COINTEGRATION AND LONG-RUN MONETARY NEUTRALITY:

A VECTOR ERROR-CORRECTION MODEL OF MONEY AND PRICE DYNAMICS IN NEW ZEALAND

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Introduction

In world terms New Zealand is a small and agriculturally trade-dependent nation. However, because the New Zealand economy also contains a well developed manufacturing base (contributing some 25 percent of GDP in 1986), the dynamic relationships between the farming and manufacturing sectors, and among monetary policy and these sectors, are important. With a large farming sector (around 20 percent of GDP including agricultural processing, and 55 percent of current account receipts in 1986) one might expect to observe somewhat different interdependences at the aggregate level than, say, in the United States.

In this regard, we investigate four related questions concerning money and prices in New Zealand: 1) Do shocks to the money supply generate responses in the time path of relative prices?; 2) Do agricultural prices and the money supply respond to autonomous manufacturing price shocks?; 3) Is there feedback from agricultural price shocks to the money supply or manufacturing prices?; and 4) Are there long-run equilibrium relationships

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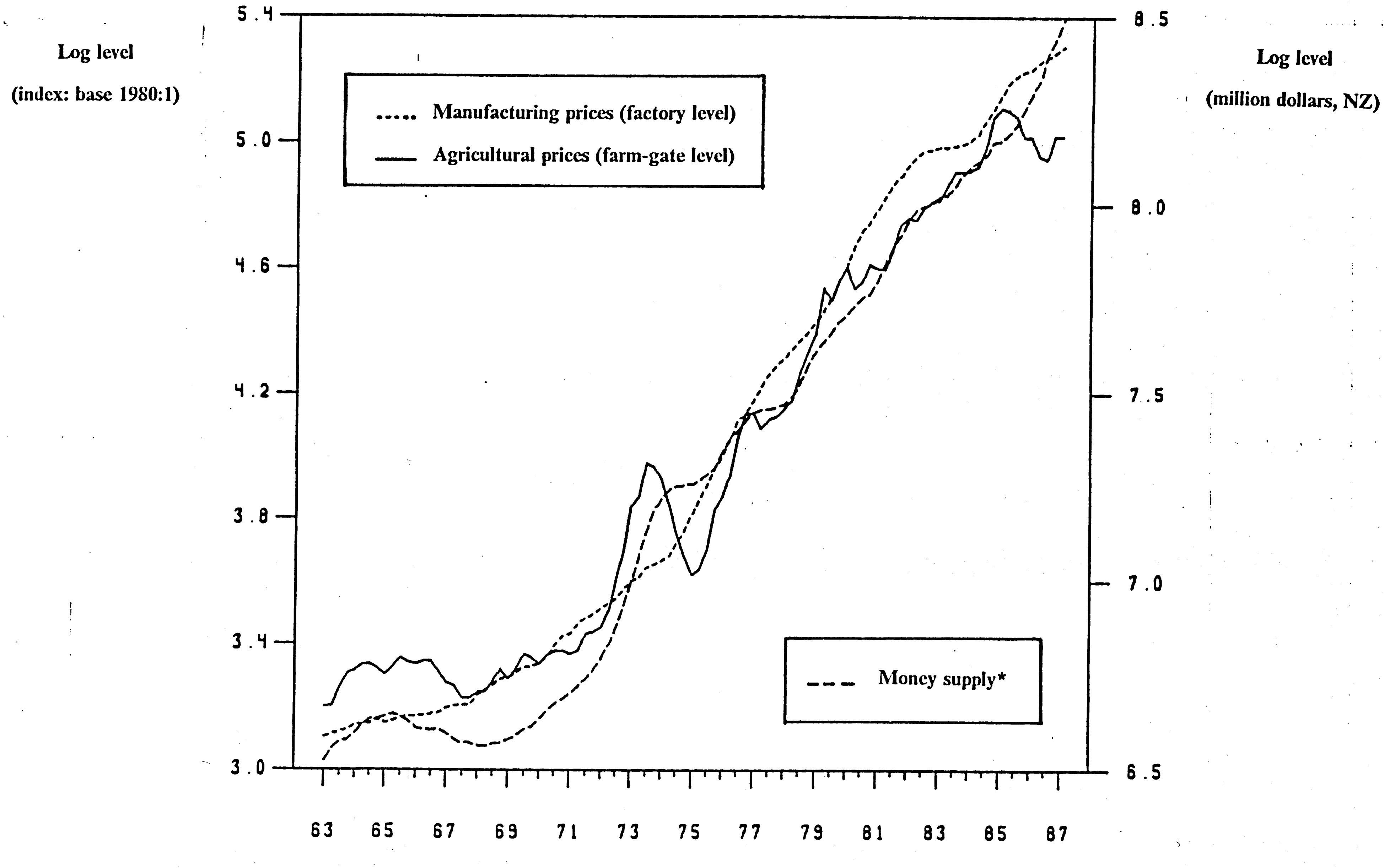
manufacturing prices?; and 4) Are there long-run equilibrium relationships among output prices (real activity) and the money supply (monetary policy)?

