Monetary Impacts on Prices in the Short and Long Run: Further Results for the United States

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This study examines the long-run neutrality of money and the short-run dynamics of farm and nonfarm prices to the monetary shock, using Johansen's approach. Results find a long-run equality of prices, but not neutrality. In the short run, farm prices adjust faster than nonfarm prices to a monetary shock.

Key words: causality, cointegration, error correction model, impulse responses, long-run equality of prices, maximum likelihood estimation, neutrality of money.

Introduction

The change in the U.S. exchange rate system in 1973 brought considerable attention to the macroeconomics of agriculture, especially as to how monetary policy affects farm prices compared to nonfarm prices. Tweeten, and Hughes and Penson argued that an expansionary monetary policy decreases the farm/nonfarm price ratio (cost-price squeeze), while Chambers (1981) and Chambers and Just argued that the policy increases the price ratio (cost-price expansion). On the other hand, Belongia, and Belongia and King concluded that the monetary policy has no impact on the relative prices (equality). Three researchers (Bordo; Frankel; and Rausser) took an intermediate position that expansionary monetary policy favors agriculture in the short run, because of sticky industrial prices and flexible farm prices, but that the policy is neutral in the long run after complete price adjustments have occurred throughout the economy (fix-price/flex-price).

Only a few agricultural economists have used unconditional vector autoregression (VAR) models to test the issue (e.g., Bessler; Chambers 1984; Devados and Meyers; Orden 1986a, b; Orden and Fackler; Taylor and Spriggs). Dynamic impulse responses of the VAR models have produced a consistent result for U.S. and Canadian data in the short run that farm prices adjust faster than nonfarm prices to monetary shock. However, in the long run, the VAR models did not produce a consistent explanation. Both Bessler and Orden (1986b) found a decrease in the farm/nonfarm price ratio to positive monetary shock, while Devados and Meyers found an increase in the ratio.

Robertson and Orden blamed the negligence of imposing the long-run behavior of money and prices in the VAR models for the inconsistency and discredited the VAR analysis. Applying the two-step procedure of Engle and Granger, they found the long-run neutrality of money[1] in New Zealander data and used it to build an error correction model (ECM). Thus, the credibility of their results relies on the acceptance of the neutrality restriction. However, a legitimate test of the neutrality restriction is not found in their paper. Moreover, using the two-step method in a trivariate system does not assure the efficiency gain over the VAR model (Hoffman and Rasche).

This study examines the long- and short-run behavior of U.S. money, farm prices, and

nonfarm prices. Johansen's (1988, 1991) maximum likelihood approach, which provides consistent estimates and test procedures for the cointegrating relationship, is applied.

Results do not support the long-run neutrality of money as defined in Robertson and Orden. However, a long-run equality of prices where farm prices are proportional to nonfarm prices is found. In the short run, farm prices adjust faster than nonfarm prices to a monetary shock, which supports the fix-price/flex-price arguments.

In the next section of this article, the ECM used in Robertson and Orden is compared to Johansen's maximum likelihood approach. The third section provides empirical findings and policy implications. The final section includes a summary and conclusions.

Time Series Models of Money and Agricultural Prices

The trivariate VAR model often used to study the monetary impacts on agricultural prices is specified as:

$$(1) \qquad y_t = \sum_{i=1}^{k} C_i y_{t-i} + \mu + \epsilon_t; \qquad y_t = [M_t \, P_t \, F_t]',$$

where M_t, P_t, and F_t are money supply, industrial prices, and farm prices, respectively (Bessler; Devados and Meyers; Han, Jansen, and Penson). If the data in y_t are generated by nonstationary processes, problems arise in estimation and inference on the VAR reduced form (Phillips and Durlauf; Sims, Stock, and Watson). When all the variables in y_t contain a unit root, the VAR in first differences,

$$(2) \qquad \Delta y_t = \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \mu + \epsilon_t,$$

is appropriate where $\Delta y_t = y_t - y_{t-1}$.

While individual macroeconomic time series data have often shown stationarity after first differencing, it is also believed that many nonstationary series have a tendency to move together in the long run (Engle and Granger; Haslag and Slottje; Kunst; Stock and Watson). If economic variables are linked by a long-run equilibrium relationship, a linear combination of the variables should not drift too far apart over time and need not be differenced. This relationship, which Granger called cointegration, reduces the number of unit roots to fewer than the number of variables. The existence of cointegration implies a restriction on multivariate time series models.

