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Won W. Koo

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Abstract

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 stable proportional relationship between prices but not the neutrality. In the short run, farm prices adjust faster than nonfarm prices to a monetary shock.

MONETARY IMPACTS ON PRICES IN THE SHORT AND LONG RUN: FURTHER RESULTS FOR THE UNITED STATES

Young Chan Choe and Won W. Koo*

The change in the U.S. exchange rate system in 1973 brought considerable attention to the macroeconomics of agriculture, especially on 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 argued that the monetary policy has no impact on the relative prices (equality). 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).

Few agricultural economists have used unconditional vector autoregression (VAR) models to test the issue (Bessler; Chambers, 1984; Orden 1986a, 1986b; Orden and Fackler; Devados and Meyers; 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. Orden (1986b) and Bessler found a decrease in the farm/nonfarm price ratio to positive monetary shock while Devados and Meyers found an increase in the ratio.

*Research specialist and professor, respectively, Department of Agricultural Economics, North Dakota State University, Fargo.

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 stable proportional relationship between prices is found. In the short run, farm prices adjust faster than nonfarm prices to a monetary shock, which supports Rausser's fix-price flex-price arguments.

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

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[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.

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

$$y_t = \sum_{i=1}^{k} C_i y_{t-i} + u_t; \qquad y_t = [M_t \; P_t \; F_t]' \tag{1}$$

where M_t, P_t, and F_t are money supply, industrial prices, and farm prices, respectively (Bessler; Devados and Meyers; Han et al.). With a nonstationary y_t, problems arise in estimation and inference on the VAR reduced form (Phillips and Durlauf; Sims et al.).

By rearranging, equation (1) can be written as

$$\Delta y_t = \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} - \Gamma_k y_{t-k} + u_t \tag{2}$$

where $\Gamma_i = \sum_{j=1}^{i} C_j - I$ and $\Gamma_k = I - \sum_{i=1}^{k} C_i$. Therefore, the VAR in levels is only appropriate when Γ_k has a full rank 3 and y_t is stationary. If Γ_k has rank zero, Δy_t is stationary, and the VAR in first differences is appropriate. 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 (3xr) 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 ß 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 (2) 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 (2). The approach is designed specially for the bivariate systems.

With no restriction on the coefficient ß, 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 ß, they found the conitegration between money and prices and between the farm and nonfarm prices, which can be represented as

$$M_t - F_t = z_{1t} \tag{3}$$

and

$$P_t - F_t = z_{2t} \tag{4}$$

where z_{1t} and z_{2t} are stationary. Robertson and Orden identified the cointegrating equations (3) and (4) 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 ß is not tested. Rejecting the stationary relationship between M and P with the unitary restriction is also critical. Subtracting (4) from (3) 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 apply 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, Δy_t and y_{t-k}, on (Δy_{t-1}, ..., Δy_{t-k+1}, μ_t) by OLS. The residuals R_{0t} and R_{mt}, respectively, from the auxiliary equations are used to form the moment matrices

$$\hat{S}_{i,j} = \frac{1}{T}\sum_{i=1}^{T} R_{it}R_{jt}. \qquad (i,j=0,m) \tag{5}$$

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

$$\left|\lambda S_{mm} - S_{m0}S_{00}^{-1}S_{0m}\right| = 0 \tag{6}$$

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

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

$$LR_r = -T\sum_{j=r+1}^{g} \ln(1 - \hat{\lambda}_i) \tag{7}$$

where $\hat{\lambda}_i$ are the g-r smallest eigenvalues. The distribution of the statistic is variant to the meaning of the nuisance parameter μ. With μ=0, the nuisance parameter does not appear in the ECM and the auxiliary equations. With μ≠0 and orthogonal to α, the nuisance parameter appears in the ECM and the auxiliary equations, representing a linear trend in the nonstationary part of y_t. With $\mu=\alpha\beta_0'$, the nuisance parameter appears in the ECM, representing the constant term β_0' in the cointegrating vector. The auxiliary equations in this case are changed to Δy_t and $\tilde{y}_{t-k}$ on $(\Delta y_{t-1}, \Delta y_{t-2}, \ldots, \Delta y_{t-k+1})$ where $\tilde{y}_{t-k} = [y_{t-k}\ 1]'$. Eigenvalues $\hat{\lambda}_i$ and eigenvectors are obtained using the same procedure. Johansen and Juselius and Osterwald-Lenum developed appropriate critical values for test statistics in each case.

