Inflationary Dynamics and the Angell-Johnson Proposals

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I. Introduction

During the 1980s, the United States economy has undergone enormous change. Among the more notable changes are the emergence of large federal budget and international trade deficits, high real interest rates, and a movement to rates of inflation and real commodity prices much lower than those of the late 1970s. Trade weighted measures of the dollar exchange rate rose to historically high levels by February 1985, and then declined precipitously.

While the Federal Reserve has maintained a fairly conservative money supply posture in its fight against inflation, fiscal authorities have more than doubled the national debt since 1981. Because of the fiscal deficit, monetary policy is the only major macro policy tool that can presently be used with any degree of flexibility.

In response, three key Washington policy makers have called for changes in the way monetary policy is conducted both here and abroad. Citing the Federal Reserve Board's goal to stabilize the price level in the face of increasingly volatile money velocity, Federal Reserve Governor Wayne Angell emphasizes that the need for reliable early warning indicators of inflationary pressures. Angell proposes determining the rate of expansion in bank reserves and the basic money supply according to the movements in a group of basic commodity prices [Angell, 1987]. Commodity prices are attractive for two basic reasons. First, they are reported daily and thus available more than a month in advance of the CPI estimates. Second, most commodity prices fluctuate freely in response to monetary disturbances and other market conditions, providing a more sensitive inflation signal than *sticky* production costs such as wages.

Commodity prices may serve as barometers for inflationary pressures. If prices are expected to rise in the future, commodity prices must be bid up today in order to bring the returns to commodity holdings down to competitive levels [Angell, 1987, pp. 9-10].

In addition to Governor Angell's proposal, Federal Reserve Vice-Chairman Manuel Johnson argues in favor of using commodity prices as well as financial auction market variables such as the spread between long-term and short-term interest rates and a trade-weighted exchange rate for the dollar as aids in developing inflation forecasts [McNamee, 1988]. Former Treasury Secretary James Baker advocates exploring a similar approach of using indicators to stabilize international exchange rates among the *Group of Seven* industrialized countries (Canada, France, Great Britain, Italy, Japan, U. S., and West Germany) [Kilborn, 1988]. One common aspect of all three proposals is their use of commodity prices as an informational signal rather than as a policy target..

To implement such policy initiatives presupposes a consistent link between basic commodity prices, interest rates, the international value of the dollar and U. S. inflation. The purpose of this paper is to examine the statistical relationships between movements in the consumer price index and the prices identified by the Angell-Johnson proposals as appropriate guides to monetary policy. If the variables move together in a consistent and predictable fashion, then commodity price movements can presumably serve as signaling mechanisms for the development of monetary policies under the Angell-Johnson proposals.

While Angell focuses on the relation between turning points of commodity price inflation in a number of existing commodity indices and turning

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points in the general inflation rate, this paper will consider the role of individual commodities and aggregates of related commodities. In practice, a simple examination of turning points may not be as useful as estimates of the complete relationship between the series. While turning points are easy to recognize, *ex post*, reliably identifying them as they occur generally requires several months of data. Thus, much of the advance warning disappears. Further, insight may be gained on the types of commodities and lag lengths likely to be useful in constructing indices specifically aimed at serving as leading indicators of inflation.

Subsequent sections of the paper include a survey of recent studies related to these policy proposals, a description of the methodology employed, empirical results, and conclusions.

II. Recent Research

Unfortunately, there is little generally agreed-upon macro-theory linking price level reactions to cost factors, demand factors, and expectations in a way which can both be derived from the aggregation of firms, micro-behavior and explain observed macroeconomic data. This hinders adequate testing of competing theories, allows theorists to specify a wide variety of models, and has left econometricians free to investigate the relationships between individual price measures and general inflation rates in a variety of ways. The Angell-Johnson proposals fall squarely within this a-theoretical tradition. As might be expected, recent papers examining the Angell-Johnson proposals utilize a wide variety of techniques to test the usefulness of potential leading indicators of general inflationary trends.

Whitt [1988] develops a model in which commodity prices can provide information regarding financial shocks in the economy. Commodity prices are useful in his model because they are immediately observable. The equations he derives indicate that commodity prices and the rate of growth in the money supply may each serve as predictors of changes in the inflation rate under different assumptions. When the velocity of money is subject to unexpected variations or the money supply is susceptible to shocks that the monetary authorities cannot anticipate, then commodity prices will serve as better indicators of future inflation.

Whitt's empirical tests utilized quarterly percentage changes of three inflation measures regressed against percentage change of the narrowly defined money supply, and the percentage change in the Raw Industrial Commodity price index. Almon polynomial distributed lags, with a Cochrane-Orcutt autocorrelation correction technique, were used to estimate equation parameters. Whitt's results indicate that the money supply serves as a good indicator of inflation during periods of price stability. Commodity prices provide better information regarding future movements in inflation for the period when the demand for money is unstable (1975-87).

Boughton and Branson [1988] also provide a theoretical basis for a leading indicator role for commodity prices. Using a two-sector model (commodities and manufactures), they assume commodities are traded in auction markets while industrial prices are administered by firms and respond only gradually to excess demand. Thus, commodity prices can react immediately to inflationary pressures and lead the price of manufactures.