To show this, take the first difference in equation (1), which results in

$$(3) \qquad \Delta y_t = \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \Gamma_k y_{t-k} + \mu + \epsilon_t,$$

where

$$\Gamma_i = \sum_{j=1}^{i} C_j - I, \qquad \text{and}$$

$$\Gamma_k = I - \sum_{i=1}^{k} C_i.$$

If y_t is stationary, Γ_k has full rank 3 and equation (3) is the same as equation (1). If Δy_t is stationary, Γ_k will have rank zero, and equation (3) will be the same as equation (2). Empirically, the individual money and price series are nonstationary, but certain linear combinations of the series with stationarity exist. In this case, Γ_k has rank r such that $0 < r < 3$, and is expressed as $\Gamma_k = \alpha \beta'$, where α and β are $(3 \times r)$ matrices. This is the case in which the number of unit roots driving the system y_t is greater than zero but less than the number of variables in the system.

The coefficient β represents the long-run relationship between variables and is called the cointegrating vector (Granger). $\beta' y_t$ represents deviations from the long-run relationship, and α represents the speed of adjustment to the deviations. Equation (3) is called the ECM (Engle and Granger). Imposing relevant cointegration restrictions during estimation is important to ensure an efficient estimation procedure and to allow the usual inference, using standard asymptotic distribution theory (Phillips).

Robertson and Orden applied the two-step procedure of Engle and Granger to New Zealander money and manufacturing and farm prices. The first step was to estimate the cointegrating relationships via OLS regression. The neutrality requires cointegrations between money and prices with unitary cointegrating parameters. The second step was to use the (stationary) residuals from the cointegrating regressions as substitutes for the unobservable $\beta' y_t$ in OLS estimation of equation (3). The approach is designed specially for the bivariate systems.

With no restriction on the coefficient β, Robertson and Orden found cointegration between money and nonfarm prices, between money and farm prices, and between two prices. With unitary restriction on the coefficient β, they found the cointegration between money and prices and between farm and nonfarm prices, which can be represented as

$$(4) \qquad M_t - F_t = z_{1t}$$

and

$$(5) \qquad P_t - F_t = z_{2t},$$

where z_{1t} and z_{2t} are stationary. Robertson and Orden identified the cointegrating equations (4) and (5) as the proof of the money neutrality.

However, the neutrality argument of Robertson and Orden is questionable since the unitary restriction on the cointegrating parameter β is not tested. Rejecting the stationary relationship between M and P with the unitary restriction is also critical. Subtracting (5) from (4) results in $M_t - P_t = z_{1t} - z_{2t}$. Since z_{1t} and z_{2t} are stationary, $z_{1t} - z_{2t}$, a linear combination of the two stationary variables, should be stationary.

We applied the maximum likelihood procedure developed by Johansen (1988, 1991) to test the cointegration and the restrictions on the cointegrating parameters. This approach involves first estimating two auxiliary equations:

$$(6) \qquad \Delta y_t = \sum_{i=1}^{k-1} a_i \Delta y_{t-i} + \mu + R_{0t}$$

and

$$(7) \qquad y_{t-k} = \sum_{i=1}^{k-1} b_i \Delta y_{t-i} + \mu + R_{mt},$$

using OLS, with $\mu \neq 0$ and orthogonal to α. The residuals, R_{0t} and R_{mt}, respectively, from the auxiliary equations are used to form the moment matrices,

$$(8) \qquad \hat{S}_{i,j} = \frac{1}{T} \sum_{i=1}^{T} R_{it} R_{jt}, \qquad (i, j = 0, m).$$

Estimates of the cointegrating parameter β can be obtained from the eigenvectors associated with the r largest eigenvalues obtained by solving

$$(9) \qquad |\lambda S_{mm} - S_{m0} S_{00}^{-1} S_{0m}| = 0$$

and the estimate of α is obtained by $\alpha = -S_{0m}\beta$.

With $\mu = 0$, the auxiliary equations are changed to

$$(10) \qquad \Delta y_t = \sum_{i=1}^{k-1} a_i \Delta y_{t-i} + R_{0t}$$

and

$$(11) \qquad y_{t-k} = \sum_{i=1}^{k-1} b_i \Delta y_{t-i} + R_{mt}.$$

With $\mu = \alpha \beta_0'$, where β_0 is the constant term in the cointegrating vector, the auxiliary equations are changed to

$$(12) \qquad \Delta y_t = \sum_{i=1}^{k-1} a_i \Delta y_{t-i} + R_{0t}$$

and

$$(13) \qquad \tilde{y}_{t-k} = \sum_{i=1}^{k-1} b_i \Delta y_{t-i} + R_{mt},$$

where $\tilde{y}_{t-k} = [y_{t-k} 1]'$.