To address these four questions we utilize a three-variable dynamic model of the money supply, agricultural prices, and manufacturing prices. Though a three-variable model is an abstract representation of the economy, interactions among these variables encompass many of the concerns expressed about linkages among sectors and macroeconomic policy. In New Zealand, domestic monetary policy and exchange-rate policy are closely coordinated and the latter has to some extent accommodated the former through nominal exchange-rate adjustments that have maintained a relatively stable real exchange rate (Rayner and Lattimore). The agricultural sector is primarily a price taker in international markets generally receiving low levels of direct producer assistance (Lattimore). The manufacturing sector has received relatively high levels of import protection and there is evidence that it has been able to administer cost-plus-markup pricing policies (Chapple). These characteristics of the economy may have implications for short-run interactions among money and prices. In the long run, however, as Batten and Belongia, Chambers and others have emphasized, the hypothesis of the monetary neutrality suggests that levels of prices and the money supply will be proportional. Consistent with this long-run hypothesis, preliminary observation shows that the money supply, agricultural prices and manufacturing prices in New Zealand have not moved apart for long periods of time (see figure 1). Thus, the answers to our first three questions, which essentially concern short-run dynamics, may be linked to the answer to the fourth question about the long run.

Recently, the use of vector autoregressive (VAR) models to examine the



*Seasonally adjusted

Figure 1. The money supply (M1), manufacturing prices, and agricultural prices in New Zealand, 1963:1-1987:1

dynamic interactions among money and prices has received emphasis (e.g.; Bessler, Chambers, Orden, Taylor, and Devadoss and Meyers). A VAR model describes the probabilistic nature of time series data within a forecasting system of autoregressive equations in which all contemporaneous variables are affected by all own and cross-variable lags (Sims, 1980). This approach recognizes the dynamic interdependence that exists among economic phenomena, and, hence, the inappropriateness of specifying particular variables as strictly exogenous when their evolution is dependent on previous values of endogenous variables. Using Sims' innovation accounting framework, a primary advantage of these models has been to allow examination of the responses of all the variables in the model to deviations from their expected time paths, without the requirement of maintaining strict ceterus paribus assumptions on the evolution of the other variables. However, a VAR model does not impose any testable, theoretically consistent, long-run equilibrium conditions. When such equilibrium conditions exist this can result in misspecification and, as we will show, can have important consequences on model performance.

It is generally recognized that the statistical validity of VAR models in estimation, inference, and forecasting procedures may be altered dramatically depending on the time series properties of the observed data (Granger and Newbold). A VAR model using nonstationary data in levels (unit roots in the generating process) can produce biased forecast errors if the data are not appropriately differenced to obtain the required stationarity.[2] Subsequent decomposition of the errors from a model

[2] A series is said to be stationary (in a weak sense) if it has a finite mean and variance that do not vary with time and covariances that depend only on the time interval between observations, not on time itself.

estimated with nonstationary data can be misleading because the covariances between series will not be independent of time. Also, the parameter estimates themselves will not be asymptotically normally distributed (Phillips and Durlauf), but will have distributions dependent on the number of variables in the model. Nevertheless, most previous VAR models of monetary-agricultural interdependencies have been estimated in levels of potentially nonstationary series, without formal tests of the levels specification.

Recent advances in understanding time series data which are nonstationary because they contain unit roots have emphasized a somewhat different approach to the stationarity issue. The basic concept is that economic variables linked by long-run equilibrium relationships should not drift too far apart over time, as individually nonstationary series do. As a consequence, equilibrium relationships among variables may be manifest in stationarity of linear combinations of variables that are themselves nonstationary (cointegration). In these cases, explicit modeling of deviations from the implied equilibria is important (Engle and Granger, Phillips). A specification which incorporates cointegration is a vector error-correction (VEC) model: a VAR in differences with additional terms measuring lagged deviations from the long-run levels relationships of the variables. A recent application by Engle and Yoo found substantially improved forecast performance for a VEC model compared to a VAR. King, Plosser, Stock and Watson find the VEC specification useful in a study of permanent versus transitory growth in income and consumption under technological innovations. These studies have illustrated the need for careful preliminary analysis of the data's time series properties before a particular model is chosen for forecasting or policy analysis.

Our paper emphasizes this concern in evaluating dynamic relationships among money and prices in New Zealand. After describing the implications of cointegration in terms of appropriate treatment of nonstationarity and model specification, a strategy to determine a dynamic model congruent with quarterly money and price data is implemented. First, we examine the time series properties of the chosen data in a set of univariate autoregressive forecasting equations. We test the null hypothesis of nonstationarity (in this case, a unit root) against the alternative that each series is stationary around a linear time trend. Based on our results, we then examine whether the series are cointegrated. In conjunction, we test for misspecification of the univariate models in terms of the appropriate order of the lag structure, predictive stability (parameter time dependence), normality of the estimated residuals, and heteroscedasticity (of both a unconditional and autoregressive conditional type).