In the case of an unanticipated monetary disturbance, commodity prices initially overshoot the equilibrium which will be attained once all prices have fully adjusted and will lead the price of manufactures. However, for a real disturbance (e.g., a supply shock), commodity prices may undershoot their eventual equilibrium level and actually lead an opposite direction change in the price of industrial products. Essentially, agents use the goods with freely variable prices (commodities) as hedges against anticipated inflation in the gradually adjusting prices. For the more realistic case in which other prices fluctuate freely (e.g., financial variables such as interest and exchange rates), the question of which prices serve as the most reliable inflation indicators is an empirical one.

Empirically, Boughton and Branson try to explain the movement of an aggregate inflation rate of the seven largest industrial countries using commodity indices which include the price of oil. They find no stable long-run relationship between the *levels* of commodity prices and the CPI. Since the order of integration of commodity prices is different than that of inflation, the series cannot be cointegrated. Moving to short-run relationships, they find some evidence that commodity prices Granger-cause inflation when oil is included in the indices (the evidence is much weaker when oil is excluded). Additionally, commodity prices are found to be more significant predictors of inflation than monetary growth in polynomial distributed lag regression estimates.

Garner [1988] is primarily concerned with defining the appropriate role for commodity prices in U. S. monetary policy. Some economists and policy-makers advocate using commodity prices as a target for monetary policy. This presupposes the type of stable link between the *levels* of commodity prices and consumer prices that Branson and Boughton fail to find. Garner concurs that such a relationship does not exist. Without such a linkage, using monetary policy to control the level of commodity prices will not necessarily control consumer prices.

The Angell-Johnson proposals, however, rely on commodity prices as purely informational variables which may improve the forecasting performance of econometric models or signal a shift in the relationship between monetary targets and inflation objectives. Garner finds evidence that commodity price indices Granger-cause inflation even when financial variables are included in the information set. Variance decompositions based on vector autoregressions provide further support for these conclusions.

Webb [1988] also conducts causality tests regarding movements in the consumer price index with respect to two monthly commodity price indices. His results indicate that commodity prices Granger-cause changes in overall inflation in the U. S. economy.

Vector autoregression models are then estimated to test the usefulness of the commodity price indices as predictors of consumer prices. The statistical results reported imply that models containing the two indices yield results only marginally superior to models containing traditional inflation forecast variables.

Among the papers dealing exclusively with commodity prices, Baillie [1988] is the only one to conclude that they offer no useful information with respect to future inflation. Baillie further claims that consumer price indices in the U. S. and five other industrial economies cannot be readily modeled with ARIMA techniques due to stationarity problems. The absence of bivariate cointegration between the consumer price index and any of six commodities (gold, beef, soybeans, coffee, corn, and cotton) leads Baillie to hypothesize that these prices contain little information concerning changes in aggregate measures of inflation. This conjecture is subsequently supported by empirical evidence provided by vector autoregression modeling results.

Koch, Rosensweig, and Whitt [1988] investigate one aspect of the Johnson proposal, namely the relationship between the exchange rate of the dollar and U. S. inflation. The authors suggest that expectation effects may cause exchange rates to temporarily lead changes in overall price levels, a hypothesis that Granger causality tests fail to reject. They conclude that movements in the dollar significantly impact consumer prices and do so with a lag greater than 36 months.

III. Methodology

This study does not attempt to test any specific theory or impose any *a priori* modeling restrictions on the equations that are estimated. Multiple transfer ARIMA analysis is used to calculate the models. The Box-Jenkins methodology selection is motivated by the lack of information regarding possible explanatory variables and appropriate lag structures. Zellner and Palm [1974] have shown that under certain conditions, ARIMA models can represent reduced forms of structural models.

Another consideration which supports the usage of multi-input ARIMA models is the time series nature of the data. Granger and Newbold [1973] point out that *spurious* regressions often result from ordinary least squares performed on series that are highly autocorrelated. Because the nominal price indices used in this study are all serially correlated, differencing is used to transform them into stationary series. Pierce and Haugh [1977] have shown that Box-Jenkins models estimated for such working series will preserve the causal relationships that exist between the variabies. Transfer functions thus seem to provide an efficient means of exploiting both the time series properties and the structural components of the relevant relationships [Box and Jenkins, 1976, ch. 11].

Transfer function modeling requires that cross correlation functions (CCF) be estimated for the prewhitened output and input series. Input series may be thought of as independent variables or regressors and output series as dependent variables. The CCFs are used to suggest the lag structure and functional form of each bivariate transfer relation. CCFs can be estimated accurately only for stationary series (a stationary series is one whose mean and variance are constant through time). Stationarity can be induced in a working series by differencing and/or transforming the observations [Pankratz, 1983, pp. 189-90].