The likelihood ratio test statistic for the hypothesis that the system of g variables contains at most r cointegrating vectors is

$$(14) \qquad LR_r = -T \sum_{j=r+1}^{g} \ln(1 - \hat{\lambda}_i),$$

where $\hat{\lambda}_i$ represents the $g-r$ smallest eigenvalues. The distribution of the statistic is variant to the meaning of the nuisance parameter μ and is provided in Johansen and Juselius and in Osterwald-Lenum.

The t-values are used to test whether μ differs significantly from 0, and the likelihood ratio statistic

$$(15) \qquad LR_\mu = -T \sum_{i=r+1}^{g} \ln(1 - \hat{\lambda}_i)/(1 - \hat{\lambda}_i)$$

is used to test whether μ represents a constant term in the cointegrating vector. The test statistic is distributed as χ^2_{g-r}.

To test restrictions on the cointegrating parameters such that $\beta = \beta^* \phi$, where β^* is a $(g \times s)$ and ϕ is an $(s \times r)$ matrix, the likelihood ratio statistic

$$(16) \qquad LR_\beta = -T \sum_{i=r+1}^{g} \ln(1 - \lambda_i^*)/(1 - \hat{\lambda}_i)$$

is used, where λ_i^* represents the r largest eigenvalues obtained by solving

$$(17) \qquad |\lambda \beta^{*'} S_{mm} \beta^* - \beta^{*'} S_{m0} S_{00}^{-1} S_{0m} \beta^*| = 0.$$

The test statistic is distributed as $\chi^2_{(g-r)}$; ϕ are eigenvectors of equation (17), corresponding to the largest r eigenvalues. The restricted cointegrating vector β can be recovered by $\beta^* \phi$.

The Johansen procedure has several distinctive advantages over the more common two-step procedure. First, Johansen's approach has an efficiency gain over the two-step estimators, since it accounts for the error structure of the underlying process with the consistent maximum likelihood procedure. Gonzalo presented Monte Carlo evidence that the Johansen method performs better than the two-step procedure. Second, Johansen's approach provides a unified framework to test and estimate the cointegrating vector (Kunst). Third, the researcher does not have to make arbitrary normalization, which causes the results to be variant in the two-step procedure (Dickey, Jansen, and Thornton). Fourth, the procedure appears to work well with a higher order model in which the two-step procedure had difficulties (Engle and Yoo; Stock).

Table 1.   Cointegration Test for Trivariate System

Test	Critical Values			LR_r
	10%	5%	1%	
$r \leq 2$	2.69	3.76	6.65	.76
$r \leq 1$	13.33	15.41	20.04	9.88
$r = 0$	26.79	29.68	36.65	31.69

Empirical Results

Seasonally adjusted quarterly data on money supply (in billion dollars) and implicit price deflators for the farm sector (1982 = 100) and nonfarm sector (1982 = 100) were collected from selected issues of the *Federal Reserve Bulletin* (Board of Governors of the Federal Reserve System) and *Survey of Current Business* (U.S. Department of Commerce) from 1948:3 through 1991:3. All variables are expressed in natural logarithms to stabilize their variances. Since the true data-generating process is unknown, five different unit root tests are applied to reduce the misspecification biases. The unit root tests of Dickey and Fuller, Stock and Watson, Dickey and Pantula, Phillips and Perron, and Schmidt and Phillips consistently suggest that all three variables are nonstationary with a single unit root (see Choe for details).

Cointegration and ECM

A VAR model was fitted to the data in levels and in first differences to find an appropriate lag structure. The Schwartz criteria (see Judge et al.) suggest two lags in levels and one lag in first differences. Sims' modified likelihood ratio test (see Judge et al.) accepted the three-lag model and rejected the four-lag model for the levels, and accepted the two-lag model and rejected the three-lag model for first differences. Thus, three lags ($k = 3$) are chosen for the ECM and the VAR in levels, and two lags are chosen for the VAR in first differences.

To test structural changes in data, the VAR in levels are estimated for three separate sample periods (48:3–73:1, 73:2–79:4, and 80:1–91:3). The Chow test (see Judge et al.) accepted a structural change in 1973 (exchange rate system change) but rejected a structural change in 1979 (Monetary Decontrol Act) at the .05 level. A dummy variable is introduced to account for the structural change in 1973.

Results from the Johansen test of cointegration among M, P, and F with no restriction on μ are shown in table 1. The hypothesis of no cointegration ($r = 0$) is rejected, and the hypothesis of at most one cointegration ($r \leq 1$) is accepted at the .05 level. Thus, a long-run equilibrium relationship among the three variables is established, indicating both VAR models in first differences and levels are not appropriate. The auxiliary equations with no restriction on μ were appropriate for the cointegration test. The hypothesis of $\mu \doteq \alpha\beta_0'$ is rejected, with $LR_\mu = 4.75$ at the .05 level; the hypothesis of $\mu = 0$ is also rejected since intercept terms in M and P equations are significant.

