The monthly dummy variables M_i incorporate seasonality effects in the non-fat dry milk market, and take into account the monthly effects of seasonal supply and demand shifters.

Next, the issue of a time-varying σ_t (and the associated heteroskedasticity) is explored. This process reflects possible changes in price volatility unrelated to the price support program. Given the standard deviation $\sigma_t = \exp[\gamma_0 + \mathbf{z}_t \bar{\gamma}]$, we consider two specifications for $\mathbf{z}_t$.

In the first specification (model I), $\mathbf{z}_t = [y_{t-1}, GS_{t-1}, (y_{t-1} * GS_{t-1}), CS_{t-1}, (y_{t-1} * CS_{t-1}), TT, R_t]$, where R_t is the interest rate and TT is a time trend capturing long-term changes in

[8] The monthly interest rates were calculated based on six-month Treasury bill rates (Federal Reserve Bank).

volatility. We include lagged non-fat dry milk price (y_{t-1}) in $\mathbf{z}_t$. This specification introduces possible feedback between market conditions and price volatility, allowing for price volatility to be affected by the price level. The market regime (government regime) would likely correspond to the situation where y_{t-1} is high (low). This specification also provides a framework capturing possible structural changes in market instability not captured by censoring effects during the sample period. For example, it is used to investigate whether the high non-fat dry milk prices of the 1990s (reflecting the prevalence of the market regime) have been associated with greater latent price volatility.

Both lagged commercial stocks (CS_{t-1}) and lagged government stocks (GS_{t-1}) are included to investigate the possibly different effects of stocks (commercial stock versus government stock) on price volatility. In this specification, the impact of lagged non-fat dry milk price (y_{t-1}) on σ_t is allowed to vary with both lagged commercial stocks (CS_{t-1}) and lagged government stocks (GS_{t-1}), motivating the introduction of interaction variables ($y_{t-1} * GS_{t-1}$, $y_{t-1} * CS_{t-1}$) as latent volatility shifters among the $\mathbf{z}_t$ variables. The monthly interest rate (R_t) is included to investigate the effects of interest rate on price volatility. As suggested by storage theory, a rise in interest rate provides a disincentive to hold private stocks, which can contribute to an increase in price volatility.

The second specification (model II) simplifies the standard deviation specification by excluding these interaction variables in $\mathbf{z}_t$. It corresponds to $\mathbf{z}_t = [y_{t-1}, GS_{t-1}, CS_{t-1}, TT, R_t]$. Implicitly, this specification restricts the impacts of lagged non-fat dry milk price (y_{t-1}) on σ_t not to vary with lagged commercial stocks (CS_{t-1}) and lagged government stocks (GS_{t-1}).[9]

These considerations result in the following two model specifications:[10]

MODEL I:

(4a) $$y_t = \max\{y_t^*, s_t\},$$

(4b) $$y_t^* \equiv f(\mathbf{X}_t, \beta) = \beta_0 + \beta_T TT + \Sigma_{i=1}^{11} \beta_k M_i + \Sigma_{k=1}^{m} \beta_k y_{t-k} + \beta_{CS} CS_{t-1} + \beta_{GS} GS_{t-1} + \beta_R R_t + e_t,$$

(4c) $$\sigma_t = \exp\big[\gamma_0 + \gamma_1 y_{t-1} + \gamma_2 GS_{t-1} + \gamma_3 (y_{t-1} * GS_{t-1}) + \gamma_4 CS_{t-1} + \gamma_5 (y_{t-1} * CS_{t-1}) + \gamma_6 TT + \gamma_7 R_t\big];$$

MODEL II:

(5a) $$y_t = \max\{y_t^*, s_t\},$$

(5b) $$y_t^* \equiv f(\mathbf{X}_t, \beta) = \beta_0 + \beta_T TT + \Sigma_{i=1}^{11} \beta_k M_i + \Sigma_{k=1}^{m} \beta_k y_{t-k} + \beta_{CS} CS_{t-1} + \beta_{GS} GS_{t-1} + \beta_R R_t + e_t,$$

(5c) $$\sigma_t = \exp\big[\gamma_0 + \gamma_1 y_{t-1} + \gamma_2 GS_{t-1} + \gamma_4 CS_{t-1} + \gamma_6 TT + \gamma_7 R_t\big],$$

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[9] Alternative model specifications were also explored. They include the use of seasonal dummies as part of $\mathbf{z}_t$ in the specification of the standard deviation σ_t. These alternative models gave results comparable to the ones reported here, with similar qualitative implications for price dynamics.

[10] Note that some of the exogenous variables (e.g., supply-demand shifters) are not included in the model. This can be interpreted as a "final equation" specification, where the dynamics of these exogenous variables are "substituted in" to generate a dynamic model where the dependent variable depends only on its lagged values (see Zellner and Palm).

where s_t is a floor price, y_t^* is the latent non-fat dry milk price at time t, and e_t is an error term distributed $N(0, \sigma_t^2)$. Except for restrictions on time-varying volatility ($\gamma_3 = \gamma_5 = 0$), model II is identical to model I. The inclusion of lagged stocks CS_{t-1} and GS_{t-1} in both the latent function $f(\cdot)$ and the standard deviation specification for σ_t allows for different effects of private and public stocks on prices. Beyond censoring, this specification can account for the market equilibrium effects of government intervention (through GS_{t-1}) as well as private storage behavior (through CS_{t-1}) on both latent price and price volatility.