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

$$LR_\mu = -T\sum_{i=r+1}^{g} \ln(1-\hat{\lambda}_i)/(1-\hat{\lambda}_i) \tag{8}$$

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 (gxs) and ϕ is a (sxr) matrix, the likelihood ratio statistic

$$LR_\beta = -T\sum_{i=r+1}^{g} \ln(1-\lambda_i^*)/(1-\hat{\lambda}_i) \tag{9}$$

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

$$|\lambda\beta^{*\prime}S_{mm}\beta^* - \beta^{*\prime}S_{m0}S_{00}^{-1}S_{0m}\beta^*| = 0. \tag{10}$$

The test statistic is distributed as $\chi^2_{(g-s)r}$. ϕ are eigenvectors of (10) 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 (Johansen). Gonzalo gave 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 et al.). Fourth, the procedure appears to work well with a higher order model in which the two-step procedure had difficulties (Engle and Yoo; Stock).

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 and Survey of Current Business from 1948:3 to 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 (p=.03740) and rejected the four-lag model (p=.50732) for the levels and accepted the two-lag model (p=.00019) and rejected the three-lag model (p=.40418) 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, 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 ($r\leq1$) 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 $\mu=\alpha\beta_0'$ is rejected with $LR_\mu=4.75$ ($p=.02933$) at the .05 level, and 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 cointegrating 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_β =.1619 (p=.6873), implying the long-run proportionality between two prices.

Now, the question is whether the cointegration in the bivariate system is the same as the one found in the trivariate system. $\beta^*=[0\ 1\ -1]$ 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_β=.5990 (p=.7411). 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 estimated ECM are shown in Table 4. R^2s are obtained after the model has been reparameterized to get an equivalent VAR in levels and indicate the model explains a significant proportion of the variation in dependent variables. Q test on the residuals indicates little evidence of residual autocorrelation, which indicates that the model does a good job of representing the autocorrelation 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 part 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 proportional price relationship.

Dynamics of Money and Prices

To detect the dynamic effects of various shocks, the ECM is reparameterized to its equivalent (restricted) VAR in levels. Recursive order used in Robertson and Orden, M-P-F, is applied to identify the structural form. 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, 1986b) and 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 4 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 to 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 (Kloek and Van Dijk; Doan and Litterman).

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 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 case of monetary shock, within three years in case of farm price shock, and within five years in case of industrial price shock.

As shown in Figure 1, the short-run overshooting of farm prices to monetary shock is apparent. The mean response of farm prices is immediate and reaches up to the 2.5% level where money supply reaches about the 2% level and nonfarm prices stay below the 1% level. The results support Rausser's fix-price flex-price argument. The response of both prices are marginally significant with their lower confidence bounds staying around zero.

A positive shock to industrial prices also raise farm prices more than industrial prices (Figure 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 the autonomous inflation causes the cost price squeeze

(Tweeten; Penn; Orden 1986a, 1986b). With the response of farm prices near the upper bound, one also can conclude that the autonomous inflation causes a farm price overshooting (Starleaf et al.; and Starleaf).

A positive shock to farm prices raises farm prices initially but does not significantly affect money supply and industrial prices (Figure 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 (Orden; Barnett et al.; Saunders; Han et al.). Thus, the sequential ordering of M-P-F is supported again.

Policy Implication and Model Comparison

The impulse responses and the long-run equality of prices provide important policy implications. Neither monetary nor farm policy would permanently change relative price as both farm nad nonfarm prices move proportionally in the long run. 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 monetary policy would have bigger and more persistent effects on farm prices than the farm policy, since the farm prices are responding more to the money shock. The monetary policy also has more significant impact on the general price level than the farm policy has. Thus, it is much safer to use the farm policy to boost farm prices when only the farm sector experiences a short period of financial stress.

As Rausser pointed out, a flexible farm policy should be practiced to reduce the short-run deviations of farm prices from the long-run equilibrium price level. An inflexible farm policy would hurt the long-run proportional relationship 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 the farm price overshooting, any farm price support should be removed to prevent over-allocation of resources into the farm sector.

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 (Γ_k=0) is imposed on the VAR model. A cyclical response pattern in farm prices suggests that the difference model suffers over-differencing 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 7 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.

Concluding Comments

This paper 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 doesn't 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 similar 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 the short-run overshooting of the farm prices to a positive monetary shock. 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 Rausser's fix-price flex-price argument. However, evidence is insufficient to decide whether there will be 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 the hand. The impulse responses also prove that macroeconomic variables change farm prices but not vice versa.