We examine further the hypothesis of a cointegration between two variables. Results of the cointegration test on three bivariate systems appear in table 2. Unrestricted auxiliary equations are used, and lag lengths are chosen, using the same procedure as in the trivariate system. The hypothesis of no cointegration is not rejected in two bivariate systems, M–P and M–F, at any significance level. Thus, the long-run money neutrality is violated. A cointegration between P and F is found at any significance level with cointegrating vector $\beta = [-6.65 \ \ 7.22]'$. Since the cointegration parameters are closer to each other, a restriction $\beta^* = [1 \ \ -1]'$ on the cointegrating vector is considered. Johansen's test accepts the restriction with $LR_\beta = .1619$, implying long-run equality between two prices.

Table 2. Cointegration Tests for Bivariate System

| Test | Critical Values | | | LR, | | |
	10%	5%	1%	M–P (lag = 3)	M–F (lag = 2)	P–F (lag = 3)
$r \leq 1$	2.69	3.76	6.65	.39	1.71	.28
$r \leq 0$	13.33	15.41	20.04	13.06	6.85	20.77

Now, the question is whether the cointegration in the bivariate system is the same as the one found in the trivariate system. $\beta^* = [0 \quad 1 \quad -1]$ is imposed on the cointegrating vector in the trivariate system. The restriction implies equality between prices and excludes money supply from the cointegrating relationship. The restriction is not rejected, with $LR_\beta = .5990$. Thus, the equilibrium relationship between prices is identified as the only cointegration among the three variables. The restricted cointegration relationship shown in table 3 is imposed via an ECM.

The estimates and evaluation statistics for the ECM are shown in table 4. R^2s are obtained after the model has been reparameterized to get an equivalent VAR in levels; the R^2s indicate that the model explains a significant proportion of the variation in dependent variables. A Q-test on the residuals indicates little evidence of residual auto-correlation, which suggests that the model does a good job of representing the autocor-relation structure of the variables. However, the standard Lagrange multiplier test indicates some evidence of ARCH effects in price equations. The constant term is significant in M and P equations, which confirms a linear trend in the nonstationary parts of the variables. The dummy variable is also significant in all three equations, which confirms the structural change in 1973. The error correction term is significant in P and F equations but not in the M equation, implying little response of M to the deviation in the long-run equality of prices.

Dynamics of Money and Prices

To detect the dynamic effects of various shocks, the ECM is reparameterized to its equiv-alent (restricted) VAR in levels. Recursive order used in Robertson and Orden, $M–P–F$, is applied to remove the contemporaneous correlation in error term. This order allows money supply shocks to affect the price variables contemporaneously (Robertson and Orden). The rationale for the ordering can be found in Orden (1986a, b) and in Orden and Fackler. The speed of adjustment parameter α also can be used to justify the structural form identification, following Engle and Granger. According to the estimated α, F adjusts over 30 times faster, and P adjusts four times faster than M to any deviation in equality of prices (table 3). Thus, the ordering of $M–P–F$ seemed the appropriate choice.

Impulse response (in percentage) over four-year periods to a 1% positive shock to each variable and their 90% confidence bounds are shown in figures 1–3. The Monte Carlo integration method with 500 draws has been used to compute the confidence interval for the posterior distribution of the impulse responses (Doan and Litterman; Kloek and Van Dijk).

Table 3. Estimated Cointegrating Vectors

| Variable | Cointegrating Vectors | |
	β	α ($\times$ 1000)
M	0	−.41
P	6.75	1.69
F	−6.75	−12.78

Table 4. Estimates and Evaluation Statistics for the ECM

Dependent Variable	Γ									Γ_k		Statistics		
	μ	DV	ΔM_{t-1}	ΔM_{t-2}	ΔP_{t-1}	ΔP_{t-2}	ΔF_{t-1}	ΔF_{t-2}	P_{t-3}	F_{t-3}	R^2	Q_{40}	ARCH	
ΔM_t	.0042 (3.49)	.0033 (2.07)	.5302 (6.64)	.0163 (.20)	.0094 (.10)	.0265 (.27)	−.0142 (−1.53)	−.0003 (−.03)	−.0028 (−.71)	.0028 (.71)	.9999	29.56 {.89}	33.93 {.74}	
ΔP_t	.0016 (1.83)	.0670 (1.17)	.0364 (.63)	.1856 (2.63)	.2729 (3.91)	.0198 (2.97)	.0111 (1.67)	.0114 (4.07)	−.0114 (−4.07)	.0030 (2.67)	.9999	45.61 {.25}	87.36 {.00}	
ΔF_t	.0017 (.16)	.4495 (.66)	1.368 (.20)	1.9483 (2.34)	−.0640 (−.08)	−.0383 (−.49)	−.1196 (−1.52)	−.0863 (−2.60)	.0863 (2.60)	−.2534 (−1.89)	.9836	52.92 {.08}	84.83 {.00}	

