A Vector Error-Correction Model of Money

and Price Dynamics in New Zealand

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Introduction

Whether monetary policy has important impacts on real activity has been an enduring source of discussion in Keynesian-monetarist-classical debates on money, prices, and business cycles in modern industrial economies. Since the publication of Sims' influential article on Granger causality relationships between money and income in 1972, the use of vector autoregression (VAR) time series models, as, for example, in Hsiao, Sims (1980a,b, 1986), Thorton and Batten, Litterman and Weiss, or Bernanke, has gained prominence among applied macroeconomists as an empirical tool for evaluating monetary impacts. These models focus on the major features of the economy, such as the money supply, the price level, real output, interest rates, and a few other variables, with only minimal restrictions on their dynamic interactions.

Unfortunately, seemingly small changes in the specification of autoregressive models have yielded conflicting answers to the money-prices-real activity question. Stock and Watson (1987b), Christiano and Ljungqvist, and Ohanian, among others, have suggested recently that this dilemma may be due in part to failure to adequately evaluate the time series characteristics of the data or take into account the implications for model specification and hypothesis testing of nonstationarity due to unit roots common among aggregate macroeconomic variables (Stulz and Wasserfallen, Nelson and Plosser, Hall, Stock and Watson (1986)).

Of particular interest in this regard is that advances in understanding time series data which are nonstationary because they contain unit roots have recently brought to light a stronger relationship between postulates of economic theory and time series properties of the data than had previously been recognized. The underlying concept is that economic variables linked by long-run equilibrium relationships should not drift too far apart over time. Since nonstationary series with independent unit roots have no tendency to move together in the long-run, independent unit roots are inconsistent with existence of long-run equilibrium. Consequently, if there are long-run equilibrium relationships among the variables, these relationships may be manifest in stationarity of linear combinations of the variables, in which case a multivariate model will contain fewer unit roots than the individual series. Such a relationship is termed cointegration (Granger 1981). Existence of a cointegrating relationship implies testable restrictions on the parameters of a VAR model of the variables in levels. Engle and Granger have shown that a convenient way to express these restrictions is a vector error-correction (VEC) model. In a VEC model, lagged deviations from any linear relationships measuring stationary long-run equilibrium among the variables affect the dynamic interactions within a system of autoregressive equations of first differences of the series. Failure to impose the restrictions implied by cointegration, or, alternatively, failure to recognize cointegration among the variables in transforming individual series to obtain stationarity (such as by first differencing) can have serious consequences for interpretation of VAR models.

The concerns about the time series properties of the data and its implications for model specification and inference also apply to the empirical analysis which has focused on the effects of monetary policy on the agricultural sector. Much of the theoretical development in this area has been based on the view that the impacts of monetary policy can be evaluated in terms of its impact on relative prices. For example, Bordo postulated that the existence of labor and sales contracts of varying lengths between the agricultural and non-agricultural sectors of the economy will cause disproportionate price responses to monetary shocks. Frankel, and Rausser, Chalfant and Stamoulis argue similarly that agriculture can be viewed as a competitive sector in which prices are flexible compared to relatively less flexible prices among non-agricultural sectors. Expansionary monetary policy then favors agriculture by causing agricultural prices to overshoot their long-run equilibrium values. Alternatively, Tweeten, suggests that because the non-agricultural sector is typically oligopolistic in nature with prices being set as a mark up over variable costs, non-agriculturally generated price shocks will place agriculture at a price disadvantage by creating a cost-price squeeze. Finally, while the overshooting and cost-price squeeze hypotheses focus on the short run, Batten and Belongia, Chambers and others have emphasized the hypothesis of monetary neutrality which suggests that in the long run monetary shocks will have equal nominal impacts and no effect on relative prices.

A number of empirical studies investigating these theoretical propositions about money, agricultural prices, and industrial prices have been conducted recently using VAR models. These models typically include a feedback equation to describe the policy-setting behavior of the monetary authorities and reduced-form equations to capture the likely responses of agricultural and non-agricultural prices to shocks of monetary and sectoral origins. In a study for Brazil, for example, Bessler estimated that after 24 months the response of industrial prices to a monetary shock is approximately twice the estimated effect of the shock on agricultural prices or the money supply itself. Devadoss and Meyers report an opposite result in a model of the money supply and agricultural and industrial prices for the United States. They show a statistically significant (under asymptotic distributional assumptions dependent on stationarity of the model) short-run shift in relative prices in favor of agriculture in response to a monetary shock. After 24 months, the estimated percentage increase of agricultural prices is about one-and-one-half times greater than the percentage increase of industrial prices and nearly *five* times greater than the increase in the money supply. Orden, in contrast, reports only a slight positive effect of a shock to the money supply on real U.S. crop prices over eight quarters in a model that includes an interest rate, an exchange rate, the GNP deflator, and agricultural exports, with a a greater response of real crop prices to interest-rate and exchange-rate shocks. Taylor and Spriggs also find the exchange rate of the U.S. dollar against other major currencies to be an important source of instability for Canadian agricultural prices and present evidence that shocks to the Canadian money supply cause agricultural prices to rise relative to industrial prices in the short run.

Differences such as these among measurements of monetary impacts on nominal and relative prices may be due to behavioral differences among economies, as has been conjectured by Bessler and Devadoss and Meyers to explain the contrast in their results. These differences could also be due to the sort of nonstationarity misspecification issues raised in the macroeconomic literature. With the exception of Taylor and Spriggs, who difference the money and price series, each of the VAR models of monetary impacts on agriculture was estimated using levels of potentially nonstationary data without investigating this possibility or its consequences.

This paper provides additional evidence about monetary impacts on agriculture with particular attention to statistical tests for nonstationarity and cointegration among aggregate money and price series and model specification based on the results of these tests. Initially, we describe the sensitivity of empirical results of VAR models to incorrect maintained hypotheses regarding stationarity. Following Engle and Granger, we also show how cointegration and the VEC model specification provide a means to incorporate theoretical equilibrium postulates into dynamic models of nonstationary variables in a plausible and testable manner. We then draw on these results to investigate the hypotheses discussed above (overshooting, the cost-price squeeze, and long-run monetary neutrality) by estimating parameters of a three-variable quarterly model of the money supply, agricultural prices, and manufacturing prices in New Zealand. For this small, open and agriculturally-dependent economy we also investigate feedback from agricultural prices to the money supply and industrial prices.