We find evidence of nonstationarity of the individual money and price series, of stationarity of each series in first differences, and of predictive stability of differenced univariate models over the 1964:2-1987:1 sample period. We also find strong evidence of cointegration among money, agricultural prices, and manufacturing prices consistent with the hypothesis of long-run monetary neutrality. We estimate the implied VEC model (again checking for the statistical validity of this specification), then use the estimated model to compute innovation statistics and evaluate money and price dynamics, assuming a recursive contemporaneous orthogonal ordering of money, manufacturing prices, then agricultural prices. Using

Bayesian integration procedures (Doan and Litterman, Devadoss and Meyers) we find statistically significant evidence that agricultural prices respond more quickly than manufacturing prices to a monetary shock but do not overshoot their new long-run equilibrium values. We also find significant evidence of a short-run price disadvantage in agriculture following a manufacturing price shock, but not of short-run or long-run feedback from agricultural prices to manufacturing prices or the money supply. These inferences are quite unlike those drawn from a VAR model in differences which would have been the specification chosen on the basis of nonstationarity of the individual series without considering cointegration.

Cointegration of Nonstationary Time Series

Following Granger, consider three series, x_t, y_t and w_t, each of which is nonstationary (integrated of order one, I(1)). Assume each series has no deterministic trend or, without loss of generality, a drift component (that is, a pure random walk). A common technique in this situation has been to difference each series to obtain stationarity.

Engle and Granger suggest, as an alternative, that although the series may be individually nonstationary, there may exist constants (cointegrating parameters), T and δ, such that the error-correction terms:

$$z_{1t} = w_t - Tx_t \quad (1a)$$

$$z_{2t} = w_t - \delta y_t \quad (1b)$$

are stationary, I(0). In this case, x_t and w_t, and y_t and w_t (and, therefore, by transitivity, x_t and y_t), share a common stochastic trend and are said to be cointegrated. The right hand sides of equations (1a,b) can be considered the long-run equilibrium relationships between the series. The z_{it}'s represent the deviations from these equilibria at a particular moment in time. However, if z_{1t} and z_{2t} are I(0), then stationarity among the three series can be expressed by any number of reparameterizations of equations (1a,b). The "equilibrium relationships" exhibited in the tri-variate model are, therefore, not unique, unlike in a bi-variate model. Nonetheless, the two specific equations (1a,b) are useful. They span the space of cointegrating relationships when nonstationarity among the three variables occurs because they share a single common stochastic trend. When x_t is the money supply, and y_t and w_t are manufacturing prices and agricultural prices, equations (1a,b) also have an intuitive interpretation: $T = \delta = 1$, implies long-run proportionality between the levels of the money supply and prices (i.e.; between the levels of agricultural prices and of the money supply, between the levels of agricultural prices and of manufacturing prices, thus, by transitivity, between the levels of manufacturing prices and of the money supply). These relationships are consistent with the long-run neutrality of money when technical change or other factors have not shifted relative prices over time.

Engle and Granger prove that when there is cointegration as in (1a,b) there exists a vector autoregressive representation:

$$\begin{bmatrix} x_t \\ y_t \\ w_t \end{bmatrix} = \sum_{s=1}^{p} \begin{bmatrix} a_{11}(s) & a_{12}(s) & a_{13}(s) \\ a_{21}(s) & a_{2}2(s) & a_{23}(s) \\ a_{31}(s) & a_{32}(s) & a_{33}(s) \end{bmatrix} \begin{bmatrix} x_{t-s} \\ y_{t-s} \\ w_{t-s} \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \end{bmatrix}, \qquad (2)$$

that is restricted such that $\left[I - \left[\{\sum_{s=1}^{p} a_{ij}(s)\} \right] \right]$ is of rank one. There also exists a vector error-correction representation:

$$\begin{bmatrix}\Delta x_t \\ \Delta y_t \\ \Delta w_t\end{bmatrix} = \begin{bmatrix}-b_{11} & -b_{12} \\ -b_{21} & -b_{22} \\ -b_{31} & -b_{32}\end{bmatrix} \begin{bmatrix}z_{1t-1} \\ z_{2t-1}\end{bmatrix} + \sum_{s=1}^{p^*} \begin{bmatrix}c_{11}(s) & c_{12}(s) & c_{13}(s) \\ c_{21}(s) & c_{22}(s) & c_{23}(s) \\ c_{31}(s) & c_{32}(s) & c_{33}(s)\end{bmatrix} \begin{bmatrix}\Delta x_{t-s} \\ \Delta y_{t-s} \\ \Delta w_{t-s}\end{bmatrix} + \begin{bmatrix}\varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t}\end{bmatrix} \quad (3)$$

where $|b_{i1}| + |b_{i2}| \neq 0$, for at least one i.[3]

Equations (3) show that deviations from the long-run equilibria expressed by the cointegrating relationships affect the dynamic interactions within a system of differenced autoregressive equations through the error-correction terms. The vector error-correction model imposes long-run constraints among the cointegrated variables, without explicitly restricting short-run dynamics.[4] As proved by Engle and Yoo, forecasts of cointegrated nonstationary economic variables from a VAR in levels are asymptotically consistent, but will be inefficient because the long-run constraints are omitted. Forecasts of cointegrated nonstationary variables from a VAR model in differences (a nested special case of the vector error-correction representation when there are no common stochastic trends) will, in principle, be inferior because the model is misspecified. In the absence of cointegration, however, nonstationary data containing a unit root should be