Notes: Figures in parentheses ( ) indicate t-values; figures in braces { } indicate probability values.
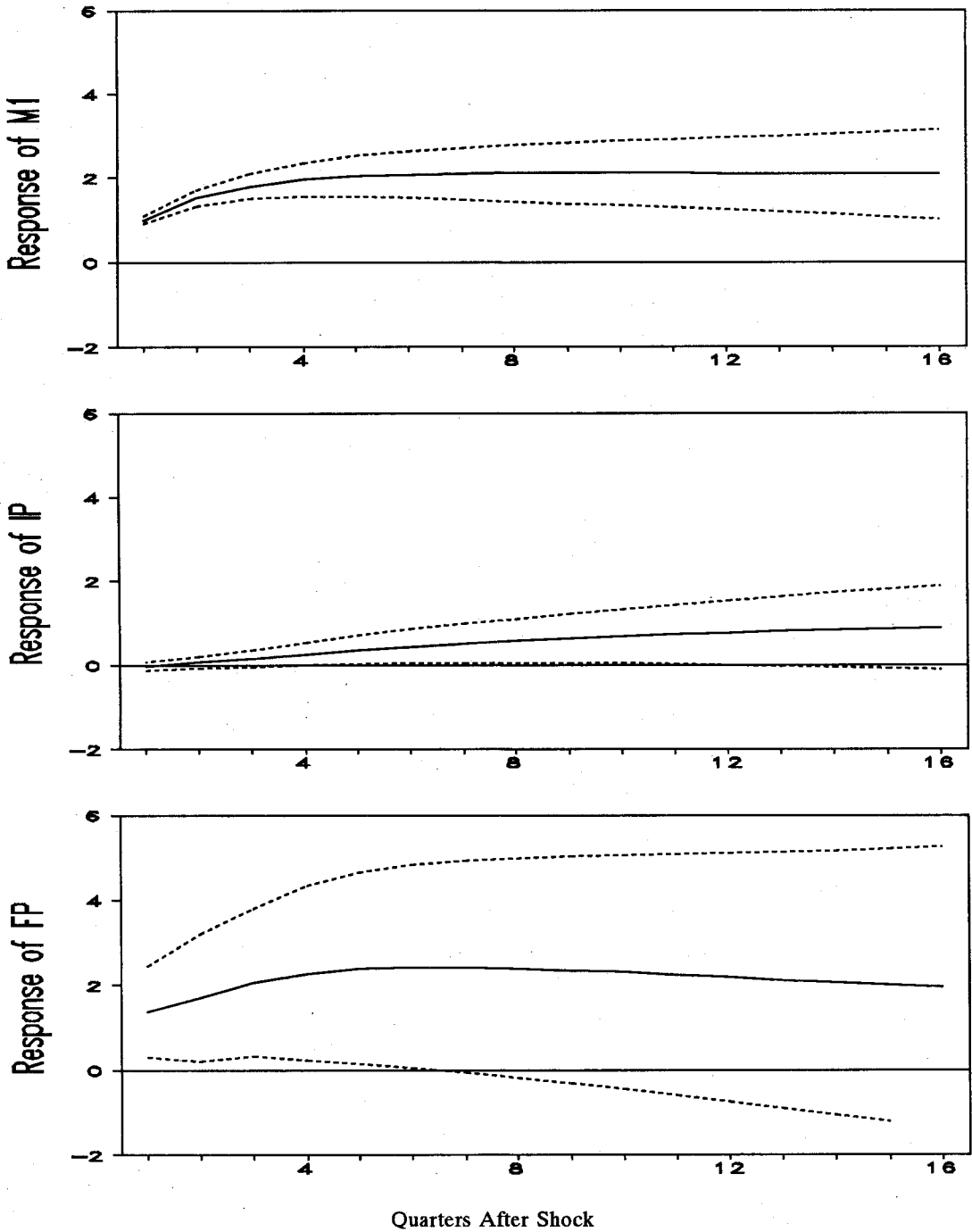
Figure 1.   Impulse responses (%) to 1% shock in M

