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Commodity prices and the CPI: Cointegration, information, and signal extraction

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Abstract: This paper provides theoretical underpinnings for the commodity price/aggregate price relationship, discusses the conditions under which commodity prices are useful information variables for monetary policy, and provides empirical results which suggest why commodity prices have not been very useful for forecasting.

Keywords: Commodity price indices, Inflation forecasts, Cointegration.

1. Introduction

Recently there has been much interest expressed by policy practitioners [Angell (1987); Whitt (1988); DeFina (1988); Garner (1989); Furlong (1989)], academics [Boughton and Branson (1988); Hall (1982); Baillie (1989)] and private sector financial analysts [Brittan (1988); Kudlow (1988)] in the use of commodity price indices as indicators of monetary policy. Much of this interest arises from the observation that commodity price indices appear to lead the consumer price index (CPI), and so may signal inflationary expectations and therefore be useful in producing better forecasts of future inflation. Intuitively, changes in inflationary expectations should appear as changes in commodity prices prior to being reflected in changes in the CPI because commodity markets are speculative auction markets in which prices immediately adjust to changes in information, while many prices comprising the CPI respond to inflationary expectations with a lag. If this is the case, a signal of inflationary expectations emanating from the commodity markets could provide the monetary authority with information which can be used to reduce price level volatility.

While an apparent correlation between commodity price indices and the CPI seems plausible, little empirical evidence supports the contention that the information in commodity prices can be used effectively to forecast the CPI or to adjust monetary policy. For example, Baillie (1989), Boughton and Branson (1988), and Garner (1989) find that there is no discernible long-run relationship between commodity prices and the CPI (the two time series are not cointegrated), yet Boughton and Branson (1988) and Garner (1989) find a weak short-run relationship (commodity prices help predict future consumer price inflation) while Baillie (1989) does not. Whitt (1988) suggests that there is not only a long-run relationship, but that commodity prices should be an intermediate target of monetary policy. This suggestion is contested by Garner (1989), and by Kitchen, Conway, and LeBlanc (1990) who claim that monetary policy directed at controlling com-

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modity prices could cause real economic disruption in agricultural markets, although they find no linkage between commodity prices and monetary policy that could be exploited. DeFina's (1988) review of the literature leads him to the same conclusion.

A motivation for this paper is to provide a theoretical foundation for long- and short-run relationships between commodity and consumer prices, and to then use the theoretical results to suggest a unifying interpretation of the disparate empirical findings. The paper develops a simple two-sector rational expectations model in which commodity prices may be informative about the CPI, defined as the price of manufactures. The model follows Fischer (1977) and Gray (1976) in that wages in the manufacturing sector are determined by long-term contracts while commodity prices are determined in speculative auction markets.

The major findings of this analysis are as follows. First, if the CPI is nonstationary, then a likely source of this nonstationarity is the monetary policy regime. This result is consistent with Pecchenino and Rasche (1990) and Goodfriend (1987). Second, while the model suggests that there is a close relationship between commodity and aggregate prices, commodity and aggregate prices will be cointegrated only under certain assumptions on the money supply rule or on the stochastic processes generating the demand and supply shocks to the system. Various empirical tests provide weak support for the hypothesis that commodity prices and the CPI are cointegrated. Third, even if a long-run relationship between consumer and commodity prices does not exist, the information in commodity prices may be useful for forecasting future consumer prices. However, signal extraction problems mitigate against using commodity prices to set monetary policy.

2. The model

The model is an adaptation of the Fischer–Gray model of long-term contracts. It is devised to capture linkages between commodity prices and the CPI by including a commodity sector as well as a manufacturing sector. A well-known feature of the Fischer–Gray model is that monetary policy can be effective in smoothing out unforeseen shocks to the system. This feature is maintained in this extension even if consumer prices are observed with a lag so long as the monetary authority can respond to the information in commodity prices before private sector agents can renegotiate their long-term contracts.