In the absence of censoring (where $y_t^* = y_t$), equations (4b) and (5b) would reduce to standard autoregressive models of order m, AR(m), with the time trend TT, monthly dummies (M_i, $i = 1, \ldots, 11$), lagged commercial stocks (CS_{t-1}), lagged government stocks (GS_{t-1}), and the rate of interest (R_t). As such, the reduced forms (4a)–(4c) and (5a)–(5c) provide an extension of such models in the presence of censoring and conditional heteroskedasticity. They constitute the econometric specifications used below in the empirical investigation of the impact of price support programs on price dynamics in the U.S. non-fat dry milk market.

Econometric Results

Following the discussion in the two previous sections, models I and II are applied to the U.S. non-fat dry milk market (based on monthly data for the period 1970–2000) to investigate the determinants of non-fat dry milk price and price volatility. The model is estimated by the maximum-likelihood method, assuming the error terms e_t are serially uncorrelated. Assuming a correct specification, the maximum-likelihood estimation method produces consistent parameter estimates. This property can still hold even in the presence of serial correlation in the error term (Robinson). In addition, if the error term is serially uncorrelated [with e_t/σ_t being white noise $N(0, 1)$], then the maximum-likelihood estimator is consistent and asymptotically efficient.

The order of the AR process (m) in (4b) and (5b) was determined using the Schwarz criterion (Judge et al., p. 426). This involves choosing m so as to maximize [ln(maximum likelihood) - $K * \ln(T)/2$], where K is the number of parameters and T is the number of observations. The Schwarz criterion selected $m = 2$ months for both models. Thus, the analysis below is based on the dynamic Tobit specification (4a)–(4c) and (5a)–(5c) with $m = 2$.

First, the issue of serial dependence in the error terms e_t is investigated for both model I and model II. This process involves performing a diagnostic test for white noise of the error terms e_t/σ_t in (4b) and (5b). Following Robinson, Bera, and Jarque, a Lagrange multiplier (LM) test is used to test the null hypothesis of serial independence in the error terms. The LM test is easier to implement than a likelihood-ratio test or a Wald test because it does not require estimating the model under the alternative hypothesis of serial correlation. The LM test results are reported in table 1. They show no statistical evidence of serial correlation in the error terms up to order 12 in both models. Thus, there is no strong evidence against e_t/σ_t being white noise. This finding suggests the model specifications appropriately capture price dynamics.

Second, we explored whether the stock effects in (4a)–(4c) and (5a)–(5c) are the same between private stocks and public stocks. Formally, this is done by testing the null hypotheses: $\beta_{CS} = \beta_{GS}$, $\gamma_2 = \gamma_4$, and $\gamma_3 = \gamma_5$ in model I; and $\beta_{CS} = \beta_{GS}$, and $\gamma_2 = \gamma_4$ in model II. Using a likelihood-ratio test, $\chi^2_{[3]} = 117.42$ (p-value = 0.000) for model I, and $\chi^2_{[2]} = 116.84$

Table 1. Serial Correlation Test Results for the Error Terms (e_t)

Serial Correlation of Order r	Lagrange Multiplier Test Statistic		Degrees of Freedom	Critical Value (at 5% level of significance)	Test Result
	Model I	Model II			
1	1.933	1.653	1	3.84	Fail to reject
2	2.301	1.958	2	5.99	Fail to reject
3	2.382	2.051	3	7.82	Fail to reject
4	2.383	2.061	4	9.49	Fail to reject
5	2.394	2.067	5	11.07	Fail to reject
6	2.549	2.233	6	12.59	Fail to reject
7	2.550	2.234	7	14.07	Fail to reject
8	3.182	2.835	8	15.51	Fail to reject
9	3.274	2.918	9	16.92	Fail to reject
10	3.318	2.971	10	18.31	Fail to reject
11	3.323	2.974	11	19.68	Fail to reject
12	4.464	4.264	12	21.03	Fail to reject

(p-value = 0.000) for model II. Therefore, we strongly reject the null hypothesis and conclude that private and public stocks have different effects on price and price volatility in both models.

Next, we investigate whether it is appropriate to introduce time-varying latent volatility (i.e., heteroskedasticity) in model I and model II. This issue was addressed by testing the null hypothesis of homoskedasticity, where $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = \gamma_5 = \gamma_6 = \gamma_7 = 0$ in (4c), and $\gamma_1 = \gamma_2 = \gamma_4 = \gamma_6 = \gamma_7 = 0$ in (5c). Using a likelihood-ratio test, the chi-squared statistics are $\chi^2_{[7]} = 354.7$ (p-value = 0.000) for model I, and $\chi^2_{[5]} = 353.4$ (p-value = 0.000) for model II. Thus, the null hypothesis of homoskedasticity is strongly rejected in both models. In other words, there is strong empirical evidence of time-varying volatility in non-fat dry milk prices during the sample period. However, this changing volatility is unrelated to the effects of the price support program because the censoring effects of the program are already captured in the Tobit specification.