   The long-run point estimates of two price responses are approximately equal to any
shock after about 60 quarters. However, the mean responses of the prices have not
converged to the level of the money response. This reflects the long-run equality of prices
and non-neutrality of money imposed by the cointegrating relationship. We are not sure
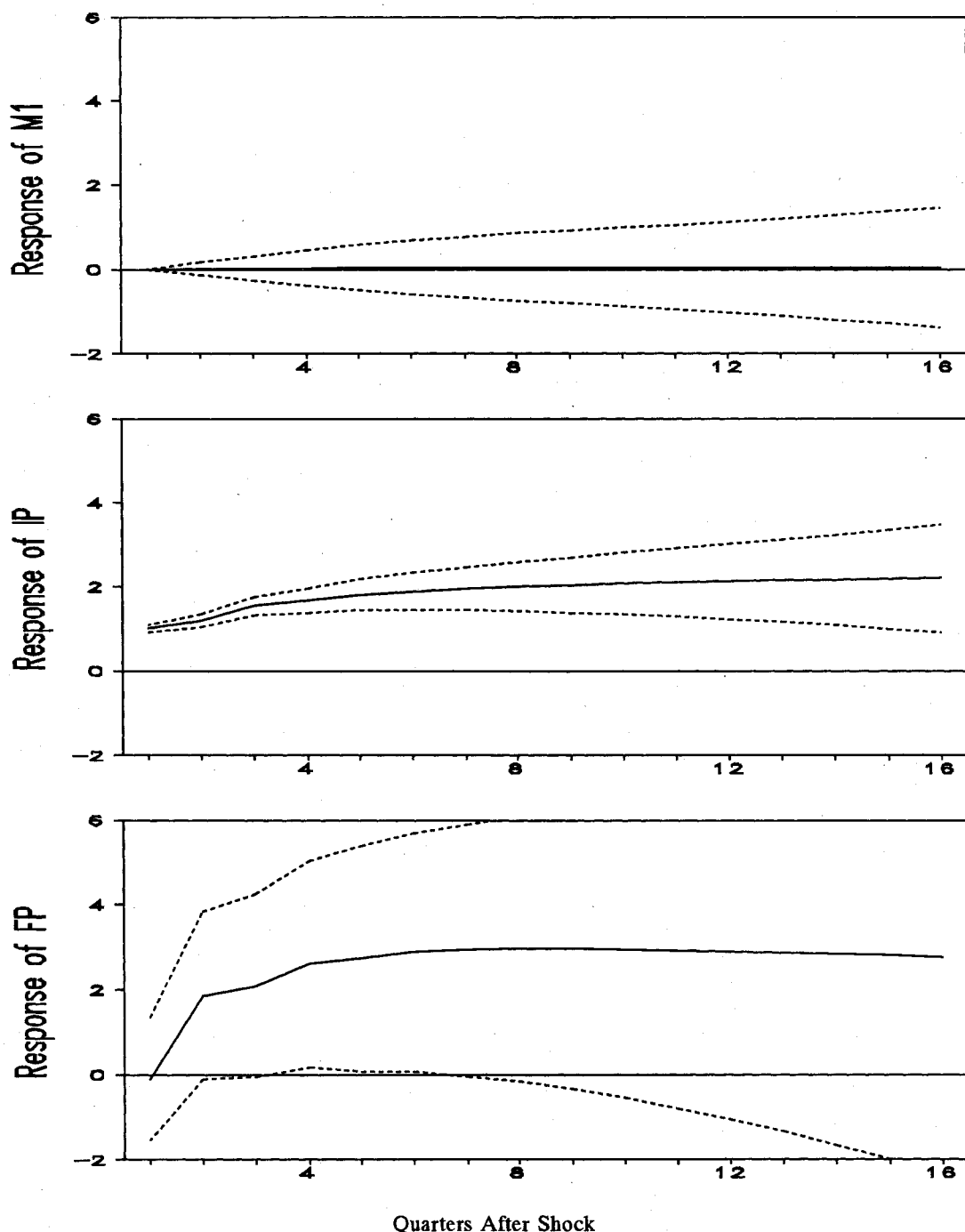whether it takes about 60 quarters to reach a new equilibrium in the prices, because the