Consider the following two sector economy. In sector 1 manufactured goods are produced by firms using a Cobb–Douglas technology with two factors of production, labor and commodities. In sector 2 commodities are produced using a fixed factor (which is subsumed in the analysis). Labor contracts set the nominal wage one period in advance, and so labor at time t is demand determined. Commodity markets are assumed to clear in each period. Consumer demand is represented by a quantity equation, and, as consumers demand manufactures only, the price of manufactures represents the CPI. The monetary authority sets the money supply according to a deterministic rule which depends only on publicly available information. Both private sector individuals and the monetary authority have the same information set at any time t, and agents are assumed to form rational expectations.

2.1. The manufacturing sector

The Cobb–Douglas production function for the representative profit-maximizing firm is

$$y_t = \alpha \ell_t^d + (1-\alpha) c_t^d + \mu_t, \tag{1}$$

where $\alpha \in (0, 1)$, y_t is the logarithm of output of manufactures at time t, ℓ_t^d is the logarithm of labor demand, c_t^d is the logarithm of demand for commodities, and μ_t is a random variable.

The representative firm demands commodities until the marginal product of commodities equals the real price of commodities,

$$g_t - p_t = \ln(1-\alpha) - \alpha c_t^d + \alpha \ell_t^d, \tag{2}$$

where g_t is the logarithm of the nominal price of commodities, and p_t is the logarithm of the price of manufactures (the price level). It demands labor until labor's marginal product equals the real wage,

$$w_t - p_t = \ln \alpha + (1-\alpha) c_t^d - (1-\alpha) \ell_t^d, \tag{3}$$

where w_t is the logarithm of the wage. Assume that the logarithm of labor supply is an increasing function of the real wage,

$$l_t^s = \theta(w_t - p_t), \tag{4}$$

where $\theta > 0$.

Nominal wages are set at time $t-1$ so as to equate expected supply and demand for labor at time t. Thus, since agents from rational expectations

$$w_t = \frac{\ln \alpha}{1+\theta(1-\alpha)} + \frac{(1-\alpha)_{t-1}c_t^d}{1+\theta(1-\alpha)} + {}_{t-1}p_t, \tag{5}$$

where ${}_{t-1}x_t$ is the rational expectation of some variable x_t taken at time $t-1$. Substituting (5) into (3) yields time t demand for labor

$$\ell_t^d = \frac{\theta \ln \alpha}{1+\theta(1-\alpha)} + \frac{1}{1-\alpha}(p_t - {}_{t-1}p_t) + c_t^d - \frac{{}_{t-1}c_t^d}{1+\theta(1-\alpha)}, \tag{6}$$

which determines manufacturing output

$$y_t = \frac{\alpha\theta \ln \alpha}{1+\theta(1-\alpha)} + \frac{\alpha}{1-\alpha}(p_t - {}_{t-1}p_t) + c_t - \frac{\alpha_{t-1}c_t^d}{1+\theta(1-\alpha)} + \mu_t. \tag{7}$$

2.2. The commodities sector

Assume that the logarithm of commodity supply c_t^s depends on the logarithm of the real commodity price and a real shock η_t (such as the weather):

$$c_t^s = \beta(g_t - p_t) + \eta_t, \tag{8}$$

where $\beta > 0$, and η_t is a random variable. Combining eqs. (8), (6) and (2), commodity market clearing implies that

$$c_t = \beta \ln(1-\alpha) + \frac{\beta\alpha\theta \ln \alpha}{1+\theta(1-\alpha)} + \frac{\alpha\beta}{1-\alpha}(p_t - {}_{t-1}p_t) - \frac{\alpha\beta_{t-1}c_t^d}{1+\theta(1-\alpha)} + \eta_t. \tag{9}$$

2.3. The monetary sector

The model is closed with money demand and supply equations. Assume money demand is represented by the quantity equation

$$m_t^d - p_t = y_t + \epsilon_t, \tag{10}$$

where ϵ_t is a random variable, and m_t^d is the logarithm of nominal money demand. Finally, assume that monetary authority follows a deterministic rule which can depend on all information available at the beginning of period t, $\mathscr{I}_t$,

$$m_t^s = m(\mathscr{I}_t), \tag{11}$$

where $\mathscr{I}_t$ may contain information that was unavailable to private sector agents when they set their wages such as the market clearing commodity price.