In addition, ARIMA modeling seems attractive relative to vector autoregression (VAR) for this application since the multivariate models often include five or six monthly input series with long lags. The use of ACFs and CCFs to suggest to lag structure restricts the actual model estimation task to a reasonable amount of parameters. VAR techniques would require the estimation of a very large number of parameters. VARs would be more appealing when considering only a single commodity index (rather than several individual commodities) and only one or two financial variables in each model (see Garner [1988] for V AR estimates of this type).

Prewhitening removes the influence of past values of a variable on its current value by eliminating serial correlation in the series. This implies that the CCF between two prewhitened series provides information about the explanatory power of the input variable after all of the explanatory power of the output variable has been taken into account. Prewhitening is accomplished in this paper by estimating univariate ARIMA models for every series being analyzed. CCFs are then calculated between the CPI model residuals and the univariate input model residuals [Chatfield, 1984, p. 173].

Schwert [1977] suggests that the "pattern and magnitude" of the relationship between the cross correlated residuals can differ from the relationship between the original stationary variables. To allow for this possibility, CCFs are also calculated directly between the working series for the CPI and each of the differenced input series [Box and Jenkins, 1977,pp. 372-74].

Based on the CCFs of the first methodology, bivariate transfer ARIMA functions are calculated for the output variable. When necessary, alternative bivariate models are tested using the CCFs of the second methodology. The maximum lag length considered is 38 months.

The final step in the procedure is estimating multivariate transfer functions for the consumer price index (CPI). Although economic theory gives little information about the dynamic mechanisms governing inflation rate movements, it does suggest that multivariate models include explanatory variables from different sectors of the economy. It is attempted, therefore, to estimate models incorporating variables from a variety of primary goods sectors. To reflect the Johnson proposal, models are also estimated using financial price series. Finally, a few equations are calculated using standard economic theory, departing somewhat from the Angell-Johnson proposals.

Monthly data (obtained from The WEFA Group) are analyzed for the period beginning in January 1972 and continuing through January 1988. This period is chosen to coincide with the effective termination of the Bretton Woods agreement following the decision to suspend dollar convertibility into gold in August 1971. It is an interesting period to analyze because of the oil shocks of 1973 and 1979, and the disinflation brought about in the U. S. by restrictive monetary policy in the early 1980s.

The general form of the transfer models estimated is the following:

$$C_t = A^{-1}(B)\,[w(B)P_t + g(B)S_t + r(B)R_t] + Q(B)U_t \qquad (1)$$

where C_t represents the stationary working series calculated for the CPI. P_t, S_t, R_t represent the stationary working series for each of the respective input variable categories: commodity prices, the interest rate spread, and the trade-weighted exchange rate. $w(B)$, $g(B)$, and $r(B)$ are the moving average polynomials associated with each respective input category. B represents the backshift, or lag, operator. $A^{-1}(B)$ and $Q(B)$ are general order autoregressive and moving average polynomials, respectively. U_t is the error term associated with the process.

The adequacy of the univariate, bivariate, and multivariate models is tested by determining whether their residuals could have been generated by a white noise process. To test this hypothesis of no serial correlation in the innovations, the Box-Pierce Q statistic is used.

While no specific lag structures are imposed, several assumptions suggested by standard economic theory are tested. Commodity prices with short lags are hypothesized as having positive impacts on the CPI.

For longer lags, *a priori* expectations of coefficient signs cannot be formulated with any degree of confidence. To see this, consider an input series which *leads* an output series in a cyclical fashion. If the lag used in the model is sufficiently long, it will create a half-cycle phase displacement, reversing the direction of the relationship between the series. Alternatively, sign reversals for long lags may simply indicate that the effects of the shock are smaller in the long- run than in the short run.

Further, the price index is expected to move inversely with respect to exchange rate. The interest rate spread, defined as the difference between 30-year Treasury bond and 30-day Treasury bill interest rates, is expected to be negatively correlated with changes in the CPI.

IV. Empirical Results

Univariate ARIMA Equations

Most of the series analyzed are successfully modeled by relatively straight forward and parsimonious univariate equations. Twenty-six of the price series are stationary after one degree of regular differencing. The CPI and money supply (currency plus demand deposits) series require seasonal as well as regular first order differencing in order to induce stationarity. Eleven of the univariate models contain statistically significant positive constant terms. In essence, this can be interpreted as evidence of a strong positive trend in the month-to-month variations in the nominal levels of these series.

The univariate CPI model is used as the starting point for the transfer function models where the input variables are included to help model the CPI working series. This equation, after regular and seasonal differencing, contained an autoregressive term at lag one, a moving average term at lag 9, and a seasonal moving average term at lag 12. The Q-statistic calculated at 38 lags for this model is 37.885, indicating that only white noise is present in the residuals. Univariate modeling results for the other series are available from the authors.

Bivariate Modeling Results

The CCFs estimated for the univariate ARIMA residuals and for the differenced data series indicate a wide variety of potential lead times for the various input series with respect to movements in the CPI. Some of the results, such as the CCFs for CPI and the spread between 30-year Treasury bond and 30-day Treasury bill interest rates, exhibit relatively short and direct leading relationships. Other CCFs, such as those for CPI and crude oil, indicate very long and complicated effects resulting from movements in the input price series. Detailed output for the CCFs is available from the authors.