Table 2 reports the parameter estimates of the heteroskedastic dynamic Tobit models presented in (4a)–(4c) and (5a)–(5c). First, all of the lagged price effects are statistically significant. This result reflects evidence of significant price dynamics in the U.S. non-fat dry milk market. Note that β_{t-1}, the coefficient of y_{t-1}, equals 1.316 for model I and 1.323 for model II, suggesting an initial overreaction to a recent price change. The roots of the estimated AR(2) for the latent prices are all in the unit circle,[11] implying the dynamic model is stationary. However, the dominant root is close to one, suggesting the latent dynamics are close to a unit-root process. Both lagged private stocks and lagged public stocks have negative and significant impacts on latent price in each of the models. This finding confirms, as expected, that increasing (decreasing) lagged stock puts downward (upward) pressure on average price. The time trend parameter is positive and significant in each model, providing evidence of the long-term impacts of inflation and technological progress on prices. Some of the monthly dummy variables are statistically significant in each model.

[11] The roots are 0.997 and 0.319 for model I, and 0.992 and 0.330 for model II.

Table 2. Parameter Estimates for Heteroskedastic Dynamic Tobit Models: U.S. Non-fat Dry Milk Price, January 1970–July 2000

		Model I		Model II	
Parameter	Definition	Estimate	Standard Error	Estimate	Standard Error
β_0	Intercept	0.878***	0.295	0.893***	0.309
β_{t-1}	Price of non-fat dry milk at time $t-1$ (y_{t-1})	1.316***	0.041	1.323***	0.039
β_{t-2}	Price of non-fat dry milk at time $t-2$ (y_{t-2})	-0.318***	0.039	-0.328***	0.037
β_R	Interest rate (R_t)	63.62	40.37	65.25	41.25
β_{GS}	Lagged government stock (GS_{t-1})	-2.337***	0.399	-2.311***	0.400
β_{CS}	Lagged commercial stock (CS_{t-1})	-10.36***	2.375	-10.15***	2.293
β_T	Time trend (TT)	0.042*	0.024	0.049**	0.024
β_{M1}	Dummy for 1st month ($M1$)	-0.370	0.267	-0.361	0.282
β_{M2}	Dummy for 2nd month ($M2$)	-0.555	0.400	-0.525	0.389
β_{M3}	Dummy for 3rd month ($M3$)	-0.133	0.283	-0.137	0.285
β_{M4}	Dummy for 4th month ($M4$)	-0.610	0.872	-0.572	0.808
β_{M5}	Dummy for 5th month ($M5$)	-1.337*	0.789	-1.136*	0.753
β_{M6}	Dummy for 6th month ($M6$)	-0.342	0.406	-0.339	0.406
β_{M7}	Dummy for 7th month ($M7$)	-0.188	0.191	-0.148	0.210
β_{M8}	Dummy for 8th month ($M8$)	0.238	0.197	0.245	0.210
β_{M9}	Dummy for 9th month ($M9$)	0.439**	0.212	0.439*	0.253
β_{M10}	Dummy for 10th month ($M10$)	0.243	0.206	0.265	0.215
β_{M11}	Dummy for 11th month ($M11$)	0.023	0.180	0.015	0.203
Intercept	Intercept for standard deviation equation	-0.539	0.610	-1.084***	0.290
γ_1	Price of non-fat dry milk at time $t-1$ (y_{t-1})	0.032***	0.007	0.038***	0.004
γ_2	Lagged government stock (GS_{t-1})	-1.936	1.412	-1.058***	0.218
γ_3	$y_{t-1} * GS_{t-1}$	-0.010	0.016		
γ_4	Lagged commercial stock (CS_{t-1})	-21.03***	5.711	-15.57***	1.380
γ_5	$y_{t-1} * CS_{t-1}$	0.065	0.066		
γ_6	Time trend (TT)	-0.028**	0.014	-0.023**	0.010
γ_7	Interest rate (R_t)	54.25***	16.19	55.17***	15.98
Log likelihood		-564.49		-565.12	
No. of observations		365			

Note: Single, double, and triple asterisks (*) denote statistical significance at the 10%, 5%, and 1% levels, respectively.

Consistent with the previous heteroskedasticity test result, most of the estimated parameters in the standard deviation equations (4c) and (5c) are highly significant (table 2). First, the coefficient estimates of γ_2 and γ_4 are statistically significant, with the exception of public stocks (GS_{t-1}) in model I. The estimates are negative and capture the stock effect on price volatility. These findings provide evidence showing both private and public stocks tend to reduce price volatility over the sample period, which is consistent with the indirect evidence found by Shively. Interestingly, the effects of private stocks on price volatility are found to be much stronger than the effects of public stocks—indicating a market liberalization involving a switch from public stocks to private stocks would contribute to market price stabilization (beyond the censoring effects already captured by the Tobit model). These effects are further evaluated below.

Next, the presence of feedback between price volatility and price level was investigated. In model I, this is done by testing the null hypothesis that $\gamma_1 = \gamma_3 = \gamma_5 = 0$ in (4c). The corresponding likelihood-ratio test statistic is 81.0. At a 5% significance level, we reject the null and conclude the lagged non-fat dry milk price is a significant determinant of σ_t. The coefficient of γ_1 is positive and significant in both model I and model II. Thus, as the lagged price increases (e.g., in the market regime of the 1990s), latent price volatility increases. Alternatively, as the lagged price decreases, latent price volatility is estimated to decrease.