3. Time series properties of the price level

Much of the empirical literature on the relationship between commodity price indices and the consumer price index conclude that commodity prices are integrated of order one and the CPI is integrated of order two [see Boughton and Branson (1988); Baillie (1989); among others]. Further, these same studies find that there is no long-run relationship between commodity price indices and the first differenced CPI series (the series are not cointegrated). [1] But, short-run relationships which may be useful for monetary policy might still exist [see Boughton and Branson (1988); and Garner (1989)].

The price series generated by the theoretical model of Section 2 will exhibit the time series behavior suggested by the empirical work under certain assumptions on the stochastic properties

[1] In the simplest terms, if a random variable follows a random walk, then its time series is said to be nonstationary in levels, or integrated of order j, $j \geq 1$ [$I(j)$]. A random variable that exhibits mean reversion is said to be stationary or integrated of order zero [$I(0)$]. Two $I(j)$ time series are said to be cointegrated if there exists a linear combination of the series that is $I(j-1)$. If two time series are cointegrated then there exists a long-run relationship that can be exploited for forecasting purposes. If two series are not cointegrated, there may still be contemporaneous correlations between realizations of the variables that can be used for one-step ahead forecasting.

of the error terms and/or on the money supply rule. To see this, solve the model for the price level:

$$\begin{aligned}
p_t &= K_1 + m_t - \eta_t - \mu_t - \epsilon_t \\
&\quad + \frac{\alpha(\beta+1)}{1+\theta(1-\alpha)+\alpha\beta}\,{}_{t-1}\eta_t \\
&\quad - \frac{\alpha(\beta+1)}{(1-\alpha)}(p_t - {}_{t-1}p_t) \\
&= K_1 + m_t - \eta_t - \mu_t - \epsilon_t \\
&\quad + \frac{\alpha(\beta+1)}{1+\theta(1-\alpha)+\alpha\beta}\,{}_{t-1}\eta_t \\
&\quad - \frac{\alpha(\beta+1)}{1+\alpha\beta}\big[(m_t - {}_{t-1}m_t) + (\eta_t - {}_{t-1}\eta_t) \\
&\qquad - (\mu_t - {}_{t-1}\mu_t) - (\epsilon_t - {}_{t-1}\epsilon_t)\big], \qquad (12)
\end{aligned}$$

where

$$K_1 = \frac{\beta \ln(1-\alpha)^2(1+\theta)}{1+\theta(1-\alpha)+\alpha\beta} + \frac{(1+\beta)\alpha\theta \ln\alpha}{1+\theta(1+\alpha)+\alpha\beta}.$$

The expectational error terms $[(x_t - {}_{t-1}x_t)$ for some variable $x]$ in eq. (12) are all $I(0)$ processes regardless of the underlying generating process for the individual variables. Thus, the order of integration of the price level depends on the processes generating the error terms, η_t, μ_t and ϵ_t, and/or the money supply, m_t.

Consider first the error terms: μ_t and η_t are real shocks and ϵ_t is a nominal shock. Suppose that the shock to output, μ_t, is a technological shock and is $I(1)$ [2]

$$\mu_t = \mu_{t-1} + v_t,$$

where v_t is a white noise process, and that the demand shock, ϵ_t is a white noise process. Further assume that the commodity supply shock, η_t, is $I(0)$, and represents, perhaps, the weather. If the money supply rule is $I(1)$, then the price level will be either $I(0)$, if m_t and μ_t are cointegrated, or $I(1)$ if they are not. If the price level process is $I(2)$, then it must the case that this money supply rule is $I(d)$, $d > 1$. [3] Thus, barring peculiar assumptions on the error processes, it appears that if the macroeconomy can be characterized by a simple rational expectations model, then the time series behavior of the CPI can be traced to the money supply process. This finding is consistent with both Goodfriend (1987) and Pecchenino and Rasche (1990).