Table 1 summarizes the transfer function modeling results for the differenced CPI series and a number of input series with lags suggested by their respective CCFs. These models indicate which variable lags should initially be included in the multivariate transfer models. Column 1 contains the name of the input series.

Numbers in parentheses represent the lags which are statistically significant for purposes of modeling the stationary CPI series. The series which do not provide significant lags in modeling CPI are denoted with hyphens. Column 2 contains the ARIMA process utilized in conjunction with the input series to model the output variable.

Notation is the well known Box-Jenkins convention represented by *(a;d;q)x(A;D;Q)*. *a* represents the autoregressive (AR) terms in each model, *d* the degree of regular differencing required to induce stationarity in the series being modeled, and *q* the moving average (MA) terms. *A* represents the seasonal autoregressive (SAR) terms, *D* the degree of seasonal differencing used

TABLE 1

CPI Bivariate Transfer Function Models

Input Series	ARIMA Specifications
1. Exchange Rate (-14)	(1;1;9) x (0;1;12)
2. Gold (-11,-17,-24,-30)	(1;1;9) x (0;1;0)
3. Silver (-6)	(1;1;9) x (0;1;0)
4. Metals (0, -24)	(1;1;9) x (0;1;12)
5. Crude Oil (0, -5, -20, -12)	(1;1;9) x (0;1;12)
6. Fuels and Energy (0)	(1;1;9) x (0;1;12)
7. Wheat (-12)	(1;1;9) x (0;1;12)
8. Com (-10)	(1;1;9) x (0;1;12)
9. Grains (-9)	(1;1;9) x (0;1;12)
10. Soybeans (-1, -12)	(1;1;9) x (0;1;12)
11. Livestock (-11)	(1;1;9) x (0;1;12)
12. Farm Products (-11, -22)	(1;1;9) x (0;1;12)
13. Processed Foods & Feeds (0, -11)	(1;1;9) x (0;1;12)
14. Lumber (-8)	(1;1;9) x (0;1;12)
15. Lumber & Wood Products (-7)	(1;1;9) x (0;1;12)
16. Plant & Animal Fibers (---)	
17. Textiles (-10)	(1;1;9) x (0;1;12)
18. Crude Rubber (-13)	(1;1;0) x (0;1;12)
19. Rubber & Plastic Products (---)	
20. Chemical Products (0, -12, -22)	(1;1;9) x (0;1;12)
21. Paper Products (-1, -20, -24)	(1;1;9) x (0;1;12)
22. Machinery & Equipment (-10)	(1;1;9) x (0;1;12)
23. Hourly Earnings (-8)	(1;1;9) x (0;1;12)
24. Per Capita Output (-20)	(1;1;9) x (0;1;12)
25. Producer Price Index (0, -1, -25)	(1;1;9) x (0;1;12)
26. Interest Rate Spread (-2)	(1;1;9) x (0;1;12)
27. Money Supply (-5)	(1,2;1;9) x (0;1;12)

to induce stationarity, and Q the seasonal moving average (SMA) terms.

Elements in each parameter category are separated by commas, the numbers represent the lag associated with each term. Because the data are monthly, the order of seasonality is 12. These results are presented, without any diagnostic statistics, to illustrate the dynamic relationships between the CPI and its components. Examples of the detailed bivariate modeling output are also available from the authors.

Results in Table 1 suggest a wide variety of interesting lead relationships to exist between the working series for CPI and the explanatory variables. Several variables such as soybean and energy prices have almost immediate impacts on the U. S. inflation rate. Others such as gold, farm, and paper prices still influence inflation more than 20 months after their initial movements.

The Angell Proposal

As mentioned above, Angell [1987] proposes the development of a commodity price guide to serve as a signalling mechanism with respect to price stability and U. S. monetary policy. Angell points out that the variability associated with monetary velocity (GNP/M1) in recent years makes it difficult to predict inflation using the money supply alone. Commodity prices, he argues, provide early indicators of general price trends. This section of the paper tests whether there exist groups of commodities whose price movements can jointly explain changes in the CPI.

There are a variety of models that yield statistically significant results with interesting implications for policy makers to consider. Federal Reserve Board Chairman Alan Greenspan warns that changes in the prices of commodities subject to large transitory shocks may have quite different policy implications from price changes symptomatic of a more widespread boom in commodity prices [Greenspan, 1987, p.5]. In utilizing commodity indicators, it is crucial to distinguish between industry specific demand factors or supply shocks likely to cause transitory or one-time adjustments in the price level and commodity price changes indicative of more general aggregate demand conditions. Using a large set of commodities or a general index may yield a more reliable signal of aggregate demand conditions.

In this context, the issue of the importance of oil and gold relative to other commodities is one of the foremost controversies concerning the use of a commodity indicator. Since oil is subject to cartel-induced supply restrictions and plays such an important role in driving input prices in many sectors of the economy, it seems appropriate to consider oil separately from an index of more competitively supplied commodities. As for gold, it may have special importance as the most traditional commodity indicator of a flight from currency.