Generally, our results demonstrate the importance of considering the time series properties of the data as part of an evaluation of monetary impacts on agriculture. We find empirical support for the hypothesis of long-run monetary neutrality and provide estimates of short-run dynamics among money and prices that incorporate this long-run restriction. Specifically, we reject the hypothesis that any of the quarterly data series are individually stationary in levels in favor of the hypothesis that each series contains a unit root and is stationary in first differences. Under the latter maintained hypothesis, we find evidence of two independent stationary linear combinations of the series that reduce to one the number of unit roots in the three-variable model. A crucial result is that our estimates of the parameters of these cointegrating relationships are consistent with long-run proportionality of the level of the money supply and the level of each nominal price. A VEC model imposing this restriction is then estimated, tested for misspecification, and used to evaluate short-run and long-run money and price dynamics. We find statistically significant evidence that agricultural prices respond more quickly than manufacturing prices to a monetary shock in the short run, but we do not find that the short-run agricultural price response is proportionately greater than that of the money supply or that short-run agricultural prices overshoot their equilibrium value in the long run. We also find significant evidence of a short-run cost-price squeeze in agriculture following a manufacturing price shock, but there is no evidence of short-run or long-run feedback from agricultural prices to manufacturing prices or the money supply. These results from the VEC model are consistent with some of the institutional characteristics of the New Zealand economy during our study period. They also provide a somewhat different description of money and price dynamics than has been reported for other countries on the basis of VAR models. For New Zealand, comparison of the results from the VEC model to estimates of the dynamic interactions among money and prices from VAR models specified under alternative maintained hypotheses about stationarity illustrates that in the absence of preliminary testing for the time series properties of the data, as well as for overall statistical validity of the chosen autogregressive model specification, there is reason to be cautious in evaluating the magnitude of the parameter estimates from a VAR model and drawing inferences on the basis of the estimates that are provided.

Stationarity, Cointegration, and Specification of Autoregressive Models

Wold's multivariate decomposition theorem states that any mean zero stationary vector time series Y_t can be represented as a weighted sum of present and past uncorrelated shocks, such as:

$$Y_t = C(L)\,\xi_t \tag{1}$$

where Y_t and ξ_t are N x 1 vectors, C(L) is an infinite polynominal of N x N matrices in the lag operator ($C(L) = C_0 + C_1L + C_2L^2 + \ldots$), and ξ_t is a white noise process with $E[\xi_t] = 0$, $E[\xi_t \xi'_t] = I_N$ and $E[\xi_t\xi'_{t-s}] = 0$, $s \neq 0$. In this representation, contemporaneous interactions among the variables in Y_t determine the coefficient of C_0.

If C(L) is invertible and its inverse can be approximated by a finite-order matrix polynominal in the lag operator, then Y_t can be expressed in the multivariate autoregressive form:

$$D(L)Y_t = \xi_t \tag{2}$$

where $D(L) = D_0 + D_1L + \ldots D_P\, L^P$ and contemporaneous interactions are measured by the coefficients of D_0. A VAR model convenient for estimation is obtained by pre multiplying (2) by D_0^{-1} to obtain:

$$A(L)Y_t = U_t \tag{3}$$

where A(L) = D_0^{-1}D(L), A_0 = I_N and U_t is an N x 1 vector of one-step-ahead forecast errors, with $U_t = D_0^{-1}\xi_t$. For a stationary system, the roots of the characteristic equation $|A(z)|$ all lie outside the unit circle, and C(L), which concisely expresses the effects of each shock on the dynamic paths of the components of Y_t, is recoverable from knowledge of A(L).

Since Wold's theorem applies to stationary vector time series, the properties of the data have important implications for the performance of empirical VAR models and the statistical validity of inferences drawn from them. If Y_t is generated by a stationary stochastic process and each equation in the VAR is of the same order, then, under normality, OLS will provide approximate maximum likelihood estimates in large samples (Harvey). In this case, usual classical hypothesis testing of restrictions on the parameters in the VAR are also asymptotically justified and, given identifying restrictions on D_0, estimates of the elements of C(L) in (1) can be recovered from the VAR specification. The presence of a deterministic time trend in each series does not complicate matters since trends in the mean and other deterministic components can be extracted before the statistical analysis begins.

Alternatively, suppose that each series in Y_t is nonstationary because it contains a unit root. Again assuming all deterministic components have been extracted, in place of (3), there exists a finite-order multivariate autoregressive form in differences:

$$A^{\star}(L)(1-L)Y_t = U_t \tag{4}$$

where $A^{\star}(L) = I + A_1^{\star}L + \ldots + A_{p-1}^{\star}L^{p-1}$ and $A_i^{\star} = -\sum_{j=i+1}^{p} A_j$.[1] The lag polynominal $A^{\star}(L)$ is of full rank and all of its roots are outside the unit circle. Hence, the differences model is stationary and there exists a multivariate Wold representation:

$$(1-L)Y_t = C^{\star}(L)\xi_t \tag{5}$$

Several implications of estimating a model such as (3) when the data are generated by model (4) have been highlighted in the literature. First, approximate maximum likelihood estimators of the parameters of the VAR may be difficult to obtain since, as shown by Phillips and Durlauf, OLS will produce consistent parameter estimates but, depending on what deterministic trend components are included in the true model and the regression equation, these estimators may not be asymptotically normally distributed. Second, as a consequence, tests on parameters can have non-standard asymptotic distributions which may have to be calculated numerically unless the null hypothesis of interest can be written as one or more restrictions on mean zero stationary variables or linear combinations of variables (Sims, Stock and Watson). In particular, distributions of sums-of-coefficients and Granger causality test statistics are not easily found. Third, since Y_t is nonstationary, the autoregressive model can not be inverted to recover an infinite-order moving average representation. This invalidates simulations based on the impulse response functions from a VAR that otherwise provide a useful means to estimate the net responses of variables in the model to specific shocks without maintaining strict ceterus paribus assumptions on the evolution of other variables (Pagan).

Before concluding that presence of a unit root in each series implies that the variables in a model should be differenced prior to estimation, a third possibility to consider is that there are linear combinations of the individual series that are stationary. This is the special case of a model with unit roots where the number of such roots is less than the number of variables. In this case, differencing each series individually leads to overdifferencing of a multivariate model.