Now solve the model for commodity prices:

$$\begin{aligned}
g_t &= p_t + K_2 + \frac{\alpha}{1-\alpha}(p_t - {}_{t-1}p_t) \\
&\quad - \frac{\alpha}{1+\theta(1-\alpha)+\alpha\beta}\,{}_{t-1}\eta_t \\
&= p_t + K_2 - \frac{\alpha}{1+\theta(1-\alpha)+\alpha\beta}\,{}_{t-1}\eta_t \\
&\quad + \frac{\alpha}{1+\alpha\beta}\big[(m_t - {}_{t-1}m_t) - (\eta_t - {}_{t-1}\eta_t) \\
&\qquad - (\mu_t - {}_{t-1}\mu_t) - (\epsilon_t - {}_{t-1}\epsilon_t)\big], \qquad (13)
\end{aligned}$$

where

$$K_2 = \frac{[1+\theta(1-\alpha)]\ln(1-\alpha) - \alpha^2\beta\theta\ln\alpha}{1+\theta(1-\alpha)+\alpha\beta}.$$

If ${}_{t-1}\eta_t$ equals zero (η_t is an $I(0)$ process), then g_t and p_t are cointegrated since the expectational error terms are all white noise. Even if g_t and p_t are not cointegrated, g_t contains information about p_t that could be valuable for predicting future movements of the price level and for setting monetary policy. [4]

To determine empirically whether or not there exists a long-run relationship between the CPI and commodity prices, a three step procedure is followed. First, using Phillips–Perron (1988) unit root tests, the order of integration of the logarithm of the CPI and the logarithms of the Commodity Research Board's All Commodities [CRBAll], Raw Materials [CRBRaw], and Foodstuffs [CRBFst] indexes for the period 1953.1–1989.7 is determined. [5] As is shown in Table 1, the test statistics are generally consistent with the

[2] The assumption that μ_t is $I(1)$ is consistent with the findings of the real business cycle literature. See, for example, Nelson and Plosser (1982).

[3] See Granger (1986) and Hendry (1986) for discussions of cointegration.

[4] These cointegration/information results carry over to a more complicated model in which commodity supply depends on inflationary expectations as well as the real commodity price. Further details are available from the author on request.

[5] There are many tests to determine the order of integration of a time series. Many are discussed and explained in Baillie (1989).

Table 1
Phillips–Perron unit root tests of monthly price indices [a]

	$t_{\tilde{\alpha}}$		t_{α^*}	
	$\ell = 4$	$\ell = 12$	$\ell = 4$	$\ell = 12$
A. 1953.1–1989.7 ($T = 439$)				
CPI	−2.800	−2.250	4.096 [b]	2.639 [b]
CRBAll	−1.851	−2.111	−0.160	−0.409
CRBRaw	−2.072	−2.385	−0.024	−0.314
CRBFst	−1.845	−1.952	−0.788	−0.861
B. 1953.1–1965.12 ($T = 156$)				
CPI	−1.880	−2.057	0.285 [b]	0.181 [b]
CRBAll	−0.601 [b]	−0.312 [b]	−1.320	−1.222
CRBRaw	−1.523	−1.848	−1.323	−1.679
CRBFst	−2.220	−2.242	−1.749	−1.766
C. 1966.1–1979.9 ($T = 165$)				
CPI	−0.184	−0.593	4.051 [b]	2.782 [b]
CRBAll	−2.148	−2.395	0.335 [b]	−0.033
CRBRaw	−2.335	−2.570	0.511 [b]	0.099 [b]
CRBFst	−1.996	−2.109	−0.393	−0.485
D. 1982.1–1989.7 ($T = 91$)				
CPI	−1.789	−1.470	0.208 [b]	0.287 [b]
CRBAll	−1.285	−1.617	−1.123	−1.496
CRBRaw	−1.413	−1.691	−0.506 [b]	−0.905
CRBFst	−1.729	−1.832	−1.384	−1.479