[3] Since x_t, y_t, and w_t are I(1), their differences will be I(0), and every variable in (3) will be I(0) if the z_{it}'s are I(0). If, for example, z_{1t} is not I(0), then the series x_t and w_t are not cointegrated. This does not rule out the possibility that z_{2t} may be still be I(0), in which case there exists one cointegrating relationship for y_t and w_t, or there may be one cointegrating relationship among x_t, y_t, and w_t. In general, there may exist m unit roots in a system of m equations each of which is I(1). There may be k stationary error-correction terms, where k is the number of unique cointegrating relationships. This implies m-k stochastic trends shared by the series (in the case presented above, m=3 and k can equal 0, 1, or 2).

[4] Another form of error-correction model, suggested by Hendry, focuses on restrictions on short-run dynamics while making them theoretically consistent with a long-run solution (see Wickens and Breusch for a discussion of the distinction). These models are viable alternatives until tested against the Engle and Granger specification, but are not considered here.

modeled in differences. Alternatively, if levels data are nonstationary but are stationary around a deterministic trend instead of a stochastic trend (a unit root) the series should be modeled in levels with a time trend included. Stationary data series can be modeled in levels, while differencing should not affect the results if the model is correct.

An Empirical Model of New Zealand Money and Prices

Descriptive Statistics

The data are quarterly, initially seasonally unadjusted series on the money supply, M1, agricultural prices, FP, and manufacturing prices, IP, from 1963:1 through 1987:1. The data are from the New Zealand Reserve Bank and Department of Statistics data bases and are transformed by taking logarithms. The measure of money is the level of M1 (currency plus demand deposits), and the prices are producer output price indices excluding taxes and subsidy payments.

The existence of seasonality in the series was investigated first by examination of the autocorrelation functions and the joint significance of seasonal dummy variables in univariate forecasting equations for each variable. The results suggested that only the money supply variable exhibited strong seasonality, which was adjusted for via the Holt-Winter smoothing technique as described in Doan and Litterman. The price series are unadjusted.

A number of characteristics of the data (plotted in figure 1) are described in table 1, which presents tests of whether the individual series are nonstationary, of whether they are cointegrated, and of misspecification of the univariate models. Tests of the hypothesis that each series contains a unit root against the alternative that it is stationary, possibly around a linear time trend, are reported in the first column of panel A. These tests are based on an OLS estimated univariate forecasting equation which provides an autoregressive approximation to short-run correlations in the series. The general form of this equation is:

$$\Delta X_t = \beta_0 + \beta_1 t + \beta_2 X_{t-1} + \sum_{i=1}^{3} \beta_{2+i} \Delta X_{t-i} + E_t \tag{4}$$

where a third-order lag structure for each equation was selected for all series on the basis of a set of F-test comparisons. The Dickey-Fuller(1979) test is used and the results suggest that the unit root hypothesis (β_2=0) can not be rejected at the 0.10 significance level for M1 and IP and at the 0.05 level for FP. The second and third columns of panel A, respectively, present t-tests on the estimated coefficients on the time trend (β_1) under the maintained hypothesis of a unit root (β_2=0), and on the drift component (β_0) under the maintained hypothesis of a unit root and no deterministic time trend (β_1=β_2=0). Only the M1 series exhibits significant trend behavior at the 0.10 level, but a zero drift component is rejected for all three series.

The last two columns of panel A present the results of joint parameter restriction tests as described in Dickey and Fuller(1981). The alternative in each case is equation (4), while the null hypotheses are, respectively, that the series are random walks with no drift and no deterministic time trend (β_0=β_1=β_2=0), or are random walks with a drift but no time trend

Table 1. Unit root and cointegration characteristics of the data

Panel A: Unit Root Statistics

Series	D-F	trend	drift	R-Walk	R-W/Drift
ΔM1	-2.666	1.890*	2.102**	5.265**	5.465
ΔIP	-2.786	0.765	2.040**	5.083**	4.192
ΔFP	-3.280*	0.560	1.725*	4.798*	5.355

Panel B: Error Correction Terms

FP	=	1.04M1	+ 0.0	- 3.62	+ Z_1	(a)
FP	=	0.0	+ 0.86IP	+ 0.55	+ Z_2	(b)

Panel C: Cointegration Statistics

Residual	ADFC	ADFC(restricted model)
Z_1	-4.574***	-3.646**
Z_2	-3.405**	-2.906*

Panel D: Summary and Misspecification Statistics

Series	$\bar{R}^2$	σ%	HS	CHOW	ARCH	KS
ΔM1	0.69	1.06	2.02*	2.08*	3.13	1.67
ΔIP	0.56	1.09	1.21	0.52	5.08	0.67
ΔFP	0.25	4.19	4.28***	1.10	8.19*	3.00

T=92 (1964:2-1987:1)

* represents rejection of the null hypothesis at the 0.10 level; ** at the 0.05 level; and *** at the 0.01 level of significance, based on the particular distributions appropriate for each test (the Dickey-Fuller(1979), t distribution, Dickey-Fuller(1981), augmented Dickey-Fuller(Engle and Yoo), and F distribution ((T2,T1), (k,T-k)), $Chi^2(4)$, and $Chi^2(2)$, respectively).