Angell model results are reported in Table 2. The Q-statistics are not reported because all were superior to that of the univariate CPI model. Several bivariate transfer function models using oil and gold prices were constructed to attempt to empirically resolve the issues surrounding gold and oil prices.

Oil price changes have a substantial and intuitively plausible effect on changes in the CPI. In one model, oil prices appear with statistically significant positive coefficients contemporaneously and with a five month lag (Model 1).

Alternatively, the five month lag can be replaced by a twelve month lag with a statistically significant, but negative, coefficient (Model 2). Both models achieve a notable reduction in the standard error of regression (SER) when compared with the univariate CPI model.

The same cannot be said for the models using gold prices. One such model has a significant positive coefficient at a lag of eleven months and an insignificant positive coefficient at a seventeen month lag (Model 3). The alternative gold model (Model 4) has significant negative coefficients at lags of twenty-four and thirty months, but no longer shows a significant relationship for the eleven month lag. However, these models leave the SER of the univariate CPI model essentially unchanged. These results concur with Garner's [1988] finding that while several general commodity indices Granger-caused the inflation rate even when a monetary variable was included in the equation, the price of gold often fails to Granger-cause the CPI.

With the traditional role of gold as a short-run inflation indicator in doubt, it appears even more plausible to investigate whether or not other commodities may play this important role. Using the bivariate results as a basis for further work, a wide variety of multivariate commodity models were estimated in order to assess the validity of the Angell proposal.

Many of the multiple input equations constructed along the lines of the Angell hypothesis contained coefficients with negative signs. Model 5, however, exhibits results that reflect well upon the Angell hypothesis. Oil prices have an immediate impact on the CPI as well as a delayed impact 5 months after the initial variation. That the oil price index has its greatest impact at lag 0 may reflect rational usage of the information contained in these components of overall inflation. The index for farm prices affects consumer prices with a lead time of 11 months.

TABLE 2

Angell Proposal Transfer Function Models

Model	Regressors					
1.	OIL (0)	OIL (5)	AR (1)	MA (9)	SMA (12)	
	0.109	0.063	0.655	0.186	-0.652	*
	5.735	3.303	11.175	2.493	-7.088	**
	S.E.R. = 0.598		D.F. = 156			
2.	OIL (0)	OIL (5)	AR (1)	MA (9)	SMA (12)	
	0.105	-0.093	0.550	0.204	-0.848	
	6.372	-5.582	10.595	3.473	-11.141	
	S.E.R. = .0556		D.F. = 174			
3.	GOLD (11)	GOLD (17)	AR (1)	MA (9)	SMA (12)	
	0.011	0.007	0.642	0.086	-0.883	
	2.193	1.537	12.180	1.364	-10.630	
	S.E.R. = 0.618		D.F. = 150			
4.	GOLD (24)	GOLD (30)	AR (1)	MA (9)	SMA (12)	
	-0.015	-0.011	0.613	0.107	-0.850	
	-3.012	-2.322	11.624	1.695	-10.544	
	S.E.R. = 0.612		D.F. = 156			
5.	OIL (0)	OIL (5)	FARM (11)	AR (1)	MA (9)	SMA (12)
	0.107	0.056	0.033	0.639	0.211	-0.649
	5.694	2.864	1.557	10.781	2.786	-7.004
	S.E.R = 0.622		D.F. = 151			

* First row contains parameter estimates.
** Second row contains computed t-statistics.

The Johnson Proposal

Vice Chairman Manuel Johnson's proposal expands the Angell hypothesis beyond its relatively narrow focus to incorporate variables reflecting domestic monetary policy and international exchange markets. Therefore, the models reported in this section of the paper include all of the inputs used in the transfer ARIMA equations above plus the exchange rate and the interest rate spread (see Table 3 for statistical results). The latter serves as a proxy for measuring the degree of tightness in Federal Reserve monetary policy. Certainly, if commodity price indicators are to be useful, they must make an incremental contribution to predictive power when the models include these financial variables.

TABLE 3

Johnson Proposal Transfer Function Models

Model	Regressors					
6.	REX (14)	AR (1)	MA (9)	SMA (12)		
	-0.051	0.625	0.189	-.813	*	
	-2.630	11.655	2.959	-10.157	**	
	S.E.R. = 0.626			D.F. = 159		
7.	SPRD (2)	AR (1)	MA (9)	SMA (12)		
	-0.139	0.621	0.144	-0.915		
	-2.077	12.731	2.510	-11.881		
	S.E.R. = 0.606			D.F. = 175		
8.	OIL (0)	OIL (5)	GOLD (17)	GRAIN (11)	SOY (12)	REX (14)
	0.112	0.062	0.007	0.029	0.007	-0.021
	5.996	3.218	1.457	3.291	1.508	-1.052
	AR (1)	MA (9)	SMA (12)			
	0.648	0.186	-0.648			
	10.831	2.451	-6.874			
	S.E.R. = 0.617			D.F. = 148		
9.	OIL (5)	OIL (20)	FOOD (11)	SPRD (2)	REX (14)	AR (1)
	0.040	0.039	0.125	-0.098	-0.038	0.619
	2.102	1.887	3.998	-1.353	-2.026	11.111
	MA (9)	SMA (12)				
	0.204	-0.761				
	2.902	-8.738				
	S.E.R. = 0.601			D.F. = 158		

* First row contains parameter estimates.
** Second row contains computed t-statistics.