[a] There are many different versions of the Phillips–Perron test. Two are used here. The first assumes that the price series can be represented by the equation:

$$x_t = \tilde{\mu} + \tilde{\beta}(t - n/2) + \tilde{\alpha} x_{t-1} + \tilde{u}_t,$$

where $\tilde{u}_t$ is a stationary process and n is the sample size. If the statistic $Z(t_{\tilde{\alpha}})$ is less than -3.41 the unit root hypothesis can be rejected against a stationary, $I(0)$, alternative. More rarely, if $Z(t_{\tilde{\alpha}})$ exceeds -0.94 the unit root hypothesis can be rejected against an explosive alternative. Both critical values are at the 0.05 level and the $Z(t_{\tilde{\alpha}})$ statistic is an adjusted version of the regualr t statistic. The second version of the test assumes, instead, that the price series can be represented by the equation

$$x_t = \zeta^* + \alpha^* x_{t-1} + u_t^*,$$

where u_t^* is a stationary process. If the statistic $Z(t_{\alpha^*})$ is less than -2.86 the unit root hypothesis can be rejected against a stationary, $I(0)$, alternative. More rarely, if $Z(t_{\alpha^*})$ exceeds -0.7 the unit root hypothesis can be rejected against an explosive alternative. Both critical values are at the 0.05 level and the $Z(t_{\alpha^*})$ statistic is an adjusted version of the regular t statistic. See Perron (1988) for an exact description of the statistics.

[b] Reject the hypothesis of a unit root at 5% level of significance.

hypothesis that the CPI series is $I(1)$. In some cases there is slight evidence that the unit root hypothesis can be rejected against an explosive rather than a stationary alternative. This unusual result is most likely due to the apparent regime changes that have taken place in the CPI series since the Federal Reserve did not follow a single money supply rule over this period. Pecchenino and Rasche (1990) identify four stable money supply regimes for the data period and show that the hypothesis that the CPI is $I(1)$ in three of the four subperiods cannot be rejected. Table 1 reports the results of Phillips–Perron unit root tests for the periods they identify (the 'free reserve' operating procedure, 1953.1–1965.12, the Treasury bill rate/federal funds rate operating procedure, 1966.1–1979.9, the 'New Operating Procedure', 1979.10–1981.12, and the 'nonborrowed reserves' operating procedure, 1982.1–1989.7). For each of these subperiods there is stronger support for the hypothesis that the CPI is $I(1)$. Further, the hypothesis that the commmodity price indices are all $I(1)$ in each of the subperiods cannot be rejected. [6]

The final step in the procedure is to test whether or not the CPI and the commodity indices are cointegrated over the subperiods. Three tests were conducted. First, since the theory predicts that if the CPI and commodity prices are cointegrated then the cointegrating vector is $(1, -1)$, real commodity price series were constructed. If these series are stationary, then the hypothesis that the CPI and commodity prices are cointegrated with cointegrating vector $(1, -1)$ can be accepted. To test for cointegration, the hypothesis that each series has a unit root was tested using the Phillips–Perron test and the Kwiatkowski–Phillips–Schmidt Stationarity test (1990). The null hypothesis of the latter test is that each of the time series is stationary. As is shown in Table 2, the Phillips–Perron test soundly rejects the hypothesis of cointegration, while the Kwiatkowski–Phillips–Schmidt test gives weak support to the hypothesis that the CPI and commodity prices are cointegrated in the middle and

[6] Since the 'New Operating Procedure' period was so short, this subperiod is ignored in the analysis.