Note from table 2 that the interaction effects between stocks and price level (as measured by the coefficients of γ_3 and γ_5) are not statistically significant from zero in model I. This was tested formally using a likelihood-ratio test. Based on a 5% significance level, the test result showed no strong statistical evidence against the null hypothesis that $\gamma_3 = \gamma_5 = 0$. We conclude there is no statistical evidence of significant interaction effects between price and stocks on latent price volatility. This result provides evidence in support of the specification given in model II. On that basis, the remainder of the article focuses on the specification in model II.[12] Table 2 shows the coefficient of the time trend parameter γ_6 is negative and significant, indicating market instability has changed during the sample period. Finally, the interest rate parameter γ_7 is found to be positive and statistically significant, confirming that interest rate has a positive impact on price volatility in the non-fat dry milk market.

The performance of the estimated model is evaluated by comparing the expected prices obtained from (3a) with actual prices. As observed from the graph in figure 3, the model has a high explanatory power during the 1970–2000 sample period. Figure 3 also provides useful information about the changing nature of the U.S. non-fat dry milk market over the last 30 years. It illustrates the stable non-fat dry prices of the 1970s and 1980s when the price support was consistently binding, while it also documents the increased volatility of non-fat dry milk prices in the 1990s.

Figure 4 shows the predicted standard deviation of non-fat dry milk price ($V(y_t)^{1/2}$), as simulated from equation (3b) over the sample period. The simulation reveals large changes in price instability. The standard deviation of non-fat dry milk price was lowest in the early 1980s. This observation can be explained as follows: (*a*) during that period the market volatility was low (as measured by σ_t), and (*b*) the censoring effects of the price support program were strong and generated a further reduction in price variance.

Figure 4 also shows that the standard deviation of non-fat dry milk price was highest in the 1990s. Again, this occurrence is due to two factors: (*a*) over that period the market volatility (as measured by σ_t) was larger, and (*b*) the censoring effects of the price support program were moderate as the price support was often lower than the market price. From (5c) and the estimates reported in table 2, the increase in latent volatility σ_t in the 1990s can be attributed to a higher price and lower public stocks during this period. As seen in figure 4, the standard deviation of non-fat dry milk price still fluctuated significantly during the 1990s. This variation is due to price and stock effects, i.e., the standard deviation σ_t decreases (increases) when the price is low (high) and/or stocks are high (low). These findings validate the important effects of storage behavior on price volatility.

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[12] Note that model I gave results which were qualitatively similar to the ones presented below.