Quarters After Shock

Figure 2. Impulse responses (%) to 1% shock in P

impulse responses become insignificant after about two to five years. The adjustment
period may vary, depending on the type of shocks in the economy. A new equilibrium
could be reached within two years in the case of monetary shock, within three years in
the case of farm price shock, and within five years in the case of industrial price shock.
    As shown in figure 1, the short-run overshooting of farm prices to monetary shock is
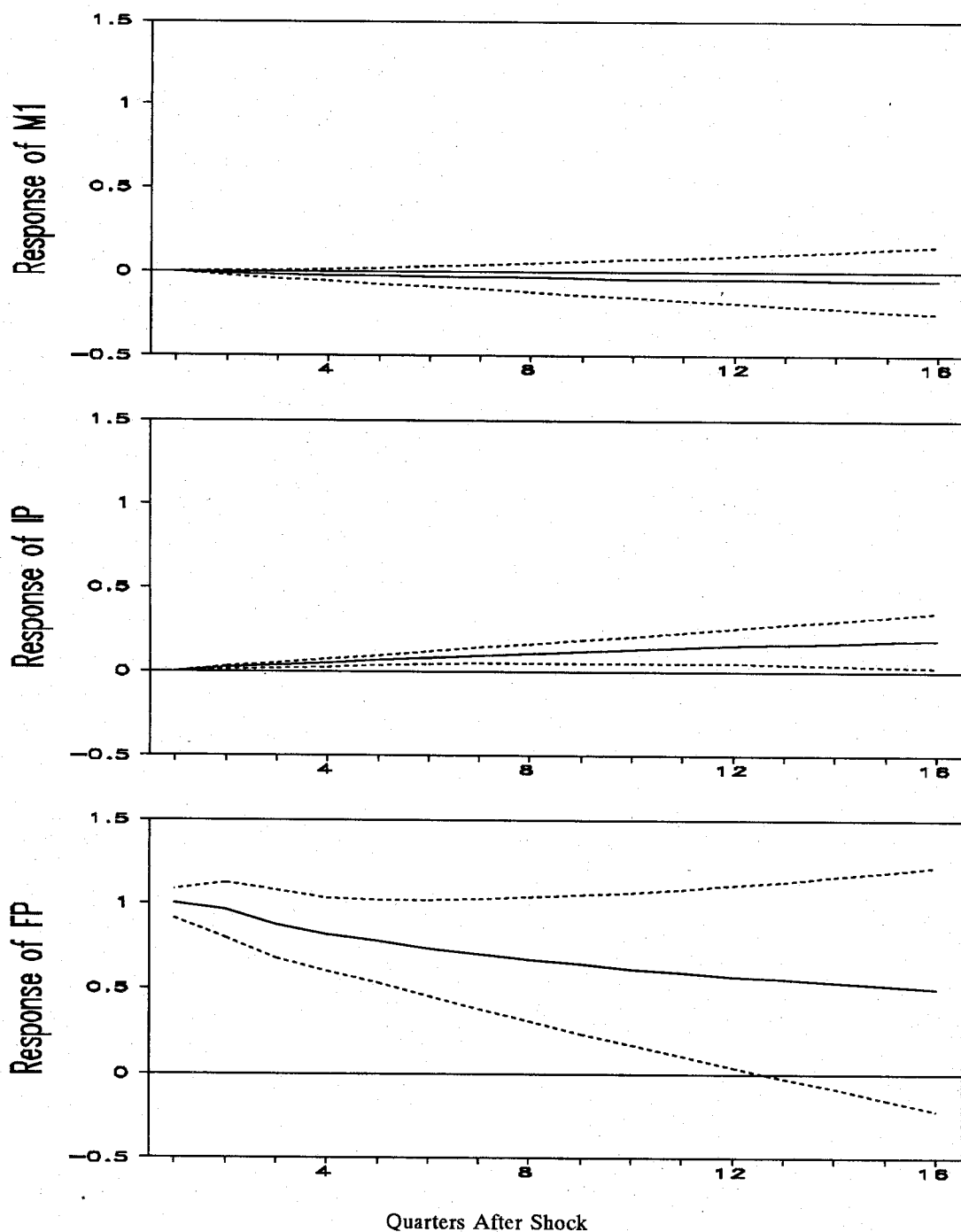
Quarters After Shock

Figure 3.   Impulse responses (%) to 1% shock in F

apparent. The mean response of farm prices is immediate and reaches up to the 2.5%
level, where money supply reaches approximately the 2% level, and nonfarm prices stay
below the 1% level. The results support the fix-price/flex-price argument. The response
of both prices is marginally significant, with their lower confidence bounds staying around
zero.

A positive shock to industrial prices also raises farm prices more than industrial prices (fig. 2). However, the mean response of farm prices is not significant statistically, while response of industrial prices is significant. Thus, whether the unexpected inflation causes an overshooting of farm prices or a cost-price squeeze in the farm sector is inconclusive. The results will depend on the economy at the moment. When the response of farm prices is detected near the lower confidence bound, one may conclude that autonomous inflation causes the cost-price squeeze (Orden 1986a, b; Penn; Tweeten). With the response of farm prices near the upper bound, one also can conclude that the autonomous inflation causes a farm price overshooting (Starleaf; Starleaf, Meyers, and Womack).