A stationary linear combination of nonstationary variables can be viewed as a long-run equilibrium relationship among the time series where the linear constraint:

$$\alpha_j' Y_t = 0, \tag{6}$$

is satisfied in the long-run equilibrium. The long-run equilibrium relationship may not hold in all periods. Rather, in some periods it will be the case that:

$$\alpha_i ' Y_t = z_{it} \neq 0 \tag{7}$$

Here, z_{it} represents a deviation from the long-run equilibrium which can be called an error-correction term. The sequence of z_{it}'s is itself a time series variable. Given an initial deviation from equilibrium, the error-correction term is not likely to return to zero in the next period. However, if the series is stationary then there will be a tendency for the long-run equilibrium relationship to be reestablished over time and the variables in Y_t are said to be cointegrated.[2]

Since any cointegrating relationship involves at least two variables, there can be at most $r \leq N-1$ linearly independent cointegrating vectors which produce z_{it}'s that are stationary, where N-r is the number of unit roots in the system. In a two-variable model, there can be at most one cointegrating relationship and, if such a relationship exists, it is unique. For models of more than two variables and more than one cointegrating relationship, however, cointegration can be expressed by any number of equivalent reparameterizations of the cointegrating vectors. For example, in a three-variable system with a single unit root the cointegrating relationships can be expressed by:

$$\alpha' Y_t = Z_t \tag{8}$$

where:

$$\alpha' = \begin{bmatrix} \alpha_1 ' \\ \alpha_2 ' \end{bmatrix} = \begin{bmatrix} 1 & \alpha_{12} & 0 \\ 0 & 1 & \alpha_{23} \end{bmatrix}; \quad Y_t = \begin{bmatrix} y_{1t} \\ y_{2t} \\ y_{3t} \end{bmatrix}; \quad z_t = \begin{bmatrix} z_{1t} \\ z_{2t} \end{bmatrix}. \tag{9}$$

The vectors $\alpha_1 ', \alpha_2 '$ span the space of relationships among the three variables that produce stationary series of error-correction terms. Any linear combination of these vectors would provide an equivalent expression of cointegration.

Granger (1981) and Engle and Granger have shown that cointegration among the variables in Y_t implies restrictions on the coefficient matrices of the moving average representation of Y_t in differences (equation (5)) and on the coefficient matrices of the autoregressive representation of Y_t in levels (equation (3)). In particular, $C^{\star}(1)$ is of rank N-r and A(1) is of rank r. Engle and Granger have also proven a representation theorem that establishes that for cointegrated variables for which there exists a finite-order autoregressive representation in levels, there also always exists a stationary VEC model of the form:

$$A^{\star}(L)(1-L)Y_t = -\gamma Z_{t-1} + U_t \tag{10}$$

where γ is an N x r matrix ($\gamma \neq 0$) such that $C^{\star}(1)\gamma = 0$. [3]

The autoregressive form of the VEC model (10) shows that one-period-lagged deviations from the equilibrium expressed by cointegrating relationships affect the dynamic interactions within a system of autoregressive equations of differences of the variables through the parameters of the γ matrix. In the VEC model, short-run dynamics are flexible while long-run constraints among levels of the variables implied by cointegration are satisfied. The model allows the error-correction process that returns the system to equilibrium to be gradual. Thus, an error-correction model can produce accurate forecasts of responses to disequilibriating shocks over reasonable horizons (Hendry, Nickell).

Presence of cointegration has several implications for estimation, inference, and forecasting. The VEC model provides a basis for efficient parameter estimation which imposes the restrictions implied by cointegration. Since all of the variables in the VEC model are stationary, usual asymptotic justifications for parameter estimates and hypothesis tests apply and the autoregressive model can be inverted, validating dynamic simulations based on the moving average representation. As pointed out above, Sims, Stock and Watson have shown that standard asymptotic distribution theory can also be applied to portions of a model in levels involving cointegrated variables. However, since a VAR in levels of cointegrated series omits rank reducing constraints on A(1), it will produce parameter estimates and forecasts that are asymptotically consistent but are inefficient (Engle and Yoo). A VAR of differenced data is nested within the VEC specification, but is misspecified if the data are cointegrated since it omits the error-correction terms.

A Model of New Zealand Money and Prices

Some Background on the New Zealand Economy

New Zealand is a small and agriculturally trade-dependent nation. The farming sector accounts for around 10 percent of GDP and, including processed agricultural products, around 55 percent of current account receipts. The New Zealand economy also contains a well-developed manufacturing base contributing some 25 percent of GDP. Hence, the sectorial effects of monetary policy and the dynamic relationships between the farming and manufacturing sectors in terms of relative prices, resource allocation, and profitability are important.

Historically, a variety of policies and characteristics of the New Zealand economy have served to insulate it from the world economy. Until 1986, domestic monetary policy and a fixed exchange-rate policy were closely coordinated. Nomin' 'change rate adjustments were undertaken periodically to maintain a relatively stable real exchange rate (Rayner and Lattimore). The manufacturing sector received relatively high levels of import protection, primarily based on import quotas on final goods. The manufacturing sector has also been characterized by a highly centralized wage-setting process and there is evidence that industry has been able to administer cost-plus-markup pricing policies (Chapple). The agricultural sector, in contrast, has been primarily a price taker in international markets. It has been subject to various stabilization schemes but has received relatively low levels of direct producer assistance (Lattimore).

Recently, there has been a general liberalization of economic policy in New Zealand. The fixed exchange rate regime has been replaced by market-determined rates (March 1985), a trend toward exposing the manufacturing sector to greater international competition has been accelerated, and producer assistance and stabilization programs in agriculture have been eliminated (1985).

Stationarity Properties of the Individual Money and Prices Series

The data used for our model of money and prices 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 nominal producer output price indices excluding taxes and subsidy payments. The series are plotted in figure 1.[4] Money and the two price indices have risen over time, with agricultural prices exhibiting more variability than manufacturing prices or money and an apparent tendency for the money and price series not to move too far apart over long periods.