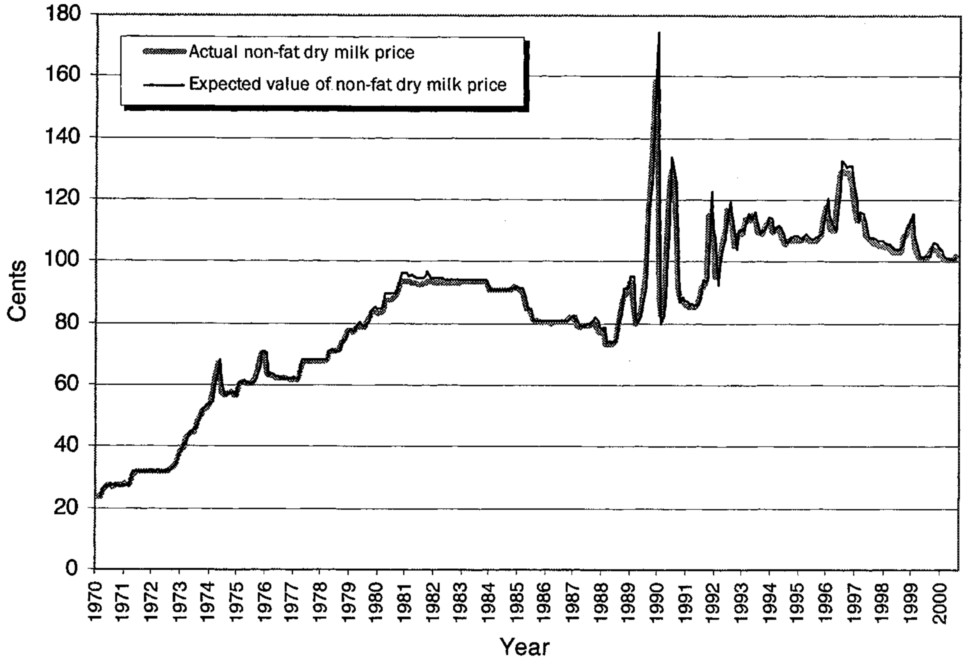

Figure 3. Expected and actual prices of non-fat dry milk

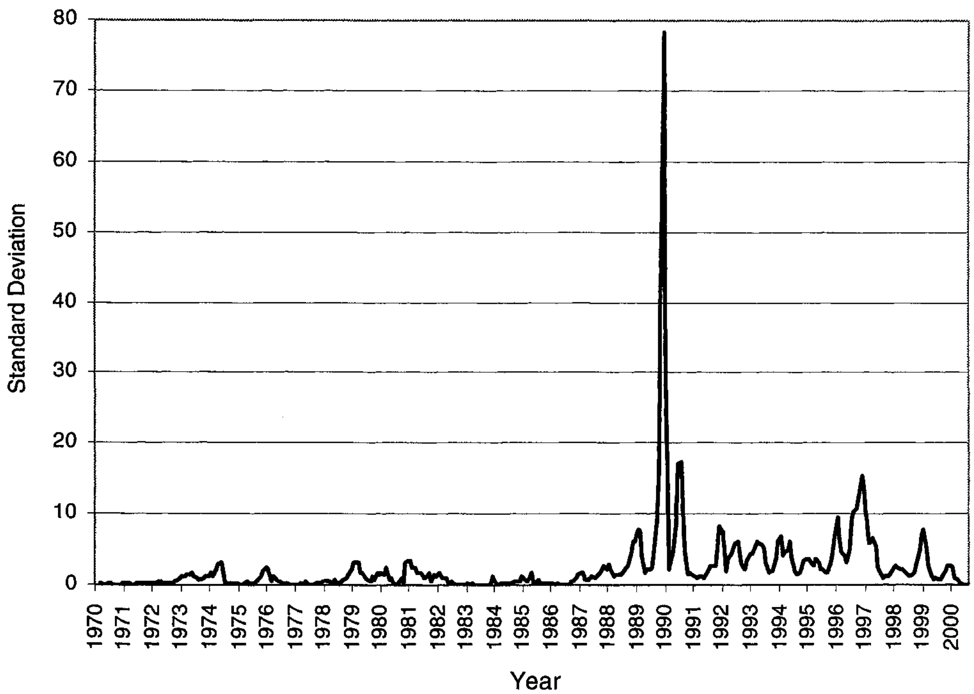

Figure 4. Estimated standard deviation of non-fat dry milk price

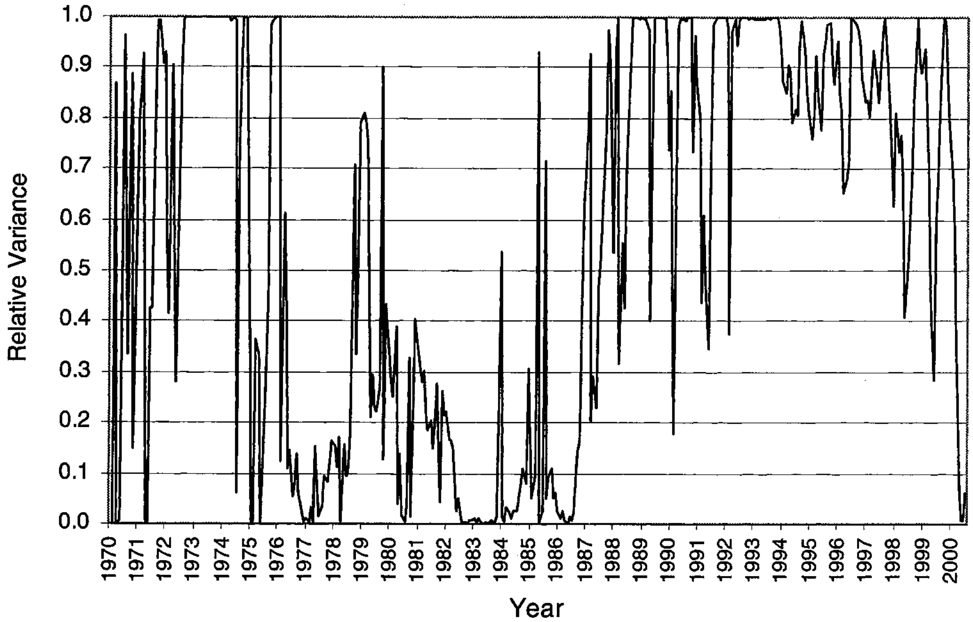

Figure 5. Relative variance, $V(y_t)/\sigma_t^2$, of non-fat dry milk price due to censoring

Finally, we investigate the relative role of the price support program in the estimated price variance. This is done by calculating the relative variance $V(y_t)/\sigma_t^2$ from equation (3b). The results are presented in figure 5. As discussed earlier, the relative variance $V(y_t)/\sigma_t^2$ is bounded between zero and one: it is equal to one in the absence of censoring, and it becomes close to zero in the presence of strong censoring effects. As expected, figure 5 indicates censoring effects are persistent for most of the sample period, except in the middle and late 1990s (when the relative variance is high). And they are strongest in the early 1980s (when the relative variance is close to zero). This pattern provides evidence that the price support program has contributed to significant reductions in price instability in the U.S. non-fat dry milk market over the last 30 years.