The bivariate models for the exchange rate and for the net spread between long-term and short-term interest rates are of particular relevance for the Johnson Proposal. Model 6 uses the trade-weighted dollar exchange rate lagged 14 months as the input variable. The coefficient is significantly different from zero, and, as expected, appears with a negative sign. Such a long lag is quite plausible given the typical behavior of importers. As the dollar depreciates in the short-run, foreign sellers are often reluctant to increase prices in order to avoid losing market share in the event that the exchange rate fluctuations are transitory.

As for the interest spread, Model 7 displays a significant negative coefficient at a lag of two months. This is consistent with the supposition that restrictive monetary policy can lead to high short-term interest rates relative to long-term rates. The relatively short lag supports the hypothesis that higher short-term interest rates can quickly divert liquid funds away from consumer and business spending.

Model 18 presents results that are in line with the premises underlying the Johnson proposal. Coefficients for the oil, gold, grains, and soybean stationary series are all positive and the coefficient for the exchange rate working series is negative. The t-statistics computed for several of the parameters do not satisfy the 5-percent significance criterion, but the bivariate results indicate that this may be due to multicollinearity. (In general, no more than 5 or 6 input series could maintain significant, independent predictive power in the same model.) The equation is also interesting because four of the series have fairly long lead times with respect to the CPl.

Model 9 contains a 20 month lag for the oil price appearing with a positive coefficient. Other variables with the hypothesized signs associated with their parameter estimates include oil at lag five, food at lag 11, the interest rate spread at lag two, and the exchange rate at lag 14. This model is not without problems, however, as the spread coefficient is not significant at the 5-percent level.

In general, the models appearing under the Johnson heading give strong support for the proposition that commodity price indicators can provide information which is incrementally useful with respect to several key financial variables. In fact, the commodity prices seem to be better predictors of future changes in the rate of inflation than the exchange rate and the net interest spread in many of the models. While many of the commodity prices maintain independent predictive power when the models control for financial variables, the opposite often fails to hold.

General Models

In addition to models falling directly under the Angell or Johnson headings, it is also useful to investigate whether various commodities can provide incrementally useful information when added to models incorporating traditional inflation indicators such as wages and capital costs. One such equation includes oil, wages, and the exchange rate (Model 10). These variables serve as proxies for several of the major sectors of the economy: energy ,labor, and imports; all represent factors traditionally considered by economic forecasters when developing inflation outlooks.

Using Model 10 as a benchmark, it can be seen that the expanded specifications can substantially improve the fit of the equations. In terms of the SER, Model 11 obtains the best fit. Machinery and food prices are added and further expand the scope of the model by proxying for capital costs and the agricultural sector. Since the average lag of the input series in this model is almost six months, this model offers the potential for inter-mediate-term advance forecasting. The addition of gold to this model fails to improve its fit (Model 12), supporting the earlier contention that gold prices may not be very helpful.

Traditional econometric models used to forecast inflation frequently include the money supply as one of the predictors. Two of the models presented in Table 4 contain the stationary component of the money supply as an input variable. When short lags of the money supply are included, the estimated coefficients are statistically insignificant (see Model 13). Model 14 contains a positive coefficient for a 29 month lag of the money supply, but the coefficient is not significant at the 5-percent level.

Within this more general framework, Model 14 includes the interest rate spread variable plus the primary metals and wage rate working series. All of the coefficients are of the hypothesized signs, but four of the ten are not significantly different from zero at the specified level. The equation is interesting because even when traditional inflation regressors are included, the commodity prices retain their independent predictive power.

Diagnostics

In order to check the validity of the models presented in this study, several key models are subjected to a battery diagnostic tests. One of the most crucial issues in producing reliable fore-