Formal tests of the stationarity characteristics of the individual money and price series are summarized in table 1. For each series, a fourth-order autogressive equation in levels was reparameterized as:

$$\Delta x_t = \beta_0 + \beta_1 t + \beta_2 x_{t-1} + \sum_{i=1}^{3} \beta_{2+i} \Delta x_{t-i} + e_t \tag{11}$$

where e_t is assumed to be distributed N.I.I.D.$(0,\sigma^2)$. These equations were estimated by OLS to test for the presence of a unit root ($\beta_2 = 0$) while allowing for the possibility of a non-zero

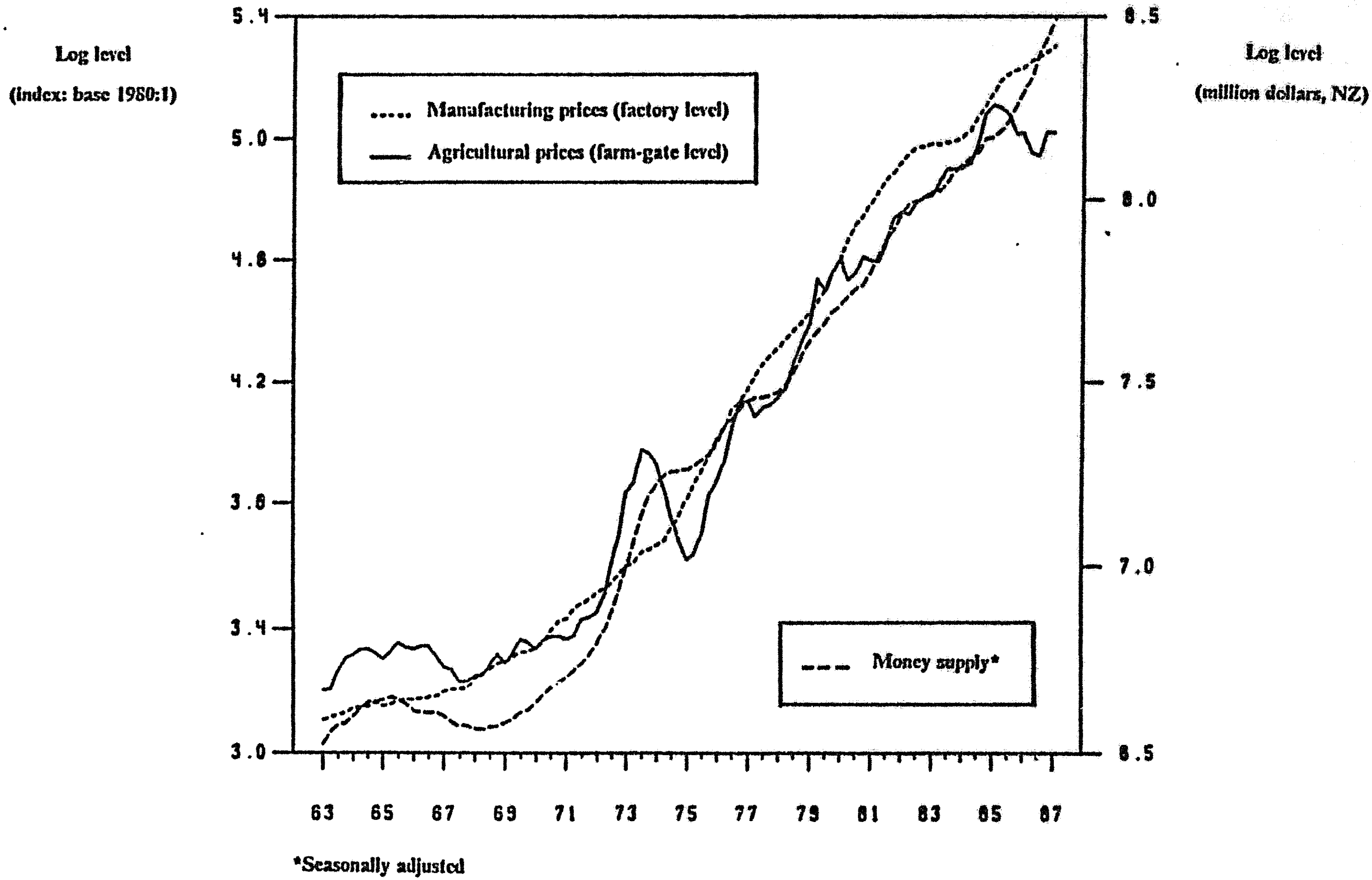

Figure 1. The money supply (M1), manufacturing prices, and agricultural prices in New Zealand, 1963:1-1987:1

Table 1. Time-series properties of money and prices in New Zeland

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

Stationarity Statistics

Hypothesis	Series		
	ΔM1	ΔIP	ΔFP
Unit root ($\beta_2=0$)			
Dickey-Fuller test [a]	-2.666	-2.786	-3.280*
Stock-Watson test [b]	-15.673	-16.087	-18.177
Deterministic trend [c] ($\beta_1 = 0$, given $\beta_2 = 0$)	1.890*	0.765	0.560
Drift [c] ($\beta_0 = 0$, given $\beta_2 = 0$)	2.102**	2.040**	1.725*

Summary and misspecification statistics

Statistic	Distribution	Series		
		ΔM1	ΔIP	ΔFP
$\overline{R}^2$		0.69	0.56	0.25
σ%		1.06	1.09	4.19
HS[d]	F(47,25)	2.02*	1.21	4.28***
ARCH	$\chi^2(4)$	3.13	5.08	8.19*
CHOW[e]	F(5,76)	2.08*	0.52	1.10
SK	$\chi^2(2)$	1.67	0.67	2.00

* reject null hypothesis at the 0.10 level.
** reject null hypothesis at the 0.05 level.
*** reject null hypothesis at the 0.01 level.

[a] Significance level of the test statistic based on the distribution reported in Dickey and Fuller.

[b] Significance level of the test statistic based on the distribution reported in Stock and Watson (1987a).

[c] Test statistic follows standard t distribution.

[d] Test for homoscedasticity over the sub samples 1964:2-1971:4 and 1974:1-1987:1.

[e] Test for post-sample predictive stability for 1986:1-1987:1 with the model estimated over 1964:2-1985:4.

deterministic trend and higher order dynamics in the series.[5] Dickey-Fuller and Stock-Watson (1986) tests of the null hypothesis that the individual series contain unit roots are reported. The results suggest that the hypothesis of a unit root can not be rejected at the 0.10 significance level for M1 and IP based on either test, or at the 0.05 level for FP under the Dickey-Fuller test or 0.10 level under the Stock-Watson test. On the basis of these results, t tests were then conducted on the significance of β_1 (the coefficient on the linear time trend) and β_0 (the constant or drift component) conditional on the maintained hypothesis of a unit root in each series. Only the M1 series exhibits significant trend behavior at the 0.10 level, but a zero drift component is rejected at the 0.10 level for all three series.[6] Of the final univariate autoregressions (with drift but without time trend), the model for FP has the lowest adjusted R^2, as shown in the lower half of table 1, and the estimate regression standard error is higher than for M1 or IP.