Implications

Given the large changes in price instability just documented, it is useful to investigate the implications of the model for price dynamics. The analysis in this section relies on dynamic multipliers. We proceed by simulating the effects of changes in selected variables on the path of expected price and the variance of price given in (3a) and (3b). Note, however, that equation (3a) involves nonlinear dynamics, because the functions ϕ and Φ are nonlinear functions of lagged prices. Due to this nonlinear dynamic nature of (3a), all dynamics are "local" in the sense that they depend on the particular path being evaluated. Therefore, we focus our attention on two scenarios: one covering the period starting in September 1985, and one covering the period starting in January 1994. Recall that these two scenarios correspond to two extreme situations related to the non-fat dry milk price support program. The first scenario ($\geq$ 1985.09) can be interpreted as representing the "government regime," where the price support is strongly binding. The second

Table 3. Elasticities of the Mean Price $E(y_t)$ and Standard Deviation $V(y_t)^{1/2}$ with Respect to Temporary Shocks in Commercial and Government Stock

Description		Commercial Stock	Government Stock
Market Regime:	Standard deviation effects	-0.967	-0.009
	Mean price effects	-0.0096	-0.0002
Government Regime:	Standard deviation effects	-6.791	-9.032
	Mean price effects	-0.003	-0.004

scenario ($\geq$ 1994.01) represents the "market regime," corresponding to a period when government purchases are inactive.

First, using (3a) and (3b), we simulated the effects of changing lagged non-fat dry milk stocks (both private and public stocks as measured by CS_{t-1} and GS_{t-1}) on the mean and standard deviation of non-fat dry milk price, $E(y_t)$ and $V(y_t)^{1/2}$. Table 3 presents the results under the two scenarios, reporting the elasticities of a temporary shock in private and public non-fat dry milk stocks (CS_{t-1} and GS_{t-1}) on the current price $E(y_t)$ and the standard deviation of the price $V(y_t)^{1/2}$. Under the government regime, the elasticities of mean price with respect to both public and private stocks were found to be negative but small: -0.003 with respect to private stock, and -0.004 with respect to public stock. Similarly, under the market regime, the elasticities of mean price with respect to private and public stocks are -0.0096 and -0.0002, respectively. This finding suggests such stock effects are very small. While the estimated stock effects in (5b) are statistically significant, our elasticity estimates indicate that their marginal effects on mean price are small. Specifically, large changes in public or private stocks are needed to have a substantial effect on expected price.

However, as observed from table 3, the effects of private and public stocks on price volatility are larger. Under the government regime, the elasticities of $V(y_t)^{1/2}$ with respect to private and public stocks are -6.791 and -9.032, respectively. Under the market regime, the elasticity of $V(y_t)^{1/2}$ is smaller with respect to private stock (-0.967), and much smaller with respect to public stock (-0.009).

These results have two implications. First, stock accumulation in both the private and public sectors can contribute to significant reductions in price volatility. The exception is for public stock under the market regime where the effect is estimated to be very small. Second, for both private and public stock, this effect is much stronger when the price support is binding, reflecting the fact that the censoring effect is large (small) under the government (market) regime. Thus, important interaction effects are identified between private and public stocks, and government policy on price volatility.

Next, the effects of a *temporary* shock in the price of non-fat dry milk were simulated. The results, graphed in figure 6 under the two scenarios, show the dynamic impact of an exogenous change in non-fat dry milk price y_t on the expected future prices $E(y_{t+j})$ and the standard deviation of future prices $V(y_{t+j})^{1/2}$ $\{j = 0, 1, 2, ...\}$. As seen from figure 6, under the government regime scenario, market price changes have only a small short-term effect on price dynamics and price volatility. Their long-term effects are negligible. In this situation, the price support is the key determining factor for the market price.

Under the market regime scenario, however, the dynamics are quite different. The simulations show both short-term and long-term price effects are large: important dynamic adjustments take place in the non-fat dry milk market in the absence of government intervention. This result reflects the near unit-root process of our dynamic model. And, as shown in figure 6 under the government regime, a temporary shock in the non-fat dry milk price generates only a small short-term effect (with negligible long-term impact) on price volatility. However, under the market regime, an increase in market price produces a larger positive and longer-term effect on price volatility. This result is attributed to the positive feedback effect estimated between price level and latent price volatility.

We then simulated the effects of a *permanent* shock in the support price in the U.S. non-fat dry milk market. The results are presented in figure 7 under the two scenarios. Figure 7 shows the dynamic impact of a *permanent* change in the support price s_t on the expected future prices $E(y_{t+j})$ and the standard deviation of future prices $V(y_{t+j})^{½}$ $\{j = 1, 2, 3, ...\}$. The support price is found to have large effects on price dynamics and price volatility under the government regime scenario. For example, when the support price is binding, a permanent increase in the price support generates almost parallel increases in the non-fat dry milk price in both the short and long run. Again, this is intuitive because under the government regime scenario, the price support is the key factor for the market price determination.

The dynamic impacts of the support price on $V(y_{t+j})^{½}$ appear more complex. Under the government regime scenario, the initial effect ($j = 1$) on the standard deviation is negative and large (with a corresponding elasticity of −0.81), establishing that the censoring effect of the price support program effectively decreases short-term price instability. However, as shown in figure 7, the next period effect ($j = 2$) is positive. This can be attributed to the short-term overshooting estimated by the model. In other words, an increase in y_t tends to generate a more than proportional increase in y_{t+1}, which reduces the negative censoring effect of the price support on the price variance at time $t+1$.