Table 2
Unit root tests of monthly real commodity price indices

	Phillips–Perron				Kwiatkowski–Phillips–Schmidt [a]			
	t_{α^*}		$t_{\tilde{\alpha}}$		η_μ		η_τ	
	$\ell=4$	$\ell=12$	$\ell=4$	$\ell=12$	$\ell=4$	$\ell=12$	$\ell=4$	$\ell=12$
A. 1953.1–1965.12 ($T=156$)								
RCRBAll	−1.430	−1.424	0.694 [b]	−0.573 [b]	2.169	1.109	3.480	0.163
RCRBRaw	−1.792	−2.025	−1.514	−1.852	1.226	0.534	0.210	0.009 [c]
RCRBFst	−1.254	−1.243	−2.267	−2.215	2.782	1.148	0.213	0.116 [c]
B. 1966.1–1979.9 ($T=165$)								
RCRBAll	−1.313	−1.665	−2.029	−2.329	1.312	0.556	0.245	0.109 [c]
RCRBRaw	−1.477	−1.848	−2.329	−2.606	1.176	0.552	0.209	0.099 [c]
RCRBFst	−1.517	−1.620	−1.830	−1.945	1.300	0.545	0.271	0.119 [c]
C. 1982.1–1989.7 ($T=91$)								
RCRBAll	−1.558	−1.760	−1.372	−1.694	0.878	0.389 [c]	0.237	0.112 [c]
RCRBRaw	−1.701	−1.977	−1.534	−1.792	0.147	0.310 [c]	0.281	0.134 [c]
RCRBFst	−0.721	−0.784	−1.771	−1.862	1.531	0.652	0.194	0.094 [c]

[a] There are two versions of the Kwiatkowski–Phillips–Schmidt test. For the first test, the null hypothesis that the time series represented by

$$x_t = r_t + \epsilon_t$$

$$r_t = r_{t-1} + u_t,\ u_t \sim \text{i.i.d.}(0, \sigma_u^2)$$

is stationary in levels cannot be rejected if the statistic, $\eta_\mu < 0.463$. For the second test, the null hypothesis that the time series represented by

$$x_t = \xi t + r_t + \epsilon_t$$

$$r_t = r_{t-1} + u_t,\ u_t \sim \text{i.i.d.}(0, \sigma_u^2)$$

is trend stationary cannot be rejected if $\eta_\tau < 0.146$. See Kwiatkowski, Phillips, and Schmidt (1990) for an exact description of the statistics.

[b] Reject the hypothesis of a unit root at 5% level of significance.

[c] Cannot reject the null hypothesis of stationarity at 5% level of significance.

final periods. Finally, the Johansen cointegration test [Johansen (1988); Johansen and Juselius (1989)] was conducted. This test does not constrain the cointegrating vector to be (1, −1). Table 3 reports the results. No cointegrating relationships are found for the first and third periods, while the CPI appears to be cointegrated with each of the commodity price series during the

Table 3
Johansen trace tests [a]

	53.1–65.12 ($T=152$)		66.1–79.9 ($T=161$)		82.1–89.7 ($T=87$)	
	$r=0$	$r\le 1$	$r=0$	$r\le 1$	$r=0$	$r\le 1$
CPI+CRBAll	1.890	0.665	22.412	2.503	4.995	0.073
CPI+CRBRaw	6.391	0.691	23.124	5.113	6.946	0.154
CPI+CRBFst	14.326	0.283	20.865	1.834	3.838	0.838

[a] The parameter r represents the number of cointegrating factors. If the trace statistic for $r=0$ exceeds 18, then the hypothesis that the series are cointegrated cannot be rejected. See Johansen and Juselius (1989) for an exact description of the statistic.

second period. There appears to be, at best, weak support for cointegration of the CPI and commodity prices.

4. Policy

Interest in commodity prices as indicators of monetary policy was rekindled by Federal Reserve Board Governor Wayne Angell (1987). He suggested that the apparent relationship between commodity price indices and the CPI indicated by their time series [7] (commodity price series lead the CPI) might, even in the absence of cointegration, provide useful information for monetary policy directed at controlling price level variance. Others [see, for example Whitt (1988)] have suggested that monetary policy could be directed at controlling commodity price level variance arguing that CPI variance would be reduced as a result.