Acceptance of the unit root hypothesis is also conditional on the statistical adequacy of the specified autoregressive models. Therefore, results of several misspecification tests for models (11) with the unit root assumption imposed are reported in the lower half of table 1. The results are generally supportive of the statistical validity of the maintained specification. Tests of whether the estimated residual variances are constant over the sub samples 1964:2-1971:4 and 1974:1-1987:1, denoted HS, yield mixed results, with homoscedasticity rejected for the FP regression residuals at the 0.01 level and marginally for the M1 residuals (see Spanos pg. 484 for the test). The ARCH statistic developed by Engle provides an indication of whether the estimated variances behave autoregressively. In our case, the possibility that the variances follow a fourth-order autoregression is not supported, relative to the null that the variances are constant, though there is weak evidence of ARCH effects in the FP equation. Recursive estimation of the parameters of each equation was undertaken to further explore the seriousness of the time dependency problem. The results (not shown) indicated quite stable time paths of the estimators for all three equations, except during the 1972-1974 period. Thus, there appears to be no underlying structural change in the parameters of the models. An indication of this parameter stability is that with the sample period restricted to 1963:1-1985:4, homoscedasticity of the sample and post-sample data is not rejected and, as shown in table 1, Chow tests for predictive stability over 1986:1-1987:1 provide little evidence that changes in the parameters have a statistically significant impact on their predictive performance. Finally, 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 as a χ^2 (2) and are denoted SK in the table.

Tests for Cointegration

Tests for the presence of cointegration among money, manufacturing prices, and agricultural prices in New Zealand were undertaken conditional on the maintained hypothesis that each series contains a unit root and a drift. Stock has shown that OLS theoretically provides consistent estimates of the cointegrating vectors (the α_i in (9)) in static bi-variate regressions of current-dated variables in levels (with a constant included to adjust for differences in means). The residuals from these regressions provide estimates of the equilibrium errors, z_{it}'s, and tests for cointegration among the series are based on testing for stationarity of these series of residuals.

Our estimates of the cointegrating vectors from bivariate regressions among money and prices are reported at the top of table 2. Any two regressions that include all three variables provide estimates of possible cointegrating vectors that span the space of bivariate cointegrating relationships that may exist among the three series. Nevertheless, for completeness we report the results from all of the regressions.

As shown in table 2, the estimated cointegrating parameters are concentrated around a value of one. This suggests that any long-run relationships among the series are almost proportional. Reversing the dependent and independent variables produces approximately

Table 2. Cointegrating vectors and tests of cointegration among money and prices in New Zealand

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

	Dependent Variable					
	M1		IP		FP	
Regressor	IP	FP	M1	FP	M1	IP
Cointegrating regressions						
Estimate of - α_{ij} [a]	0.85	0.95	1.14	1.13	1.04	0.86
R^2	0.98	0.97	0.98	0.98	0.97	0.98
Cointegration tests [b]						
Unrestricted model	-2.88*	-3.65**	-3.00*	-5.41***	-3.77**	-5.45***
Restricted model (α_{ij} = 1)	-2.65	-3.69**	-2.65	-3.29**	-3.69**	-3.29**

* reject null hypothesis at the 0.10 level.
** reject null hypothesis at the 0.05 level.
*** reject null hypothesis at the 0.01 level.

[a] The coefficient α_{ij} is the only unknown coefficient of a bi-variate cointegrating vector (α_i in equation (9)) normalized so the coefficient of the dependent variable is one.

[b] Augmented Dickey-Fuller test for a unit root in the residuals from the cointegrating regression. Significance level of the test statistic based on the distribution reported in Engle and Granger (1987).

inverse parameter estimates (up to 1-R^2, see Engle and Granger, pg. 261). The parameter estimates also come very close to satisfying the transitivity that would arise from determining the coefficient of any third equation algebraically from estimated parameters of any two other equations.

Results of tests for cointegration based on the residuals from the bivariate regressions are reported in the lower part of table 2. The null hypothesis is that there is no cointegration and the residuals from the bivariate regressions contain a unit root. The alternative hypothesis is that the linear combination of variables in the corresponding regression are cointegrated so that the residuals are stationary. The statistical test is a Dickey-Fuller test (allowing for the possibility of higher order dynamics in the error process by including four lagged residuals in the model) adjusted for the dependence of the z_{it}'s on the estimated cointegrating parameters (Engle and Granger). On the basis of this test, we reject the null hypothesis that the residuals contain a unit root at the 0.01 level for an IP,FP regression in either direction (reversing the dependent and independent variables). For the M1,FP regressions, the null hypothesis is rejected at the 0.05 level. The results for the M1,IP regression are less conclusive, but a unit root in the residuals is rejected at the 0.10 level.[7]

Because the initial parameter estimates from the bivariate regressions among money and prices are close to unitary this restriction was imposed on the cointegrating regressions and new residuals (z_{it}'s) were estimated and tested for unit roots. Stationarity of the residuals from cointegrating regressions imposing this restriction implies long-run proportionality between the level of the money supply and the level of each of the nominal prices. The Dickey-Fuller test results shown in table 2 for the restricted model suggest that the money and price series are cointegrated when this restriction is imposed, though again the evidence is weakest (cointegration is not supported) for the estimated M1,IP relationship.

To summarize, our tests of the time-series properties of the data for the money supply, agricultural prices, and manufacturing prices in New Zealand suggest that each series contains a unit roots and is stationary in first differences with a drift. The three variables appear to be cointegrated, with linear combinations of the variables FP-M1, IP-FP, and M1-IP stationary (in logarithms and with means extracted). Thus, the number of unit roots in a system of these three variables is reduced to one. This provides evidence of long-run equilibrium relationships between money and prices. Unrestricted estimates of the cointegrating vectors are consistent with long-run proportionlity of the levels of money and the nominal prices and residuals from the cointegrating regressions are stationary when this restriction is imposed.[8]

A V .:or Error-Correction Model of Money and Price Dynamics

Based on the preceding evidence, a three-variable VEC model (in the form of (10)) was estimated by OLS using first differences of each series, a constant term, and the error-correction residuals from the restricted IP,FP and FP,M1 cointegrating regressions. A third-order lag structure for the system was chosen based on a series of Sims' modified likelihood ratio tests (1980a) and Akaike's Information Criterion. The lag selection test results are reported at the top of table 3.