The flexible farm price policy is recommended to reduce the short run instability in the farm prices. The policy should not hurt the long run proportional relationship between the farm and the nonfarm prices.

Table 1. Cointegration Test for Trivariate System

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

Table 2. Cointegration Tests for Bivariate System

Test	Critical Values			LR_r		
	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

Table 3. Estimated Cointegrating Vectors

Variable	Cointegrating Vectors	
	ß	α(x1000)
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	.0033	.5302	.0163	.0094	.0265	-.0142	-.0003	-.0028	.0028	.9999	29.56	33.93
	(3.49)	(2.07)	(6.64)	(.20)	(.10)	(.27)	(-1.53)	(-.03)	(-.71)	(.71)		{.89}	{.74}
ΔP_t	.0016	.0670	.0364	.1856	.2729	.0198	.0111	.0114	-.0114	.0030	.9999	45.61	87.36
	(1.83)	(1.17)	(.63)	(2.63)	(3.91)	(2.97)	(1.67)	(4.07)	(-4.07)	(2.67)		{.25}	{.00}
ΔF_t	.0017	.4495	.1368	1.9483	-.0640	-.0383	-.1196	-.0863	.0863	-.2534	.9836	52.92	84.83
	(.16)	(.66)	(.20)	(2.34)	(-.08)	(-.49)	(-1.52)	(-2.60)	(2.60)	(-1.89)		{.08}	{.00}

( ) indicates t values.
{ } indicates probability value.