(β_1=β_2=0). This evidence generally supports the existence of three unit roots in the series (one per series) with a drift but no deterministic time trend.

Next, Stock has shown that OLS theoretically provides a consistent estimate of the cointegrating parameters in a static regression of current-dated variables in levels (with a constant to adjust for differences in scale) and can form the basis for tests of cointegration. Panel B presents the estimates of these cointegrating regressions. The estimated parameters indicate that if the series are cointegrated the long-run relationships between levels of M1 and FP, and between levels of FP and IP, are approximately proportional.

Column one of panel C displays the results of a test of the hypothesis that the money supply, agricultural prices, and manufacturing prices are cointegrated. The test is obtained by observing the extent of autocorrelation (or equivalently, a unit root) in the residuals from the cointegrating regressions in panel B. Statistics that have been suggested to test the cointegration hypothesis include the Durbin-Watson statistic, Dickey-Fuller cointegration tests (DFC) (the Dickey-Fuller test for unit roots adjusted for the dependence of the z_{it}'s on the estimated cointegrating parameters), and the Stock and Watson test. Here, the results of the augmented DFC test (adjusting to allow for the possibility of higher order dynamics in the error process) are presented (Engle and Yoo). For both equations, the results reject the presence of unit roots in the residuals (at the 0.05 level of significance for the agricultural price-money equation and at the 0.01 level for the agricultural-manufacturing price equation) suggesting that there are two cointegrating relationships among the three series.[5]

Because the initial parameter estimates are close to unitary, this restriction was imposed on the cointegrating regressions and new residuals (error-correction terms) were estimated and tested. These error-correction terms impose strict long-run proportionality among money and prices and makes interpretation of impulse responses from a VEC model consistent with long-run monetary neutrality. The augmented DFC test results reported in column two of panel C suggest the series are cointegrated when this restriction is enforced.[6]

Finally, a number of summary and misspecification statistics for the univariate models are reported in panel D of table 1. The explanatory power of the models is not high. In particular, the FP autoregression has a low adjusted R^2 and the estimate regression standard error is higher than for any of the other regressions. The HS test (Spanos(pg. 484)), distributed F(47,25)), is a simple test of whether the estimated residual variances are homoscedastic when two sub samples (1964:2-1971:4 and 1974:1-1987:1) are

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[5] The space of possible bi-variate cointegrating relationships spanned by the two reported estimates also contains that between M1 and IP. A regression of IP on M1 yields an estimated cointegrating parameter of 1.19, and cointegration is not rejected at the 10 percent level of significance of the ADF test. This parameter value is also obtained by transitivity from the two reported estimates (1.04/0.861). Reversing the order of the dependent and independent variables produced approximate inverses of the parameter estimates in all three regressions (up to $1-R^2$, Engle and Granger p.261). Also, Banerjee, Dolado, Hendry, and Smith find that, in practice, OLS estimates from static regressions may be biased in small samples because the dynamic relationships are not specified, and that tests for cointegration based on a static regression often lead to an erroneous over rejection. They suggest that a rule of thumb as to the usefulness of the static regression estimates is the size of the R^2. In our case, for all three equations (and their inverses), the R^2 is greater than 0.90.

[6] Formal tests of the equivalence of these parameters would involve nonstandard distribution theory (see Sims, Stock and Watson).

compared. The results suggest that the FP regression residuals exhibit heteroscedasticity at a 0.01 level of significance. Despite this result, CHOW tests (distributed $F(5,81)$) for predictive stability over the period 1986:1-1987:1 indicate that the parameter estimates for each equation may not be seriously time dependent. Any changes over the sample period do not have a statistically significant impact on the predictive performance of the equations. The ARCH statistic (distributed $Chi^2(4)$) provides an indication of whether the estimated variances behave autoregressively. Squared values of the estimated residuals were regressed against 4 lags of the squared residuals and a constant. The ARCH effect does not appear pronounced, relative to the null that the variances are constant, though there is weak evidence of ARCH effects in the FP equation. Last, the results from a series of skewness-kurtosis tests (Bera and Jarque) are consistent with the estimated residuals being normally distributed. These tests are distributed $Chi^2(2)$, and are denoted KS.

The overall implications of our preliminary statistical tests are that M1, IP, and FP contain unit roots (random walks with drift) and that univariate autoregressive models using differenced data do not seriously violate the classical assumptions concerning the statistical properties of the estimated models. Moreover, while the individual series are nonstationary, they are cointegrated and the ratios FP/M1 and FP/IP (and, hence IP/M1) are stationary. As observed in figure 1, the series may move apart for periods of time but there is a long-run tendency for them to come together with proportionality between levels of the money supply and prices.