Basic estimation results and misspecification statistics for the VEC model are presented at the bottom of table 3. The adjusted R^2's from the estimated forecasting equations are higher than from the univariate models (except for the ΔIP equation). There is weak evidence of heteroscedasticity, and ARCH effects in the $\Delta M1$ equation, but less evidence of either effect in the ΔFP equation than in the univariate model. The Chow tests for predictive stability are consistent with stable parameter estimates and the SK statistic indicates normally distributed residuals for all three equations. Together, these results imply that the VEC model provides a reasonably robust statistical model. The evidence from our cointegration tests in support of a VEC model as an appropriate specification for evaluating dynamic interactions among the money and price series is reinforced by rejection of the

Table 3. Summary and misspecification statistics for a vector error-correction (VEC) model of money and prices in New Zealand

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

Lag length selection

Test	Number of lags					
	6	5	4	3	2	1
Modified likelihood ratio test [a]		7.23	9.98	8.36	21.71***	16.93**
Akaike's information criterion	-23.74	-23.84	-23.91	-24.00	-23.93	-23.92

Summary and misspecification statistics

Statistic	Distribution	Series		
		ΔM1	ΔIP	ΔFP
$\overline{R}^2$		0.71	0.54	0.40
σ%		1.03	1.13	3.73
HS[b]	F(47,25)	2.19*	1.19	3.58**
ARCH	$\chi^2(4)$	8.52*	3.07	4.34
CHOW[c]	F(4,70)	1.99	0.99	1.64
SK	$\chi^2(2)$	1.03	0.44	1.78

* reject null hypothesis at the 0.10 level.
** reject null hypothesis at the 0.05 level.
*** reject null hypothesis at the 0.01 level.

[a] Test statistic follows a χ^2 distribution with 9 degrees of freedom.

[b] Test for homoscedasticity over the sub samples 1964:2-1971:4 and 1974:1-1987:1.

[c] Test for post-sample predictive stability for 1986:1-1987:1 with the model estimated over 1964:2-1985:4.

restriction that the coefficients of the error-correction terms are jointly zero. The calculated statistic (not shown in table 3) is $F(6,71)=5.37$, which is significant at the 0.05 level.

Structural Specification and Simulation Results

To trace the dynamic effects of monetary, manufacturing-price, and agricultural-price shocks, the estimated VEC model is reparameterized to its equivalent VAR formulation in levels of the variables. With this reparameterization, the error-correction terms are incorporated into the lagged variables of the autoregressive model. The resulting model can be inverted to obtain impulse response functions that capture the effects of deviations from long-run equilibrium on the dynamic paths followed by each variable in response to initial shocks. In contrast, a standard procedure of inverting the autoregressive components of the VEC model to obtain impulse response functions will misrepresent dynamics of the economy because the error-correction terms will be treated as deterministic constants.

Structural interpretation of the VEC model also requires identifying assumptions about contemporaneous interactions among the variables. We assume a recursive orthogonal order that allows a Choleski decomposition of the covariance matrix of the model's one-step-ahead forecast errors. Impulse response functions and their 90-percent confidence bounds were calculated for several recursive orders using a Monte carlo integration algorithm with 200 draws from the posterior distribution of each response estimate (Doan and Litterman, chpt. 17). The results for the orthogonal order M1,IP,FP are plotted in figure 2 for a twenty-quarter forecast horizon. This orthogonal order allows money supply shocks the most opportunity to affect the price variables by attributing to the monetary shock any contemporaneous correlations between the forecast errors for money and prices.[9]

The point estimates from the impulse response functions in figure 2 show that the dynamic responses of money and prices depend on the source of the initial shock. The long-run percentage responses to a given shock, however, are approximately equal among

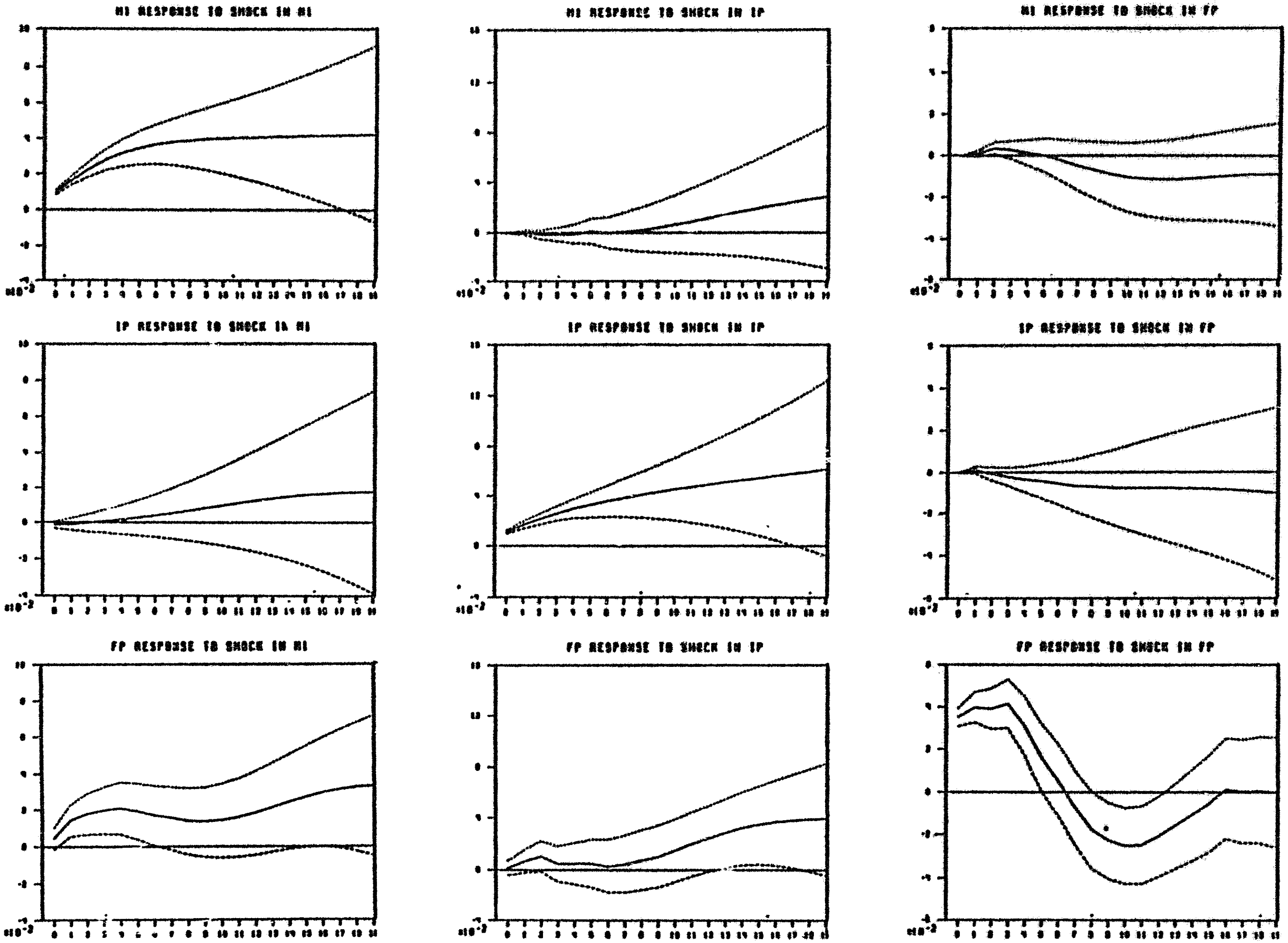

Figure 2. Impulse response functions from a VEC model of money (M1), manufacturing prices (IP), and agricultural prices (FP) in New Zealand

the three variables. This reflects long-run equilibrium and the unitary cointegrating parameters. After any given shock, equality of the point estimates of the impulse responses from the VEC model occurs around 30 quarters.

