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Abstract

This paper estimates a structural VAR model of U.S. consumer and world commodity prices. An equiproportional long-run response of nominal price levels to a monetary shock yields identifying restrictions. Exogenous innovations to monetary policy account for a sizable share of the co-movement of these series, including during episodes more commonly attributed to “supply shocks.”

JEL Categories: C32, E31, E49.

Keywords: Commodity price determination, vector autoregression, long-run restrictions, co-integration, monetary shocks.

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1 Introduction

Commodity prices are widely thought of as a potential leading indicator of general inflation, especially in the financial press, despite a mixed forecasting record.¹ Their use in empirical macroeconomics as “indicators of future inflation” has become fairly standard in vector autoregressive models of monetary policy since Sims (1992).² More recently, commodity prices have appeared in the instrument set of estimated forward-looking monetary reaction functions.³

In many of these cases, commodity prices are treated conceptually as an exogenous factor — often labeled a “cost-push” shock — or at least predetermined with respect to monetary policy. Moreover, most of this work examines the short-run relationship between consumer and commodity prices. In contrast, this paper focuses on the long-run co-movement between consumer and commodity prices. Consistent with almost all modern macroeconomic theories, my approach assumes that purely nominal (i.e. monetary) shocks eventually will feed equiproportionally into all prices, leaving relative prices unaffected. It matches several properties of the data illustrated in section 2 while remaining agnostic with respect to the short-run sources of commodity price volatility or aggregate price sluggishness. Other shocks are left unrestricted in the model, and may have permanent or temporary effects on relative prices.

As noted in section 3, a parallel may be found between the treatment of commodity prices here and the treatment of exchange rates in the monetary models of Dornbusch (1976) and Mussa (1982). Extensions of the overshooting model specifically to commodity prices have been undertaken by Frankel (1986) and Boughton and Branson (1991). This paper formally tests the long-run implications of these models of commodity prices.⁴ Section 4 reports that monetary shocks actually account for a sizable portion of the variability in commodity prices — notably in periods traditionally associated with “supply shocks” — and that like exchange rates, commodity prices overshoot with respect to monetary shocks. These dynamics arise from the data and are not imposed by the identifying assumptions. The implication is that commodity prices more likely are an indicator of inflationary pressures originating with monetary policy rather than a factor contributing directly to aggregate price inflation.

¹For a discussion, see, for example, Blomberg and Harris (1995) or Furlong and Ingenito (1996).

²See also Christiano et al. (1999). Hanson (2004) offers a critique of this practice.

³See Clarida et al. (2000), for example.

⁴Using techniques similar to those in section 3, Lastrapes (1992), Clarida and Galí (1994), and Rogers (1999), amongst others, have estimated the importance of nominal (monetary) shocks for exchange rates.

2 Statistical Properties of the Data

The aggregate price measure used in this study is the U.S. consumer price index (CPI) less shelter for all urban consumers.⁵ A broad non-fuels, non-precious metals index from the International Monetary Fund is used to measure commodity prices. Figure 1 plots the logarithmic levels of each series, with the implied real commodity price (log IMF index less log CPI) in the lower panel. Over the 1959 – 2002 sample period, the real commodity price trends downward, interrupted by a few sizable jumps. Table 1 summarizes the average level and volatility of the growth rate of the CPI and the IMF index. Commodity prices are far more volatile than consumer prices, although the IMF overall commodity price index used here generally is less volatile than each of its component commodity prices, as the idiosyncratic shocks that impart variability to these components occasionally will offset each other in the overall index.

Table 1 also reports moments for three sub-periods. The 1972 break-point is motivated by the collapse of Bretton Woods and the poor worldwide harvests that occurred late that year and into 1973. In the 12-year interval through the end of 1983 (suggested by the conclusion of the Volcker disinflation in the U.S.), the average level and volatility of both the price series are much higher than the sub-samples that precede or follow. Although the commodity price growth rate exceeded consumer price inflation during the 1972 – 1983 interval, the converse was true throughout the rest of the sample.

While the contemporaneous correlation between consumer price inflation and the growth rate of commodity prices is usually quite low, the short-run relationship between the price series more commonly is investigated by testing for Granger causality. The null hypothesis that, conditional on lagged CPI values, lagged commodity prices do not provide any incremental forecasting power for the CPI is rejected in table 2. This result holds both in levels and in first differences, while the converse does not: consumer prices show little incremental forecasting power for commodity prices.⁶

To investigate the long-run relationship between consumer and commodity prices, first the stationarity of each series is examined. For two series to share a common stochastic trend (i.e. to be cointegrated), both individually must be integrated of order one (or higher). Consistent with the upward trend in each series, table 3 reveals that the null of a unit root cannot be rejected for the log level of either series, while that null is strongly rejected in the first difference of commodity prices and is weakly rejected for

⁵This measure is used to avoid problems with the accounting of mortgage interest prior to 1982. The GDP deflator and the all-items CPI index (not reported here) yield similar results.