The Vector Error-Correction Model

A VEC model (in the form of (3)) was estimated using first differences of each series, a constant term, and the error-correction terms from the restricted cointegrating regressions. The model was estimated with OLS. The lag structure for the system was determined by a series of F tests and Akaike's Information Criterion. In both cases, a third-order model was suggested.

The basic estimation results, and residual checking statistics, are presented in table 2. The adjusted R^2's from the estimated forecasting equations forming the VEC model are higher than from the univariate models (except for the IP equation) or from an unreported VAR in differences without the error-correction terms, particularly for the FP equation. The HS statistic (distributed $F(47,25)$) suggests heteroscedasticity in the FP equation, but the CHOW test (distributed $F(5,76)$) for predictive stability is again consistent with stable long-term parameter estimates. There is no evidence of ARCH effects, while the Q statistic suggests no higher order autocorrelation in the residuals from the estimated equations. The KS statistic indicates normally distributed residuals. These results imply that the VEC model is not severely misspecified, and provides a robust statistical model. Also, consistent with the cointegration tests, although a VAR in differences is also a statistically valid model, an F test of the restriction that the error-correction terms in the VEC model are jointly zero ($F(6,81) = 5.37$) is rejected at the 0.05 level of significance in favor of the maintained VEC representation.

Table 2. VEC model summary statistics

Statistic	$\Delta M1$	ΔIP	ΔFP
$\bar{R}^2$	0.71	0.53	0.40
$\sigma\%$	1.03	1.13	3.73
HS	2.19*	1.19	5.28***
CHOW	1.99*	0.99	1.64
ARCH	8.52*	3.07	4.34
KS	1.03	0.44	1.78
Q(27)	36.09	34.35	32.92

T=92 (1964:2-1987:1)

Notes as for Table 1 and as discussed in the text.

Inference Results

For the VEC model orthogonalized in the order M1, IP, FP, the impulse responses to a unit innovation in each variable's expected time path and their estimated 90-percent confidence bounds are plotted in Figure 2.[7] To trace the dynamic effects of these innovations through the system it is necessary first to expand the estimated VEC into a levels form in which the parameters of the error-correction terms are incorporated into the parameters of the variables being impulsed. Otherwise, impulsing procedures will treat the error-correction terms simply as unrelated deterministic constants.

The point estimates from the impulse response functions show that each innovation will eventually lead to long-run percentage responses that are approximately the same for all three of the variables, as expected from the imposition of the theoretical error-correction model. The estimated standard errors of the impulse responses diverge rapidly, so interpretation of the point estimates must be qualified.[8] Nonetheless, when asking about the significance of various responses, we are interested in knowing whether the posterior probability is concentrated on zero, negative, or positive values (Sims, 1987). The estimated 90-percent confidence bounds suggest that a number of the responses are positive, while some are concentrated

[7] Each point estimate in the impulse response function was calculated as the mean from a Monte carlo integration algorithm using 200 draws from the posterior distribution of each response estimate (Doan and Litterman, chpt.17). The confidence bounds for each estimate were also calculated from this procedure.

[8] Runkle has shown that the asymptotic t-statistic for the nth term in the impulse response function goes to zero at the rate a/n, where a is the point estimate of the nth response. The effect of a current shock e_t gets multiplied by a^n, so its importance can diminish rapidly.