As illustrated in figure 7, in the longer term, the effects of a permanent increase in the price support on $V(y_{t+j})^{½}$ are found to be positive but small. Again, this finding is attributed to the estimated positive feedback effect between price level and latent price volatility. Thus, while a permanent increase in the price support reduces short-term price volatility due to censoring effects, these price stabilization effects are found to disappear in the intermediate term and long term. In other words, under the government regime scenario, while the price support program reduces short-term price instability, it does not contribute to a significant reduction in long-term price instability. As such, the findings identify the need to differentiate short-run versus long-run effects of price stabilization in the analysis of a price support program.

Next, we examine the impact of the price support on price dynamics and price volatility under the market regime scenario. In the market regime, the short-term elasticity of $V(y_{t+1})^{½}$ with respect to the support price is −0.026. As shown in figure 7, the impact on price volatility is small in both the short and long run. Thus, when the price support is lower than the market price, a permanent increase in support price does not have a large effect on price volatility. As anticipated, the effects of the support price on expected price vary greatly between the government regime and the market regime.

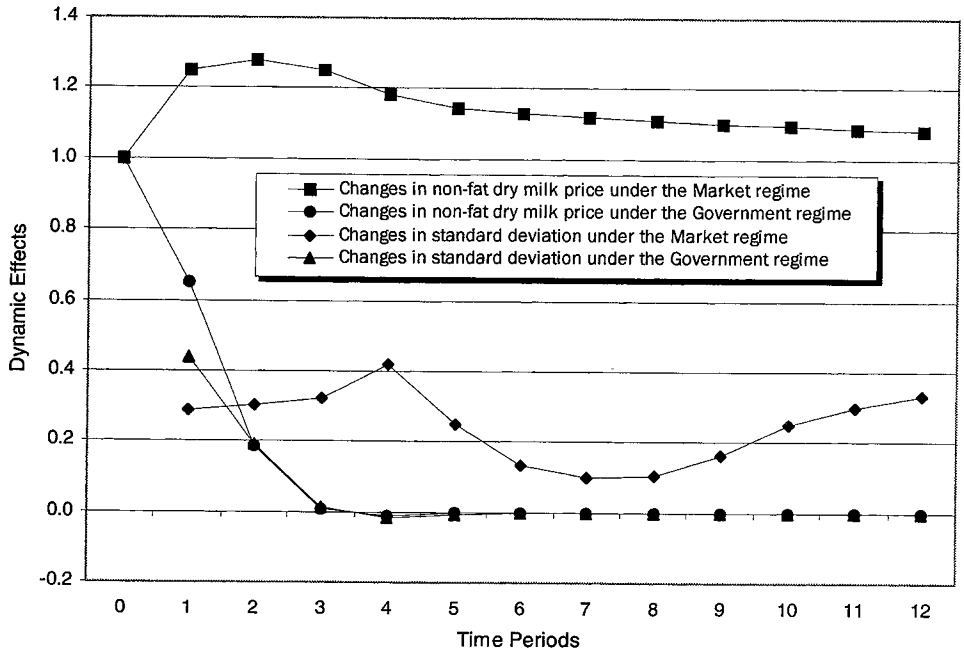

Figure 6. Effects of a temporary shock in non-fat dry milk price on expected future prices $E(y_{t+j})$ and standard deviation of future prices $V(y_{t+j})^{½}$

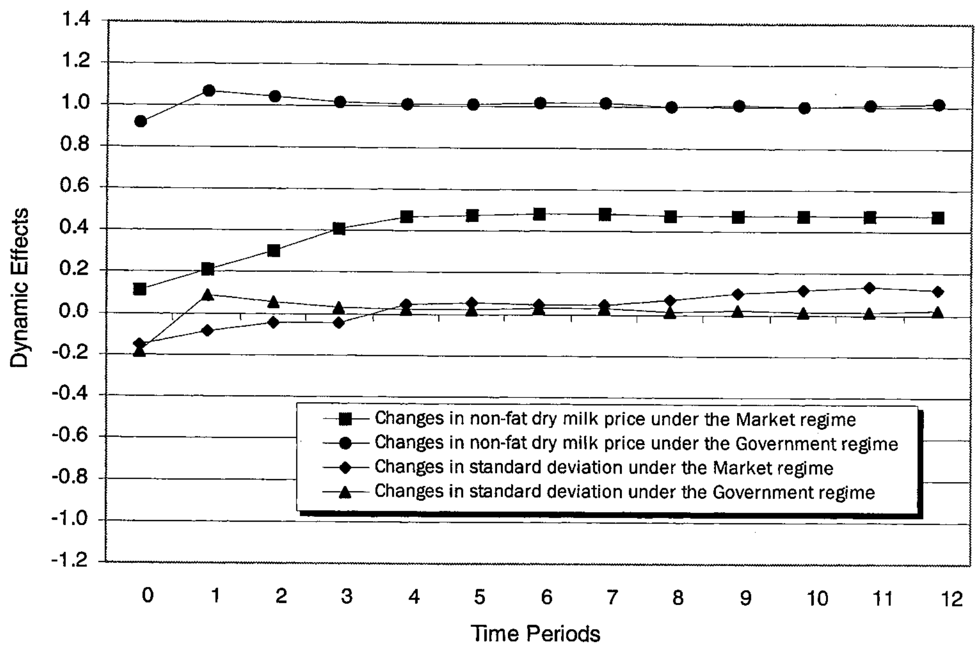

Figure 7. Effects of a permanent shock in the support price of non-fat dry milk on expected future prices $E(y_{t+j})$ and standard deviation of future prices $V(y_{t+j})^{½}$