⁶All tests use 4 lags of data. Repeating the tests in tables 2 through 4 with 3, 5, or 6 lags returned similar conclusions.

CPI inflation.⁷

Nonetheless, the results in table 4 do not indicate cointegration between consumer and commodity prices. The Phillips and Ouliaris (1990) test posits an unrestricted equilibrium relationship and tests whether the residual is stationary, as cointegration requires. The Horvath and Watson (1995) test imposes a specific cointegrating vector under the alternative; here the relevant cointegrating vector is $[1, -1]$: the log difference of the two price series should be stationary. At conventional levels of significance, the null hypothesis of no cointegration cannot be rejected for either test.

3 Model and Estimation

Section 2 established several stylized facts about the properties of consumer and commodity prices with which the model developed in this section must be consistent. The cointegration results in table 4 suggest that other shocks may have permanent effects on the relative price of commodities; the model below allows for permanent effects of both real and nominal shocks on both consumer and commodity prices.⁸ The sole imposed restriction is that nominal shocks cannot have a permanent effect upon real commodity prices, as nominal shocks are by definition neutral for any real variable in the long-run.

This identifying restriction can be implemented in a bivariate model of consumer and commodity prices using the approach of Blanchard and Quah (1989) and Shapiro and Watson (1988). The vector moving average (VMA) representation of this structural model can be written as:

$$\Delta c_t - \Delta p_t = A_{11}(L)\psi_t + A_{12}(L)v_t \quad (1)$$

$$\Delta p_t = A_{21}(L)\psi_t + A_{22}(L)v_t \quad (2)$$

where Δp_t represents the rate of consumer price inflation and $\Delta c_t - \Delta p_t$ represents the growth rate of real commodity prices. The model decomposes these two series into two distinct innovations: a monetary or nominal component, v_t , and a commodity-specific or real component, ψ_t . These structural innovations are mutually and serially uncorrelated by assumption.⁹

The model is identified by imposing that real commodity prices should be unaffected by a monetary

⁷While there has been some debate in the literature as to the order of integration of U.S. consumer prices, for the remainder of the paper I maintain that CPI inflation is stationary — that consumer prices are $I(1)$ — in my sample.

⁸The trend in real commodity prices in the lower panel of figure 1 is consistent with this interpretation.

⁹Although included in the estimation, constant terms are suppressed for ease of exposition.

innovation in the long-run: $A_{12}(1) = 0$ in equation (1). The rest of the model is left unrestricted, so that v_t can have temporary effects on both variables (and permanent effects on P) while ψ_t can have both permanent and temporary effects on either variable. In other words, the impact effects and short-run dynamics from either shock on either variable are unrestricted. Thus the consumer price level potentially could be shifted permanently by either shock. However, only the real (commodity-specific) shock, ψ_t , can have a permanent effect on the relative commodity price. Permanent changes in ψ_t preclude a simple cointegrating relationship between P and C . Transitory changes in ψ_t that reflect shifts in the supply or demand for commodities can account for the greater volatility of C and low contemporaneous correlation between consumer and commodity prices reported in table 1.

While this specification is not inconsistent with a “cost shock” view of commodity price fluctuations that eventually feed into consumer prices, it is more in the spirit of treating commodity prices as forward-looking asset prices. Most of the commodities in the index are traded internationally with prices set in auction markets. The Granger causality results in table 2 are consistent with commodity prices that react to monetary shocks more quickly than sluggishly-adjusting consumer prices. Theoretically, in the short-run following a monetary shock, sticky consumer prices lead to higher real balances and lower interest rates. In a Hotelling model of storable commodities, an arbitrage condition links a reduction in interest rates to a jump in (real) commodity prices — implying that, like exchange rates, commodity prices should overshoot in response to a monetary shock. Such transitory dynamics are consistent with the economic rationale for the above identification scheme, but not imposed in estimation.

Finally, unlike identifying assumptions that focus on short-run relationships, the long-run relationship modeled and examined here is invariant to changes in monetary policy regimes, international financial structure or exchange rate arrangements, which in turn allows for consideration of a much longer time period for the empirical estimates.