TABLE 4

General Information Model Results

Model	**Regressors**						
10.	OIL (0)	OIL (12)	WAGE (9)	REX (14)	AR (1)	MA (9)	SMA (12)
	0.082	-0.063	4.020	-0.050	.0559	0.348	-0.571 *
	4.356	-3.353	3.140	-2.573	8.871	4.766	-6.797 **
	S.E.R. = 0.617		D.F. = 158				
11.	OIL (0)	OIL (12)	FOOD (11)	MACH (0)	MACH (2)	WAGE (9)	
	0.096	-0.102	0.117	0.694	0.649	2.423	
	6.404	-6.708	4.503	4.640	4.294	2.345	
	AR (1)	SMA (12)					
	0.519	-0.866					
	9.646	-10.656					
	S.E.R. = 0.499		D.F. = 162				
12	OIL (0)	OIL (12)	GOLD (24)	GOLD (30)	FOOD (11)	MACH (0)	
	0.092	-0.109	-0.010	-0.010	0.186	0.678	
	5.769	-6.783	-2.256	-2.124	5.046	3.898	
	MACH (2)	WAGE (9)	AR (1)	SMA (12)			
	0.569	2.832	0.446	-0.690			
	3.446	2.477	7.282	-8.182			
	S.E.R. = 0.530		D.F. = 151				
13.	OIL (0)	OIL (5)	FOOD (0)	REX (2)	MONEY (5)	AR (1)	MA (9)
	0.101	0.041	0.093	-0.093	-0.005	0.570	0.191
	3.667	1.854	2.783	-1.509	-0.374	8.756	2.575
	SMA (12)						
	-.0588						
	-6.693						
	S.E.R. = 0.645		D.F. = 143				
14.	OIL (0)	OIL (5)	FOOD (11)	REX (14)	MONEY (29)	AR (1)	MA (9)
	0.084	0.116	0.062	-0.022	0.028	0.671	0.275
	3.354	4.456	2.054	-0.979	1.742	10.584	3.667
	SMA (12)						
	-0.636						
	-6.693						
	S.E.R. = 0.592		D.F. 140				

15.	OIL (0)	OIL (5)	FOOD (11)	METAL (0)	WAGE (9)	REX (14)
	0.105	0.052	0.090	0.140	2.976	-0.032
	5.449	2.634	2.763	1.650	1.736	-1.635
	SPRD (2)	AR (1)	MA (9)	SMA (12)		
	-0.042	0.655	0.225	-0.500		
	-.548	10.900	2.876	-5.390		
	S.E.R. = 0.633		D.F. = 155			

* First row contains parameter estimates.
** Second row contains computed t-statistics.

casts is structural stability. If there is either a structural shift or a significant stochastic component in the parameters, it is difficult to attain a high degree of predictive accuracy.

To this end, Chow tests for structural change were conducted. The time series are split in order to estimate separately for the early (1972-1980) and late (1981-1988) periods. This choice of periods corresponds with the shift from the strong commodity prices of the 1970s to the weaker price environment of the 1980s. None of these F- tests give support for the alternate hypothesis of structural change at the 5 percent significance level. Thus, what may appear to be a structural break is more likely to be indicative of the type of commodity price innovations which may reliably lead changes in the general inflation rate.

Similarly, the evidence of a substantial stochastic component in the parameter values is weak. The most straight-forward means of testing this proposition is to check for the presence of the statistical consequences of random coefficients: heteroscedasticity and serial correlation. The Breusch-Pagan heteroscedasticity test is often recommended in this context. None of the models examined showed significant evidence of heteroscedasticity. As for autocorrelation, the *Q*-statistics at various lag lengths show no signs of autocorrelated residuals.

Another potential complication in time series analysis is the possible presence of autoregressive conditional heteroscedasticity (ARCH) [Engle, 1982.] ARCH processes occur when the magnitude of past disturbances provides information about the magnitude of future disturbances (i.e., non-constant forecast variance). A Lagrange multiplier test is used to detect ARCH. Most models do not show evidence of ARCH processes at the 5 percent significance level. However, a few models have weak signs of ARCH at certain lag lengths. For Model 8, the null hypothesis of no ARCH cannot be rejected for lag lengths of one, three, six, and sixteen months. But at a twelve month lag, the test statistic is marginally significant. Since the evidence of ARCH is scant, the estimation procedures used in the study are not modified.

Despite the fact that many t-statistics become quite low whenever more than five input series are used, multicollinearity between the price series used as predictors does not appear to be a crucial problem. Sample correlation coefficients between various explanatory variables generally remain below 0.5 in absolute value. In order to detect collinearity between larger sets of variables, one explanatory variable is regressed on a set of other explanatory variables in order to calculate the variance inflation factors $[1/(I-R^2)]$. Generally, the variance inflation factors lie between one and three, well below the range considered to indicate a high degree of collinearity.

V. **Conclusion**

The fall and variability in velocity has lead economic policy makers in the U. S. to consider alternative means of determining monetary aggregate targets. Several proposals that have garnered attention from the popular press seek to utilize commodity prices as leading indicators of inflation and exchange rate movements. This study employs transfer ARIMA analysis to investigate some of the empirical problems associated with identifying individual price variables that may provide reliable signals to policy analysts regarding future movements in inflation rates.

Similar to other papers using vector autoregression and polynomial distributed lag estimation techniques, the authors' estimation results imply that examining the linkages between commodity prices and aggregate price measures can advance the understanding of short term inflationary trends in the U. S. The results indicate the existence of a wide variety of dynamic relationships. The models satisfy the general criteria used to analyze the empirical characteristics of transfer ARIMA models (low Q statistics, low standard errors of regression relative to the univariate CPI equation, etc.) [Jenkins, 1979, pp. 52-5].

It also appears that the effects of the initial movements in some of the explanatory series may be shorter than anticipated by the policy proponents. Angell points out that the turning points of a moving average commodity index has led turning points in the inflation rate by an average of seven and one half months [Angell, 1987, pp. 3-4]. The results of this study indicate that the lag lengths in the relationships between inflation and individual commodity prices are considerably more variable and do not cluster tightly around the seven and one-half month average lag found when using turning points of the general indices.