Under the market regime scenario in figure 7, the short-term effect of the price support on expected price is positive but small, as anticipated due to small censoring effects. However, the longer-term impact of a permanent increase in the price support on expected price is larger. For example, figure 7 shows, in the market regime, the long-term marginal impact of a permanent change in the price support on expected price is 0.48, suggesting the cumulative impact of a higher support price on expected market price is not negligible even when the level of support price is relatively low. This result is attributed to the near unit-root process of the latent prices which tends to amplify small positive impacts over time, confirming that limited government intervention can still have a significant effect on long-term price dynamics. This would correspond to a market regime scenario where the support price is set lower than the long-term average price, implying infrequent government purchases taking place only when the price is "unexpectedly low." This finding suggests it is possible for government policy to have a significant effect on long-term market prices at a relatively low cost to taxpayers.

Concluding Remarks

This study has investigated econometrically the effects of a price support program and stocks (both private and public) on price dynamics and price volatility. A dynamic Tobit model was specified and estimated under time-varying volatility, where the price support provides a censoring mechanism for prices. The model is applied to the U.S. non-fat dry milk market.

The econometric analysis provides empirical evidence on the dynamics of non-fat dry milk prices and their changing volatility. First, we found evidence that both private and public stocks have significant effects on the reduction of price volatility. As expected, public stock accumulation contributes to market stabilization. Further, the negative effects of private stocks on price volatility are consistent with the literature on the economics of storage (e.g., Williams and Wright; Deaton and Laroque 1992, 1996).

Second, we document how price volatility has changed in the non-fat dry milk market over the last few decades. Based on our analysis, the period of market liberalization (the 1990s) has been associated with an increase in price volatility.

Third, the results provide evidence on the price stabilization effects of the price support program. The price support program has been effective in reducing short-term price volatility. However, the estimated price dynamics reveal that such price stabilization effects are short term and tend to disappear in the longer term. This finding emphasizes the need to investigate price stabilization policies in a dynamic context.

Fourth, the simulation results identify some important dynamic aspects of price adjustment in the U.S. non-fat dry milk market under market liberalization. As expected, increasing the price support raises expected price when the price support is set relatively high. But under the market regime scenario (where the support price is set below the market price), the analysis indicates the support price program can still affect expected price. While this effect is small in the short term, our findings demonstrate that the price support program can still contribute to significant increases in expected prices in the longer term. Therefore, it is possible for government policy to have long-term effects on market prices at a relatively low cost to taxpayers. Indeed, setting the support price below the long-term expected price would imply infrequent government purchases (only in situations of "unexpectedly low" prices). Yet the dynamic analysis suggests such a policy may still contribute to increasing the long-term expected price.

Because these findings were obtained in the context of the U.S. non-fat dry milk market, it is unclear whether similar results would hold in other markets. In particular, it would be of interest to examine the role of government in terms of changes in price floors across major dairy products (i.e., butter, American cheese, and non-fat dry milk). Finally, further research is needed to investigate the interactions among policy reform, price dynamics, and storage behavior.

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Appendix: The Variance of the Observed Market Price, y_t

Consider the standardized residual $\varepsilon_t = [y_t - f(\mathbf{X}_t, \beta)]/\sigma_t$. Using $h_t = [s_t - f(\mathbf{X}_t, \beta)]/\sigma_t$, we have

$$\text{(A1)} \qquad E(\varepsilon_t) = \left[E(y_t) - f(\mathbf{X}_t, \beta)\right]/\sigma_t = h_t * \Phi(h_t) + \phi(h_t)$$

from text equation (3a). In addition,

$$E(\varepsilon_t^2) = \int_{-\infty}^{h_t} h_t^2 \phi(u)\, du + \int_{h_t}^{\infty} \varepsilon_t^2 \phi(u)\, du.$$

From Maddala (p. 365), we have

$$\begin{aligned}\int_{h_t}^{\infty} \varepsilon_t^2 \phi(u)\, du &= \left[1 - \Phi(h_t)\right] * E\left[\varepsilon_t^2 | \varepsilon_t > h_t\right] = \left[1 - \Phi(h_t)\right] * \left[1 + h_t * E(\varepsilon_t | \varepsilon_t > h_t)\right] \\ &= \left[1 - \Phi(h_t)\right] * \left[1 + h_t * \phi(h_t)/(1 - \Phi(h_t))\right].\end{aligned}$$

It follows that

$$\text{(A2)} \qquad E(\varepsilon_t^2) = 1 - \Phi(h_t) + h_t * \phi(h_t) + h_t^2 * \Phi(h_t).$$

Using $V(y_t) = \sigma_t^2 * V(\varepsilon_t) = \sigma_t^2 * [E(\varepsilon_t^2) - (E(\varepsilon_t))^2]$, (A1) and (A2) yield text equation (3b).