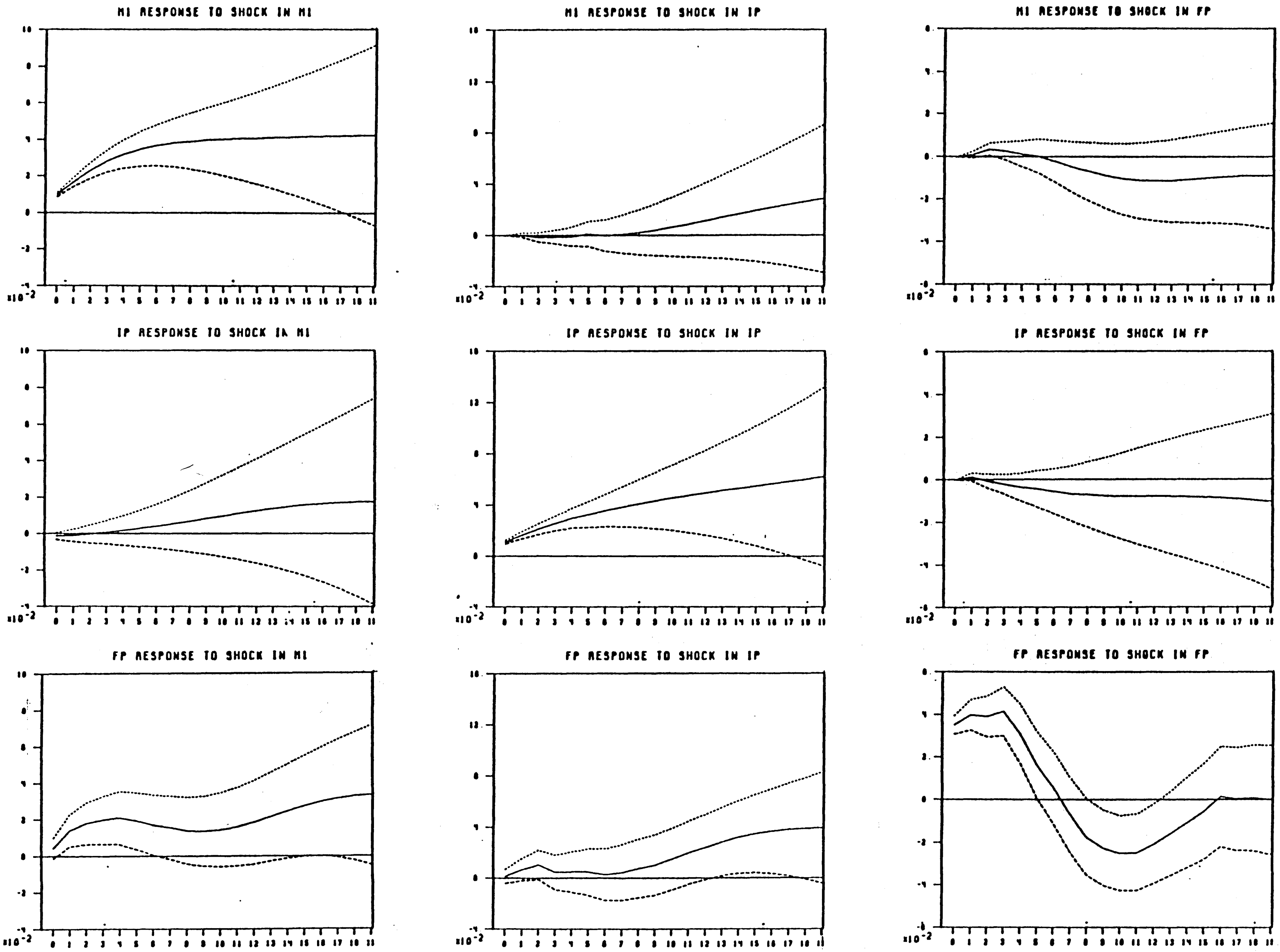

Figure 2. Impulse response functions from a VEC model of money (M1), manufacturing prices (IP), and agricultural prices (FP) in New Zealand

around zero. Thus, we can make several general observations about the likely time paths for the variables under specific shocks. In particular:

1) A positive monetary shock raises the levels of the money supply and agricultural and manufacturing prices in the long run. Agricultural prices are immediately responsive to the monetary shock and display some indication of a cyclical response pattern, while manufacturing prices start to respond only after a lag of 4 quarters. Thus, positive monetary shocks induce a shift in relative prices in favor of agriculture in the short run. Agricultural prices do not overshoot their new long-run equilibrium level in response to a monetary shock, but they move toward the new equilibrium faster than manufacturing prices.

2) A positive shock to manufacturing prices also raises the long-run levels of all three series. A relatively slow response of agricultural prices initially places agriculture at a price disadvantage. Moreover, agricultural prices start to rise only when the money supply begins to respond to the shock to manufacturing prices.

3) The own response to a shock to agricultural prices displays a dampening oscillatory pattern. Neither the money supply nor manufacturing prices shows a significant short-run response. Further, in contrast to the permanent shift in the level of each series induced by shocks to the money supply or manufacturing prices, new equilibrium of the series are established at around the same levels as before the agricultural price shock occurred. Thus, there is neither a short-run nor long-run response of the

money supply or manufacturing prices to an agricultural price shock.

Inferences from the impulse responses of the VEC model are supported by the associated forecast error-variance decompositions. These (unreported) decompositions suggest that both the money supply and manufacturing prices behave exogenously, and, in turn, account for a large proportion of the forecast-error variance of agricultural prices. There is no evidence of feedback from agricultural prices to money.

For comparative purposes, it is also interesting to examine the impulse responses from alternative (unreported) VAR specifications. The point estimate impulse responses from a VAR in levels are generally similar to those from the VEC model except that long-run convergence is less evident and the estimated confidence bounds are more imprecise. In contrast, inferences quite unlike those from the VEC model are derived from a VAR model in differences. The differences model suggests more-than-proportionate short-run responses of both manufacturing and agricultural prices to money supply shocks, and that agricultural price shocks induce responses in the level of the money supply, implying a feedback relationship. In the long run, agricultural price shocks persist in a VAR in differences, while manufacturing price shocks do not. The long-run positions of the point estimates diverge in response to specific shocks, contrary to the data evidence which suggests that they should converge. These empirical results conform to Engle and Yoo's observation that a VAR in levels is asymptotically equivalent to the VEC model (but inefficient) and a VAR in differences is misspecified when the data are cointegrated.