4 Empirical Results

The model of section 3 is estimated as a vector autoregression (VAR) from 1959 Q1 to 2002 Q4, with four lags of each variable.¹⁰ Figure 2 illustrates the dynamic response of each series to a one-time innovation to the monetary (left column) or real (right column) shock. The point estimates are given by the solid

¹⁰The third month of each quarter is used as the quarterly observation. The Akaike and Schwarz Information Criteria suggested lag lengths of 5 and 3 quarters, respectively.

line, while the 68% and 95% confidence regions are shaded in dark and light grey, respectively.¹¹

Consistent with their relative sluggishness, consumer prices rise mildly (less than one-half of one percent) in the period the nominal shock occurs. They continue rising for three to four years, ultimately increasing by almost 2.9%. The identifying restriction ensures that nominal commodity prices rise by the same amount in the long-run. Upon impact, nominal commodity prices overshoot their long-run response by more than a percentage point. They continue to rise for about a year before slowly declining to their long-run response of 2.9%.¹² Importantly, these dynamics are not imposed by the long-run identifying restriction, which only pins down the asymptotic response to the nominal shock.

Turning to the right-hand column of figure 2, the real shock initially causes a 2.9% rise in nominal commodity prices. Within two years the full effect of this shock is reached, at 4.8%. Unlike the response to the nominal shock, there is little statistically discernible overshooting by real commodity prices to the commodity-specific shock.¹³ The real shock has a limited effect (economically or statistically) on consumer prices at any horizon. It is important to recognize that the identifying assumption does not constrain the impulse responses to the real shock.

The forecast error variance decompositions in table 5 indicate that real shocks do impart some high frequency variability to commodity prices, as they are responsible for 33% of the forecast error variance upon impact and over 11% one year later.¹⁴ Asymptotically, the real shock accounts for just over 14% of the forecast error variance for CPI inflation (note shown), but effectively none for the price level.

Conversely, monetary shocks are a significant source of forecast error variability for commodity prices for many years. Initially the nominal shock accounts for 66% of the variance in the nominal commodity price index (55% for the real commodity price); three years later, over half (30%) of the forecast error variance in nominal (real) commodity prices still can be attributed to the monetary disturbance. While the identifying assumption does imply that the nominal shock has a negligible effect on real commodity prices as the forecast horizon approaches infinity, the contribution is non-trivial even ten years out.

The final pieces of evidence for the importance of nominal shocks are historical decompositions,

¹¹These bootstrapped bias-corrected error bands are constructed per Kilian (1998), with 2500 replications.

¹²While inconsistent with rational expectations and complete information, exchange rates also exhibit “delayed overshooting” in response to a nominal shock. This behavior may reflect learning about the nature and duration of the shocks on the part of traders in these markets.

¹³These results suggest that the mean-reversion in commodity prices noted by Deaton and Laroque (1992) is due primarily to the nominal shock.

¹⁴The variance decompositions report what fraction of the k -period ahead forecast error of a series is due to each of the identified shocks.

which use the vector moving average representation to measure the relative contributions of each shock through time.¹⁵ Figure 3 shows the estimated decomposition for consumer prices. Consistent with the evidence and discussion above, much of the high frequency and almost all of the low frequency movement in (detrended) consumer price inflation is due to the accumulated history of monetary shocks. As with the impulse responses and variance decompositions above, changes in the CPI appear to be ultimately a monetary phenomenon — despite the identifying restriction having nothing to say about the relative importance of either shock for consumer prices.

More striking perhaps is the historical decomposition of real commodity price growth in figure 4. Although the real component does play a more sizable role here, the nominal component is still the dominant factor underlying much of the variability in real commodity prices. In particular, a greater proportion of the large increases — and subsequent declines — in the mid- and late-1970s (both preceding the two oil price shocks by about a year) can be attributed to the accumulated history of the nominal shocks.¹⁶ A large positive then negative spike in the late 1980s is about evenly split between the nominal and real components.

5 Conclusion

This paper investigates the endogenous nature of commodity prices with respect to monetary policy by studying the long-run relationship between consumer and commodity prices. Despite the relatively high variability observed for the commodity price series, the identification scheme nonetheless associates a significant portion of the dynamics of commodity prices with nominal shocks to the economy. A positive impulse to this shock yields short-run overshooting and persistent increases in nominal commodity prices. The nominal component accounts for most of the movements in consumer prices as well.¹⁷

This approach has several advantages. The restriction that all prices should eventually respond equiproportionally to a nominal (monetary) shock is consistent with most mainstream macroeconomic theories. It avoids “incredible” identifying restrictions on the timing of the short-run co-movement of the variables in the model. It is invariant to many structural changes in the economy and in policy making. Moreover, the estimated results are consistent with several theories and statistical evidence about

¹⁵The historical decompositions are constructed with eight-year rolling windows of the VMA representation.