Interpretation of the point estimates of the impulse responses shown in figure 2 must be qualified because their estimated standard errors increase rapidly.[10] Nonetheless, in terms of whether the posterior probability is concentrated on zero, negative, or positive values, the confidence bounds suggest several general observations about the likely time paths for money and prices after 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. Since agricultural prices move toward the new long-run equilibrium level faster than manufacturing prices, positive monetary shocks induce a shift in relative prices in favor of agriculture in the short run. However, agricultural prices do not exhibit a greater proportional response than the money supply in the short run, nor do they overshoot their own long-run equilibrium level in response to a monetary shock.

2. A positive shock to manufacturing prices also raises the long-run levels of all three series. Because of a relatively slow response of agricultural prices, a shock to manufacturing prices initially places agriculture in a cost-price squeeze. 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 and is not persistent. Neither the money supply nor manufacturing prices shows

any significant short-run response, so there is a short-run shift in relative prices in favor of agriculture. In the long run, in contrast to the permanent shift in the level of each series induced by shocks to the money supply or manufacturing prices, a new equilibrium of each series is established at around the same level 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 and nominal and relative prices eventually return to their initial levels.

The point estimates of the impulse response functions from the VEC model provide an interesting empirical account of money and price dynamics in New Zealand. With respect to policy issues, it appears that monetary policy has affected the agricultural and manufacturing sectors differently over the short to medium term, though in the long run the two price levels move together. Moreover, the response of the money supply to manufacturing price shocks is consistent with the monetary authorities accommodating domestic price increases in a situation in which quantitative import restrictions providing high levels of protection have allowed domestic manufacturing prices to deviate from world price levels. Though we detect little evidence of structural change so far, an important question will be whether this price autonomy and monetary responsiveness will persist if current policies to liberalize manufacturing are pursued.

Our failure to detect any effect of agricultural prices on the money supply stands in contrast to the long-run responsiveness of the money supply to shocks to manufacturing prices and also has implications related to the recent liberalization of the New Zealand economy. Over the sample period, the agricultural sector has been a price taker on world markets under a regime of 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 and provides evidence that the authorities have succeeded with such a policy. Stabilization schemes that locked farm earnings in the Central Bank during periods of high world prices may have contributed to the ability of the monetary authorities to insulate New Zealand from external inflationary and deflationary pressure arising from agriculture and spread over time the effects that agricultural price shocks have on New Zealand's current account. Since 1985, these stabilization programs have been abandoned. Again, if the current liberalization is pursued an important policy issue will be whether past dynamic patterns of responses to agricultural price shocks persist or whether these shocks will induce greater instability in the money supply or manufacturing prices.

In terms of the policy issues raised by our evaluation of the money and price dynamics in New Zealand, it is interesting to compare the results shown in figure 2 to impulse responses from VAR models specified under alternative assumptions about the time series properties of the money and price data. Two VAR models were estimated: a fourth-order model in levels, under the maintained hypothesis of stationarity, and a third-order model in first differences (without the error-correction terms), under the maintained hypothesis of independent unit roots. The point estimates of impulse responses from the VAR in levels are generally similar to those from the VEC model.[11] This is consistent with Engle and Yoo's theorem that when nonstationary variables are cointegrated, parameter estimates from an unrestricted levels model are consistent but inefficient. In the levels model, the point estimates of the impulse responses following any specific shock fail to converge to common values even after 40 quarters.

In contrast to the similarity of the results from the VEC model and VAR model in levels, impulse response functions quite unlike those from the VEC model are computed for the VAR model estimated in differences without the error-correction terms. Estimates from the differences model (reparameterized to evaluate impacts of each shock on levels of the variables) suggest both a more-than-proportionate short-run response of agricultural prices to a money supply shock and that in the short run agricultural prices overshoot their long-run level. The own response to a manufacturing price shock is persistent in the differences model and is largely matched by agricultural prices, while the money supply responds slower and to a lower level even over a long horizon. Finally, own responses to agricultural price shocks are persistent but neither the money supply nor manufacturing prices respond, suggesting a long-run shift in relative prices in favor of agriculture. Inferences that would be drawn from these results are quite different than those drawn on the basis of the impulse response functions from the VEC model.

Conclusions

In this paper, analysis of the stationarity properties of the data has been shown to be important to developing a description of the dynamic relationships among money and prices in New Zealand and to be closely related to evaluation of the hypothesis of long-run monetary neutrality. Tests for stationarity failed to reject the presence of unit roots in the individual series for money, manufacturing prices, and agricultural prices. There is also evidence of long-run equilibrium relationships among levels of the money and the price series that reduces the number of unit roots in a system of equations for these variables from three to one. Further, estimates of the coefficients of the cointegrating vectors reflecting these equilibrium relationships are consistent with the long-run neutrality of money and long-run proportionality between levels of the money supply and each of the nominal prices. Thus, our results provide strong empirical support for the long-run monetary neutrality hypothesis that is widely assumed in theoretical models but the implications of which are often ignored in empirical analysis.

When the long-run equilibrium relationships found to characterize the money and price data are incorporated in a VEC model an interesting data-congruent description of short-run money and price dynamics in New Zealand emerges. In terms of broad hypotheses about these dynamics, we find that agricultural prices respond differently to monetary shocks than manufacturing prices, as postulated by Bordo. The argument that industrial price shocks place agriculture in a cost-price squeeze also receives support. However, we do not find evidence that agricultural prices rise proportionately more than the money supply in the short run nor that they overshoot their own long-run level in response to a monetary shock.

The importance of preliminary evaluation of the stationarity properties of the money and price data and its implications for model specification is illustrated by comparing the impulse response functions from the VEC model with those estimated from VAR models under alternative maintained hypotheses about unit roots. In the New Zealand case, the VEC model provides more precise parameter estimates than an unrestricted VAR model in levels. A VAR model in differences provides impulse response functions that are quite different, and in several ways less plausible, than those from the VEC model. Impulse response functions from the differences model would support different policy inferences than those drawn from the error-correction specification.