Conclusions

One objective of applied research is to produce explanations (or at least predictions and descriptions) for actual economic phenomena. The model utilized should be an adequate approximation of the data generating process giving rise to the phenomena of interest in terms of statistical validity, explanatory performance, and theoretical consistency. In this paper, testing for the presence of nonstationarity was an important aspect of obtaining a statistically valid model (in conjunction with diagnostic checks). Testing for cointegration had important implications for model specification and performance. Cointegration of individually nonstationary money and price time series was found to be consistent with long-run monetary neutrality.

When the long-run equilibrium relationships in the data are accounted for in a VEC model which does not explicitly restrict short-run dynamics, with a contemporaneous causal ordering of money, manufacturing, then agricultural prices, an interesting data-congruent explanation of money and price behavior in New Zealand emerges. In terms of the questions to be addressed about short-run money and price dynamics, first, it seems that money does matter to prices and that in the short run relative prices shift to favor agriculture in response to monetary shocks, without agricultural prices overshooting their new long-run equilibrium values. Second, autonomous manufacturing price shocks induce long-run responses in the money supply and agricultural prices with lags resulting in a short-run price disadvantage in agriculture. Third, agricultural prices appear sensitive to own shocks, but this response quickly dissipates and the money supply and manufacturing prices are largely unaffected.

An important interpretation of our results is that monetary policy (through adjustments in the growth of M1) may be affecting the two commodity sectors disproportionately over the short to medium term, although, in the long run market and other forces drive prices together. Moreover, if the manufacturing sector is able to administer output prices for periods of time (for example, through import protection or as a markup over normal costs), then current policies to expose the agricultural sector further to world trade competition in isolation from manufacturing will require agriculture to bare a greater share of the burdens of economic adjustment. Finally, the lack of feedback from agricultural prices to the money supply stands in contrast to the long-run responsiveness of the money supply to shocks to manufacturing prices. Since historically the agricultural sector has behaved as a price taker on world markets under a regime of historically gradually adjusting nominal exchange rates, the lack of feedback from agricultural prices to the money supply suggests unwillingness of the monetary authorities to accommodate world agricultural price shocks. This behavior may isolate the New Zealand economy from external inflationary and deflationary pressure arising from agriculture, while failing to respond to effects that price shocks may have on New Zealand's current account.

A comparison of our results for New Zealand to other studies of money and price dynamics highlights the importance of the model specification issue. Previous studies have been based on VAR models estimated in levels (an exception is Taylor who differences some series but not others). Little attention has been given to diagnostic checking of the statistical validity of the levels specification, such as unit root tests, or to modeling long-run dynamic constraints. Long-run impulse responses have generally not been shown (for example, Chambers presents responses over only 12 months, while Bessler, and Devadoss and Meyers provide estimates of impulse responses over twenty-four monthly periods, and Orden over eight quarters).

One is uncertain whether short-run inferences from these levels VAR models are statistically valid or are consistent with long-run equilibrium relationships among the included variables. For Brazil, Bessler fails to reject monetary neutrality (under usual distributional assumptions) based on a statistical comparison between the summed coefficients on money in the autoregressive equations for agricultural and industrial prices. Even so, his point estimate of the response of industrial prices to an initial monetary shock is approximately twice the estimated effect of the shock on agricultural prices or the money supply itself after 24 months. For the United States, Devadoss and Meyers note an opposite result. They show a statistically significant (again under usual distributional assumptions) short-run shift in relative prices in favor of agriculture in response to a monetary shock. After 24 months, their point estimate of the percentage increase of agricultural prices is 1.3 times greater than the percentage increase of industrial prices and nearly five times greater than the impact on the money supply. Conflicting results such as these may be due to behavioral differences between economies, as has been conjectured, but could be due to model misspecification. Likewise, Orden reports little effect of a shock to the money supply on real U.S. crop prices in a model that includes an interest rate, an exchange rate, the GNP deflator, and agricultural exports; with a greater response of real crop prices to interest-rate and exchange-rate shocks. Again, model misspecification is an issue in interpreting these results. Although VAR models in levels are asymptotically valid in the presence of cointegration, in the absence of cointegration tests each of these models may or may not have been misspecified.

Finally, though the specification of the three-variable VEC model reported in this paper has been relatively carefully tested, the model is easily tractible, and the characteristics of the New Zealand economy described by the model can be given plausible economic interpretations, some caveats should be noted. As Hendry and others have emphasized, the acceptability of a model based on empirical evidence should be considered only tentative until it has been successfully tested against new data, different criteria, and rival models. Also, the model is a closed system that omits the potential influence of other variables, and may be limited by the contemporaneous structure imposed (Bernanke, Fackler). A recursive causal ordering of money, manufacturing prices, then agricultural prices is useful in economic environments where orthogonal shocks can be identified with distinct sources, but may not be a realistic approximation to the actual contemporaneous relationships among money and prices (for instance, consider the apparent short-run exogenous nature of manufacturing price behavior). Assuming that the data itself is rich enough to reveal the underlying phenomena of interest, the development of larger models with more appealing contemporaneous structures is a priority for future research. Together with the improvements in techniques for econometric modeling of time series data, the future for these types of dynamic models is encouraging.

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