¹⁶This evidence is consistent with Barsky and Kilian’s (2001) interpretation of the U.S. experience with stagflation.

¹⁷While not reported here, these results are robust to alternative consumer and commodity price indices. Not surprisingly, the quantitative importance of the commodity-specific shock becomes larger for more narrowly-defined commodity price indices.

commodity price behavior. The results also support other studies that emphasize the contribution of monetary policy to price dynamics, such as Bernanke et al. (1997), who claim that monetary accommodation was the primary cause for the observed inflation following the oil price shocks, and Barsky and Kilian (2001) (as noted previously).

The primary disadvantage of this approach is the possible convolution of shocks.¹⁸ As a bivariate model, at most only two independent structural shocks can be identified. In this study, the real shock is likely a combination of various, conceptually distinct, economic innovations. As such, the model can be viewed as “semi-structural” in the sense that only the nominal shock — the focus of this study — has a straightforward economic interpretation. To the extent that other, non-monetary, factors could yield equiproportional long-run effects on both consumer and commodity prices as well, these results may suggest an upper bound for the contribution of nominal shocks to the time series behavior of commodity prices. It is difficult, however, to name a specific candidate that might significantly contaminate the nominal shock as identified above. Subsequent research that extends this model to a larger collection of macroeconomic variables should help to clarify whether swings in commodity prices have a direct causal role for inflation, or are mostly indicative of structural shocks originating elsewhere.

¹⁸Faust and Leeper (1997) discuss additional considerations for long-run identifying restrictions in VAR models.

Table 1: Comparison of Nominal Price Indices
Annualized Quarterly Growth Rate over Specified Period

	Sample Period			
	1959 Q1 – 2002 Q4	1959 Q1 – 1971 Q4	1972 Q1 – 1982 Q4	1983 Q1 – 2002 Q4
<i>CPI Inflation (excl. Shelter)</i>				
Mean	3.92	2.52	7.24	2.78
Standard Deviation	3.17	1.82	3.38	2.09
<i>IMF Overall World Commodity Price Index</i>				
Mean	2.46	1.10	8.02	-0.11
Standard Deviation	19.09	7.72	17.17	10.25
Correlation w/ Inflation	0.207	-0.039	0.119	0.182

Table 2: Granger-Causality Tests of Consumer and Commodity Prices
Test results for 1959 Q1 – 2002 Q4 sample

Null Hypothesis	<i>F</i> -Statistic	Significance Level
<i>c</i> does not Granger-cause <i>p</i>	5.40	0.0004
<i>p</i> does not Granger-cause <i>c</i>	1.02	0.4000
Δc does not Granger-cause Δp	4.33	0.0023
Δp does not Granger-cause Δc	0.63	0.6445

Notes: Tests constructed using four lags of each variable. Test statistic distributed as F(4,167). ‘*c*’ represents log nominal commodity price index, ‘*p*’ represents log consumer price index, ‘ Δ ’ indicates annualized first differences.

Table 3: Tests of Stationarity of Nominal Price Series
Augmented Dickey-Fuller test results for 1959 Q1 – 2002 Q4 sample

Transformation	Consumer Prices		Commodity Prices	
	ADF Statistic	Largest Autoregressive Root	ADF Statistic	Largest Autoregressive Root
Log level	-1.01	0.999	-1.15	0.989
with trend	-1.56	0.994	-2.05	0.968
First log difference	-2.58	0.846	-6.19	0.296
with trend	-2.59	0.845	-6.23	0.284
Second log difference	-7.22	-1.167	-7.60	-1.143

Notes: A constant and 4 lags were included in each test. Sample size is 176 observations for all tests. Critical values: without trend, -2.88 at the 5% level, -2.57 at the 10% level; with trend, -3.44 at the 5% level, -3.14 at the 10% level. Reject null of a unit root if ADF statistic is less than critical value.

Table 4: Cointegration Tests of Consumer and Commodity Prices
Test results for 1959 Q1 – 2002 Q4 sample

Cointegration Test	Test Statistic	Sample Size	Parameters
Phillips - Ouliaris (1990)	$Z_\rho = -7.332$	176	$n - 1 = 1$
Horvath - Watson (1995)	$W = 2.451$	176	4 lags

Notes: For Phillip-Ouliaris test, null hypothesis is no cointegration; alternative is a spurious cointegrating relationship. Reject null at 5% level if $Z_\rho < -21.5$; reject null at 10% level if $Z_\rho < -18.1$. For Horvath-Watson test, null hypothesis is no cointegration; alternative has cointegrating vector $[1, -1]$. Reject null at 5% level if $W > 10.18$; reject null at 10% level if $W > 8.30$.