In general, two results of the modeling effort stand out. First, numerous commodity prices exhibit consistent relationships with the CPI. Although not all commodity price indicators provide as much advance warning as the Angell and Johnson proposals imply, the breadth of these relationships indicates that this is not an insurmountable problem. It should also be noted that the CPI is published with a lag and subject to revision. Thus, even commodities appearing with short lags in the equations may be useful for policy purposes.

Second, commodity prices are incrementally useful predictors of inflation. This result holds true even when financial variables and indicators such as wages and prices of intermediate goods are included in the transfer functions. There appears to be, therefore, a complementary role for commodity prices in modeling changes in the CPI. The supplementary nature of the role of commodity indicators is important to bear in mind. Rigidly tying policy decisions to fluctuations on commodity markets may not only be inappropriate, but also may change the established relationships necessary for such an approach to be useful.

In order to provide wide coverage of economic sectors, petroleum, food products, and precious and non-precious metals should be considered. Since the relevance of commodity prices lies in the presumption that commodity markets may provide sensitive indicators of current inflationary pressures, commodities appearing with very long lags may not be highly relevant for short and intermediate-term policy purposes. The results also indicate that intermediate inputs, such as machinery and chemicals, may also help policymakers formulate a reliable, broadly-based model to supplement more traditional indicators.

Although this paper focuses on U. S. domestic inflation, monetary authorities in other nations may also find the information derived from commodity markets useful in a purely domestic context. The Baker exchange rate proposal moves this approach into an international setting. By requiring the *Group of Seven* industrial nations to monitor the same bundle of commodity prices as an indicator of global inflationary trends, coordination of monetary and exchange rate policies would be facilitated. Similarly, U. S. Federal Reserve Chairman Greenspan emphasizes the potentially differing policy implications when commodity prices rise in terms of only one currency as opposed to rising in terms of all currencies [Greenspan, 1987, p. 4].

W. D. Angell, "A Commodity Price Guide to Monetary Aggregate Targeting," Paper presented to the Lehrman Institute, New York, December 10, 1987.

R. T. Baillie, "Commodity Prices and Aggregate Inflation: Would a Commodity Price Rule be Worthwhile?," Paper presented at Carnegie Rochester Public Policy Conference, Pittsburgh, November 18, 1988.

J. M. Boughton and W. Branson, "Commodity Prices as a Leading Indicator of Inflation," IMF Working Paper WP/88/ 87, International Monetary Fund, 1988.

G. E. P. Box, and G. M. Jenkins, *Time Series Analysis: Forecasting and Control,* San Francisco: Holden-Day, 1976.

C. Chatfield, *The Analysis of Time Series,* New York: Chapman and Hall, 1984.

R. F. Engle, "Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation," *Econometrica,* 50, 4, July, 1982, pp. 987- 1007.

C. A. Gamer, "Commodity Prices: Policy Target or Information Variable?," Research Working Paper RWP 88-10, Federal Reserve Bank of Kansas City, 1988.

C. W. J. Granger, and P. Newbold, "Spurious Regressions in Econometrics," *Journal of Econometrics,* 2, 2, July, 1974, pp.111-20.

A. Greenspan, Testimony presented to Subcommittees on Domestic Monetary Policy and on International Finance, Trade, and Monetary Policy, *U.* S. House of Representatives, December 18,1987.

L. D. Haugh, and D. A. Pierce, "Causality in Temporal Systems: Characterizations and a Survey," *Journal of Econometrics,* 5, 3, August 1977, pp. 265-93.

G. M. Jenkins, *Practical Experiences with Modeling and Forecasting Time Series,* Kendal, Great Britain: Titus Wilson, 1979.

P. T.Kilbom, "Accord Indicated on Stable Dollar," *The New York Times,* April 8, 1988.

P. D. Koch, J. A. Rosensweig, and J. A. Whitt, "The Dynamic Relationship between the Dollar and U. S. Prices: An Intensive Empirical Investigation," *Journal of International Money and Finance*, 7, 2, March 1988 pp.181-204.

M. McNamee, "The Mellowing of a Supply Side Warrior," *Business Week*, April 25, 1988.

A. Pankratz, *Forecasting with Univariate Box Jenkins Models*, New York: John Wiley and Sons, 1983.

G. W. Schwert, "Tests of Causality: The Message in the Innovations," Working Paper 77-4, Graduate School of Man- agement, University of Rochester, 1977.

R. H. Webb, "Commodity Prices as Predictors of Aggregate Price Change," Federal Reserve Bank of Richmond *Economic Review*, 74, 6, November/December 1988, pp. 3-11.

J. A. Whitt, "Commodity Prices and Monetary Policy:' Working Paper 88-8, Federal Reserve Bank of Atlanta, 1988.

A. Zellner, and F. Palm, "Time Series Analysis and Simultaneous Equation Econometric Models," *Journal of Econometrics,* 2, 1, April 1974, pp. 17-54.

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Volume 19, Pages 1-14.