While our analysis demonstrates the importance of considering the stationarity properties of the data as part of an evaluation of money and price interactions in New Zealand, some caveats about our results should also be noted. The acceptability of an empirical model should be considered only tentative until it has been successfully tested against new data, different criteria, and rival models. The VEC model we have presented is a small closed system that omits the potential influence of other variables. Inferences drawn from the model may also be limited by the recursive structure imposed. Assuming that the data itself is rich enough to reveal the underlying phenomena of interest, the development of larger dynamic models with more appealing contemporaneous structures is a priority for future research.

Footnotes

[1] The relationship between A(L) and $A^{\star}(L)$ can be seen by reparameterizing $A(L)Y_t$ as

$$A(L)\,Y_t = IY_t + [A_1 + \dots + A_P]Y_{t-1} + [-A_2 - \dots - A_P](Y_{t-1} - Y_{t-2})$$

$$+ \dots + [-A_P]\,(Y_{t-(P-1)} - Y_{t-P})$$

When each series contains an independent unit root $-[A_1 + \dots + A_P] =$ and the model can be rewritten in differences as in (4). The relationship between C(L) and $C^{\star}(L)$ is less direct, since an infinite-order moving average representation does not exist for the levels of a nonstationary Y_t. Instead, assuming $\xi_t = 0$ for $t \leq 0$, from (5):

$$Y_t = (1 - L)^{-1} C^{\star}(L)\,\xi_t$$

which yields:

$$Y_t = C_t(L)\,\xi_t$$

where $C_t(L) = C_0 + C_1L + \dots + C_{t-1}L^{t-1}$ and $C_i = \sum_{j=0}^{i} C_j^{\star}$.

[2] Formally, the elements of a vector Y_t are said to be cointegrated of order (d, b) if all components of Y_t are integrated of order d (denoted I(d); for example, series that are nonstationary because they contain a unit root are I(1)), and there exists a vector $\alpha \neq 0$ such that $\alpha' Y_t = z_t$ is I(d-b), b > 0. The vector α is called the cointegrating vector.

[3] In the case of cointegration, $-[A_1 + \dots + A_P] \neq I$. The reparameterized model $A(L)Y_t$ can be expressed as:

$$A(L)Y_t = I(Y_t - Y_{t-1}) + [I + A_1 + ... + A_P]Y_{t-1} + ... + [-A_P](Y_{t-(P-1)} - Y_{t-P})$$

Comparing this expression to (10), one can see using (7) that:

$$\gamma\alpha' = [I + A_1 + ... + A_P] = A(1)$$

which is part of Granger's representation theorem.

[4] The existence of seasonality in the series was investigated 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.

[5] The lag length for the univariate models was chosen on the basis of a series of Sim's modified likelihood ratio tests at the 0.05 probability level. For each series a fourth-order model in levels was selected.

[6] As discussed in the introduction, there is widespread evidence that unit roots are common among aggregate macroeconomic variables. The presence of a drift parameter and a unit root in the autoregressive model of a time-series suggests that the series consists of a deterministic linear trend, a random-walk component, and a stationary component. The inclusion of a time trend suggests that there is a nonlinear deterministic time component of the series in levels in addition to the other processes. For money and prices specifically, additional evidence on the time series properties of aggregate data has been reported for the U.S. in Stock and Watson (1987b). For the sample period 1960-1985, they

find that nominal M1 is stationry in differences around a drift and linear time trend and the wholesale price index is stationary in differences with a drift. As in our case, these results rest on the assumption that the underlying process is at least approximately an autoregressive model. Schwert (1987) also finds evidence for unit roots in several money and price series for the U.S. based on the autoregressive model. However, he suggests that a moving average process may be a more appropriate model. We find no evidence to suggest the autoregressive model is inconsistent with the data for money and prices in New Zealand.

7 By transitivity, if M1,FP and FP,IP are cointegrated, then, equivalently, IP,FP are also cointegrated. Differences in the test statics may be due to the low power of the tests. Also, Banerjee, Dolado, Hendry, and Smith find that, in practice, tests based on a static bivariate regression often lead to an erroneous over rejection of cointegration.

8 Given this evidence of long-run neutrality of money and proportionality of the level of the money supply and the level of each of the nominal prices, one might anticipate little impact of money on real output in New Zealand in the long run. Using real GNP as a proxy for output, in a third-order univariate autoregression we find evidence of a unit root in the real output series (the Dickey-Fuller test statistic is -2.1 and the Stock Watson test statistic is 14.3 which are not significant at the 0.10 level). We fail to reject the hypothesis that money and real output are not cointegrated at the 0.10 level of significance. We also reject Granger causality from money growth to real output growth F(8,73) = 2.02 in a bivariate VAR which is not significant at the 0.10 level). This evidence corroborates the monetary neutrality results from our model of money and prices.

[9] The Choleski decomposition has been the common means of identification of VAR models and choice of an orthogonal order is an important part of model interpretation. Granger causality tests provide some useful information about lag relationships among the variables. In our case, for models in levels of the variables, the p-values of the F-tests for lags of M1, IP and FP are, respectively, 0.00, 0.08 and 0.01 in the M1 equation; 0.27, 0.00 and 0.40 in the IP equation; and 0.12, 0.01 and 0.00 in the FP equation. These results are suggestive of feedback relationships among money and prices, the magnitude and direction of which can be evaluated efficiently and conveniently based on the impulse response functions from the VEC model. The Granger causality results are also consistent with our cointegration tests, since cointegration implies Granger causality in at least one direction (Granger 1986). However, the Granger causality tests provide little in the way of explicit criteria for choosing an orthogonal order since they do not provide any information on the contemporaneous interactions among the variables. Choice of an orthogonal order must rest on a priori logical grounds. Comparing impulse response functions for alternative orthogonalizations of a VAR or VEC model may then shed some light on the importance of different assumptions about the contemporaneous correlations among the variables. For our model, the results are largely invariant to the orthogonal order, with impulse responses and their confidence bounds for two alternatives (IP,FP,M1 and FP,IP,M1) almost identifical to those shown in figure 2. Nonrecursive identifying assumptions can also be considered, as in Blanchard, Bernanke, Sims, or Orden and Fackler, but were not examined in our study.

[10] 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.

[11] Results from the levels and differences models are not shown but are available from the authors on request.

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