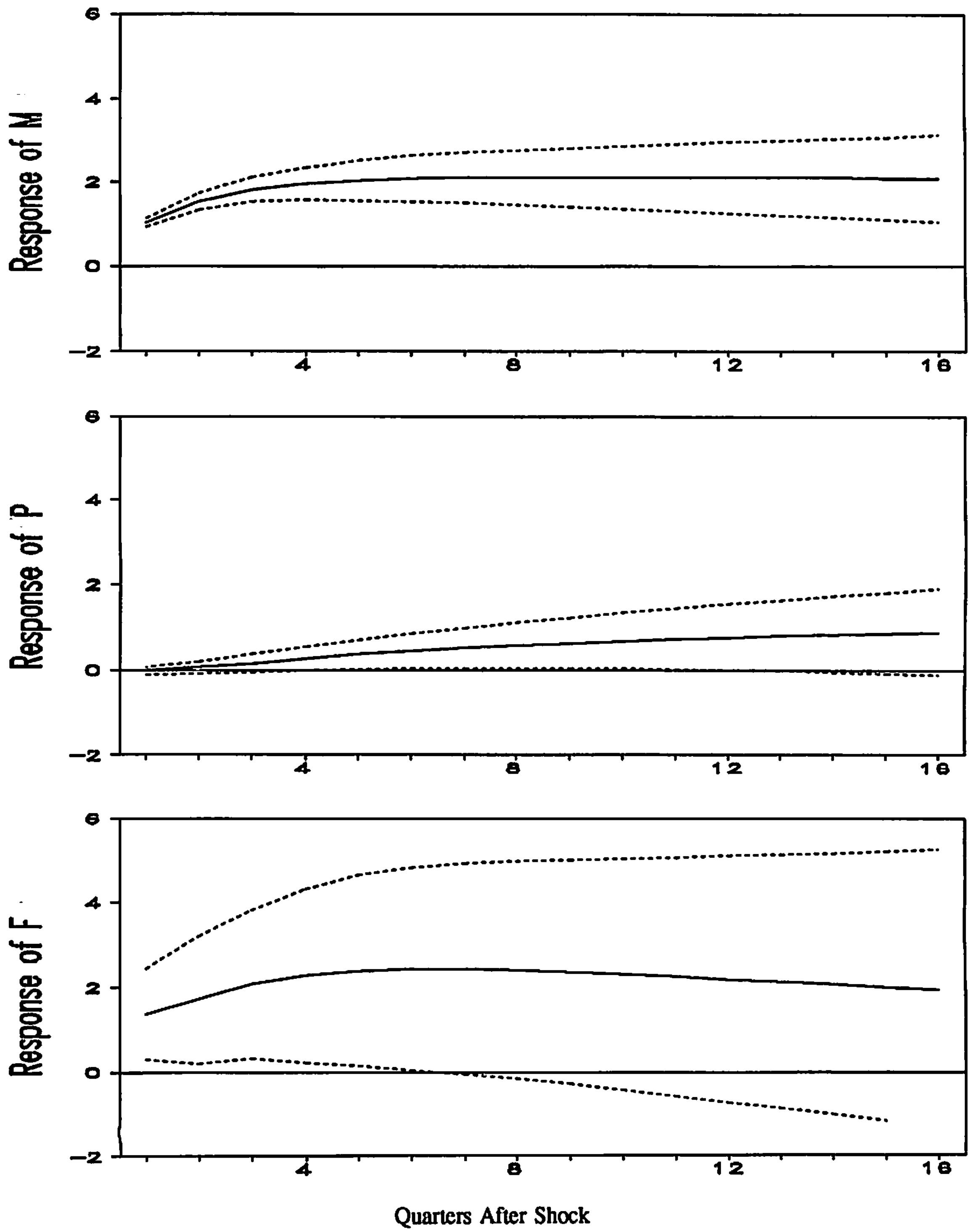

Figure 1. Impulse responses (%) to 1% shock in M

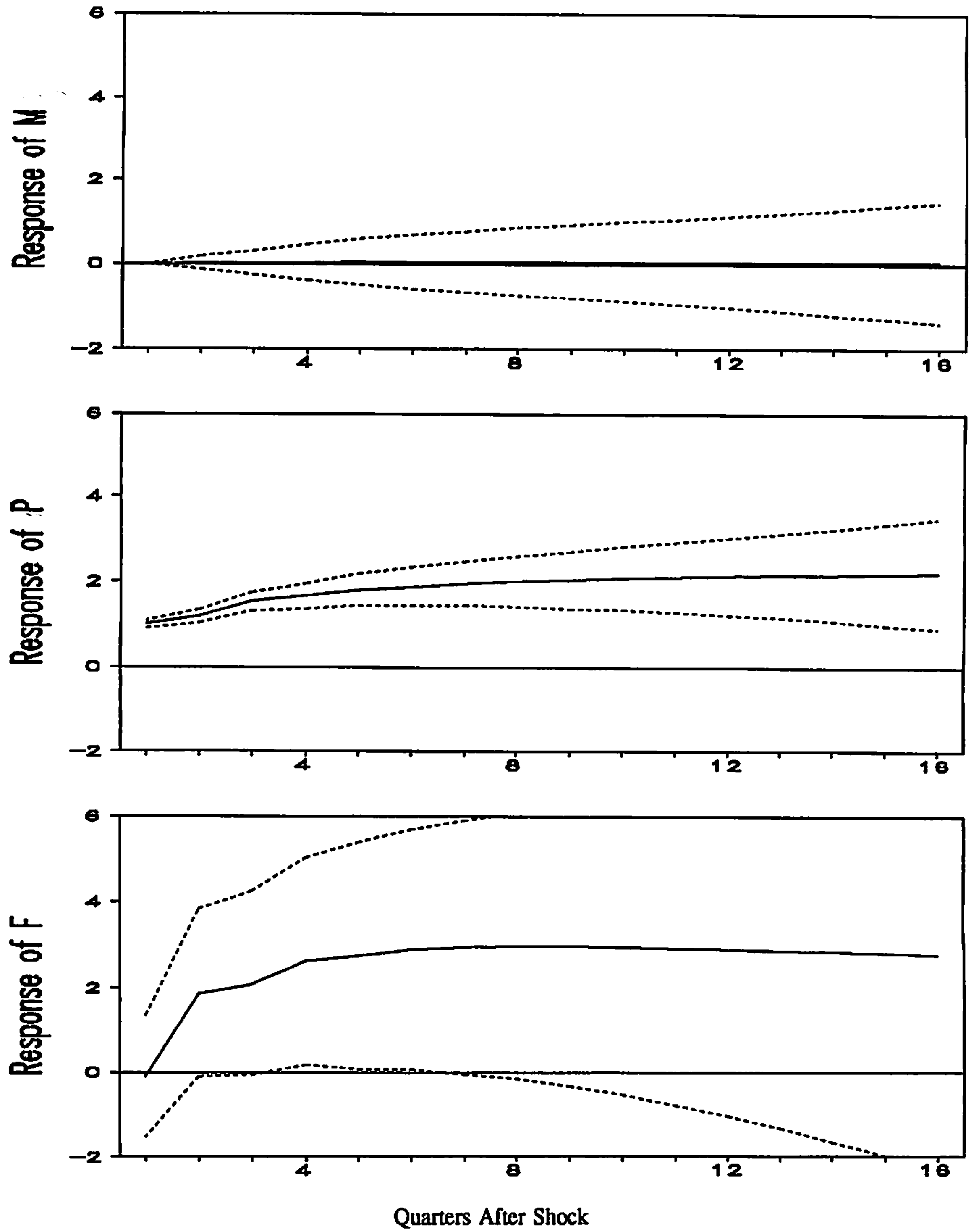

Figure 2. Impulse Responses (%) to 1% shock in P