Whitt's suggestion of monetary policy controlling commodity prices cannot be supported by this model since the monetary authority does not have an information advantage over agents operating in the commodities market. However, if, following Angell, the monetary authority's goal is to minimize the variance of the CPI, then it has one potential source of information upon which to base its policy: the innovation in commodity prices $(g_t - {}_{t-1}g_t)$. With this information it can improve its forecast of the price level at time t, and, potentially, choose policy parameters to minimize the variance of the price level forecast.

Under the assumption that p_t and g_t are not cointegrated, let μ_t and η_t be $I(1)$ where

$$\mu_t = \mu_{t-1} + v_t,$$

$$\eta_t = \eta_{t-1} + h_t,$$

where v_t and h_t are mean zero white noise processes with finite variances σ_v^2 and σ_h^2, respectively. Further let ϵ_t be a mean zero white noise process with finite variance σ_ϵ^2. The shocks v_t, h_t, and ϵ_t may be contemporaneously correlated. At time t the monetary authority and individual agents' information sets are

$$\mathscr{I}(t) = \{\mu_s, \eta_s, v_s, h_s, \epsilon_s, \Sigma, \Omega, m_s^s, \forall s < t, \text{ and } g_t)$$

where Ω is the contemporaneous correlation matrix, Σ is the variance/covaraince matrix, and

$$m_t^s = \delta(g_t - {}_{t-1}g_t). \tag{14}$$

Combining (12) and (13) with the monetary authority's money supply rule, (14), yields the forecast of the price level

$$p_t = \frac{\alpha - 1}{1 - \delta + \alpha\beta}(h_t + v_t + \epsilon_t) + \frac{\delta - \alpha(1+\beta)}{1 - \delta - \alpha\beta}(\eta_{t-1} + \mu_{t-1}) + K_1. \tag{15}$$

This forecast, under the assumptions of the model, is exactly the price level. The price level (forecast) variance is defined by

$$\operatorname{var} p_t = \left[\frac{\alpha - 1}{1 - \delta + \alpha\beta}\right]^2 \left[\sigma_h^2 + \sigma_v^2 + \sigma_\epsilon^2\right] + \left[\frac{2(\alpha - 1)}{1 - \delta + \alpha\beta}\right]\left[\operatorname{corr}(h_t, v_t) + \operatorname{corr}(h_t, \epsilon_t) + \operatorname{corr}(v_t, \epsilon_t)\right]. \tag{16}$$

Since η_{t-1} and μ_{t-1} are known at time t, they have zero variances. Further, since they were known when wages were set at time $t-1$, these values do not provide the monetary authority with additional information upon which to base its policy.

The choice of δ that minimizes the variance of the price level forecast is negative infinity: the authority fully offsets the exogenous shocks, both nominal and real. Generally, the monetary authority only wants to respond to nominal shocks, thereby not moving the economy away from the path determined by real factors. This may not be possible given the signal extraction problem the monetary authority faces: it observes the sum of the shocks not their individual realizations. Any policy action taken by the authority would, in this situation, have both benefits and costs. If the nominal shock is the primary contributor to the variance of the price level and the authority evaluates the cost of offsetting real as well as nominal

[7] Visual inspection of commodity price series superimposed over the CPI series indicates that commodity prices lead the CPI.

shocks as being low, then the authority could follow the policy procedure suggested. Otherwise, if the economic costs are high or if the main contributors to price level variance are real shocks, the authority should be more circumspect in its policy choice.

5. Conclusion

The apparent relationship between the CPI and commodity prices has proved to be a seductive lure to policy-makers and academics looking for a variable useful in the conduct of monetary policy. Previous studies have tended to ignore the components of the CPI and the commodity price relationship while focusing on the aggregate relationship. This paper provides a theoretical foundation for short and long-run relationships between the CPI and commodity prices, and examines the components of these relationships. This examination reveals why the aggregate relationships may not generally be useful from a policy or forecasting perspective.

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