A positive shock to farm prices raises farm prices initially but does not significantly affect money supply and industrial prices (fig. 3). Compared to other shocks, the mean responses of all three variables, especially the macroeconomic variables, are smaller in the case of farm price shock. The mean responses of M and P to farm price shock are less than .25%, while the responses of M and P to other shocks range from 1 to 2%. The results are consistent with other empirical studies where macroeconomic variables changed agricultural prices but not vice versa (Barnett, Bessler, and Thompson; Han, Jansen, and Penson; Orden 1986a, b; Saunders). Thus, the sequential ordering of M–P–F is supported again.

Sensitivity Analysis

To compare the results shown in figures to impulse responses with the different Choleski decompositions, triangularization orderings of M–F–P and F–P–M used in Bessler are considered. The point estimates of the impulse responses are found insensitive to the orderings. The point estimates of the impulse responses with alternative orderings (not shown but available upon request) are almost identical to those with M–P–F ordering. The only difference is that F–P–M ordering produced slightly smaller responses of farm prices to monetary shock compared to alternative orderings. The differences are expected since F–P–M ordering imposes a restriction that the response of farm prices has to start off in the quarter following the monetary shock, where the other orders allow the farm prices to respond immediately to monetary shock. The difference, however, does not change the impulse response pattern of the farm prices to monetary shock.

Sensitivity of the results with respect to the different model specifications is also considered. The VAR in levels produced similar point estimates of the impulse responses to the ECM, but both prices do not converge to common values even after 60 quarters. The VAR in first differences produces quite different estimates of impulse responses. The point estimates from the difference model show that all responses are immediate and permanent, implying all three series contain a unit root. The results are not surprising since a nonstationarity restriction ($\Gamma_k = 0$) is imposed on the VAR model. A cyclical response pattern in farm prices suggests that the difference model suffers overdifferencing of series by ignoring an existing cointegration.

The difference model also results in more overshooting of farm prices in the case of monetary and farm price shock than the other models. The mean responses of farm prices are 10 times bigger than those of industrial prices in the case of monetary shock, and 20 times bigger in the case of farm price shock. Moreover, the response estimates show a big cost-price squeeze in the farm economy to industrial price shock. Nonfarm prices increase about seven times more than farm prices to a nonfarm price shock. Thus, inferences drawn from the VAR model in first differences would differ from those of the ECM or the VAR in levels.

The results support an argument raised by Engle and Granger (p. 259) that vector autoregressions estimated with cointegrated data will be misspecified if the data are differenced. Also, they will have omitted important constraints if the data are used in levels. The constraints will be satisfied asymptotically, but efficiency gains and improved multistep forecasts may be achieved by imposing them.

Concluding Comments

This article has examined the long-run neutrality of money and the short-run dynamics of prices for the U.S. economy. Johansen's maximum likelihood method does not find a long-run neutrality of money as defined in Robertson and Orden. However, a stable equilibrium relationship in which farm prices are proportional to nonfarm prices in the long run is found. By restricting the VAR model with the long-run relationship, an ECM is estimated. Though the VAR in levels performed similarly to the ECM, the long-run property of the prices is not found in the VAR. The VAR in differences performed differently from the ECM.

Impulse responses from the ECM found that a positive monetary shock causes the short-run overshooting of the farm prices. The mean responses of the farm prices are immediate and bigger than those of the industrial prices. The long-run and short-run results together support the fix-price/flex-price argument. However, evidence is insufficient to decide whether there is a cost-price squeeze or a farm price overshooting after autonomous inflation in the nonfarm sector. The explanations of inflation impact on farm prices may differ, depending on the economy at hand. The impulse responses also suggest that macroeconomic variables change farm prices but not vice versa.

The impulse responses and the long-run equality of prices provide important policy implications. In the short run, the monetary policy will change the relative price, since it will initiate the shock to the money supply. The farm policy also will affect the relative price in the short run, since it will initiate the shock to the farm prices. The long-run equality of prices will be restored unless the policy shock is unprecedented, since farm prices move proportionally to nonfarm prices in the long run. Though the monetary policy shock has bigger and more persistent effects on farm prices than the farm price shock has, it also has significant impact on the general price level.

A flexible farm policy can reduce the short-run deviations of farm prices from the long-run equilibrium price level. An inflexible farm policy would hurt the long-run equality between two prices and, hence, would introduce more instability into the farm economy. When an inflation in the nonfarm sector causes a cost-price squeeze in the farm sector, a farm price support is recommended. When an inflation in the nonfarm sector causes a farm price overshooting, any farm price support should be removed to prevent overallocation of resources into the farm sector.

Note

[1] Robertson and Orden defined the long-run neutrality of money as a stable equilibrium relationship that both manufacturing and agricultural prices are proportional to money.

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