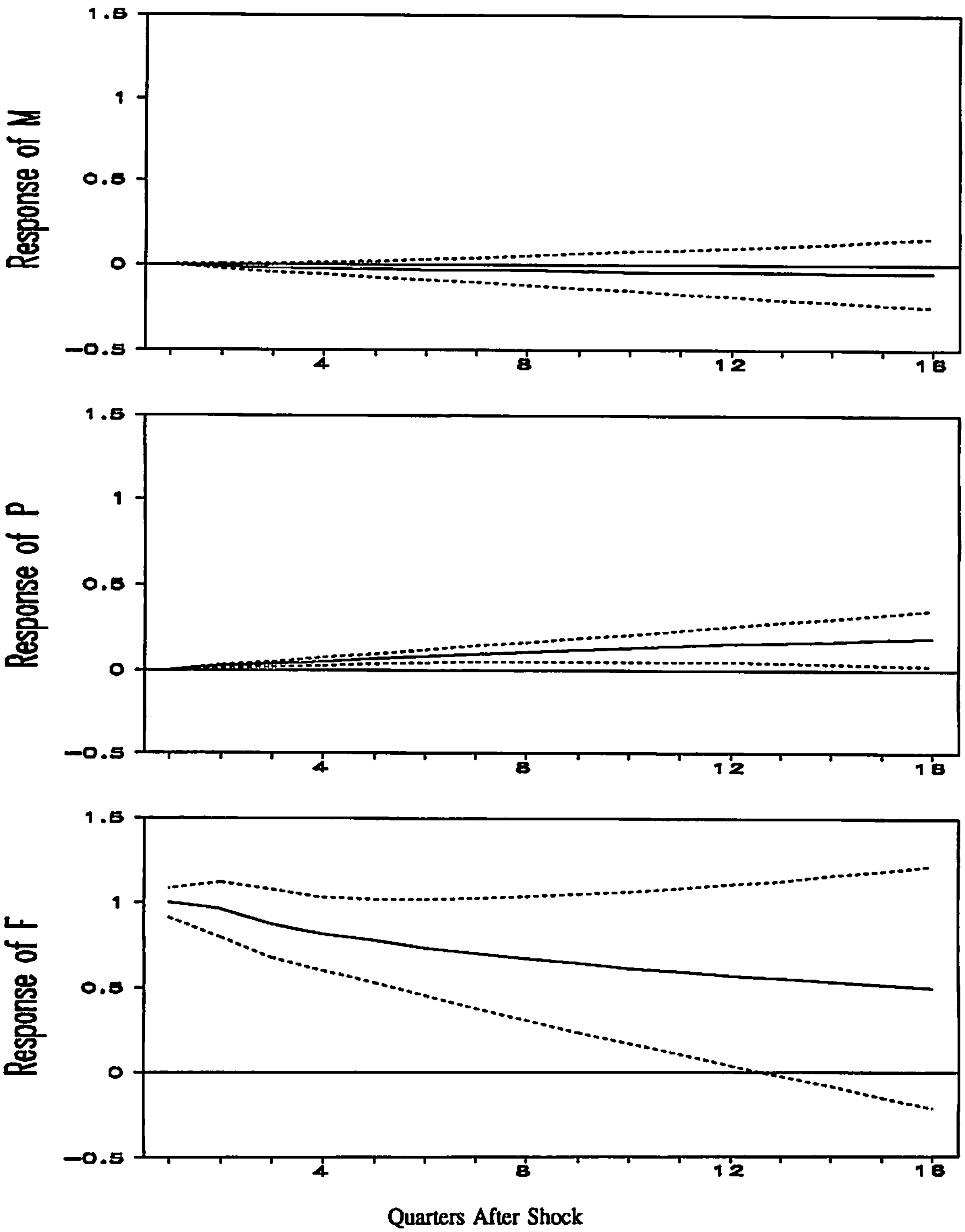

Figure 3. Impulse Responses (%) to 1 % shock in F

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