Table 5: Share of Variance Decompositions Due to Nominal Shock*Percent, 1959 Q1 – 2002 Q4 sample*

Forecast Horizon (Quarters)	Consumer Prices			Nominal Commodity Prices			Real Commodity Prices		
	0	<i>66.7</i>	66.7	<i>66.7</i>	<i>66.2</i>	66.2	<i>66.3</i>	<i>55.4</i>	55.4
1	<i>69.3</i>	74.6	<i>78.9</i>	<i>58.2</i>	64.2	<i>69.8</i>	<i>47.7</i>	53.7	<i>59.8</i>
2	<i>72.8</i>	80.3	<i>85.9</i>	<i>53.3</i>	62.2	<i>70.2</i>	<i>42.1</i>	51.1	<i>59.6</i>
3	<i>76.6</i>	84.6	<i>90.5</i>	<i>50.7</i>	61.4	<i>70.8</i>	<i>39.0</i>	49.8	<i>59.7</i>
4	<i>80.3</i>	88.5	<i>94.0</i>	<i>48.3</i>	60.1	<i>70.5</i>	<i>39.0</i>	47.5	<i>58.6</i>
6	<i>85.1</i>	93.2	<i>97.3</i>	<i>42.8</i>	57.5	<i>69.9</i>	<i>29.1</i>	43.0	<i>56.3</i>
8	<i>88.0</i>	95.7	<i>98.4</i>	<i>38.2</i>	55.0	<i>69.5</i>	<i>24.0</i>	38.8	<i>54.3</i>
12	<i>90.7</i>	97.9	<i>99.1</i>	<i>31.2</i>	50.8	<i>69.3</i>	<i>18.0</i>	31.8	<i>50.8</i>
16	<i>92.0</i>	98.7	<i>99.4</i>	<i>26.0</i>	47.4	<i>69.0</i>	<i>15.1</i>	26.6	<i>48.0</i>
20	<i>92.6</i>	99.1	<i>99.5</i>	<i>22.9</i>	44.7	<i>69.4</i>	<i>13.5</i>	22.9	<i>45.6</i>
32	<i>93.1</i>	99.6	<i>99.7</i>	<i>16.9</i>	39.3	<i>70.2</i>	<i>10.1</i>	16.1	<i>41.7</i>
40	<i>93.3</i>	99.7	<i>99.8</i>	<i>14.7</i>	37.0	<i>70.7</i>	<i>8.7</i>	13.4	<i>40.1</i>

Notes: Middle number for each series is point estimate of forecast error variance percentage due to the nominal shock at the specified forecast horizon. First and last number for each series (in italics) are the lower and upper 90% bootstrapped confidence intervals for the forecast error variance, respectively, at that horizon.

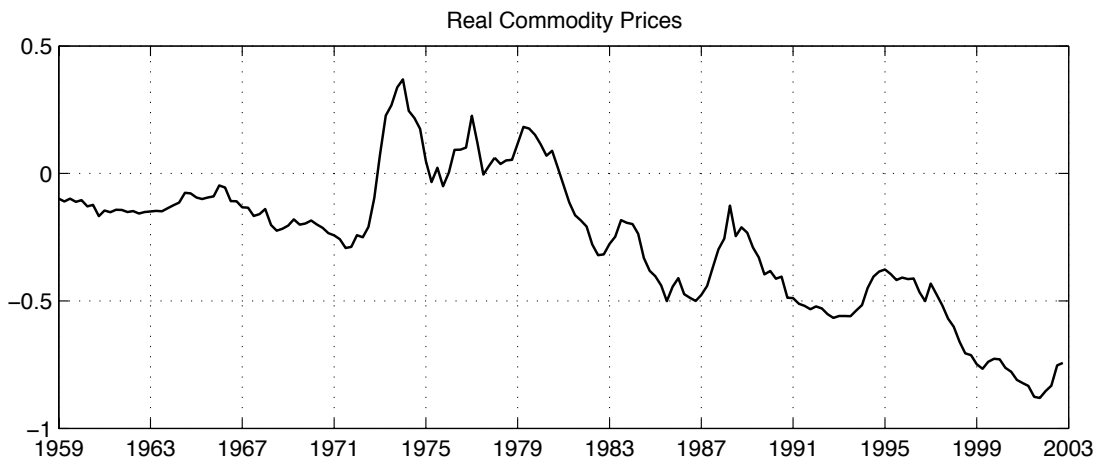
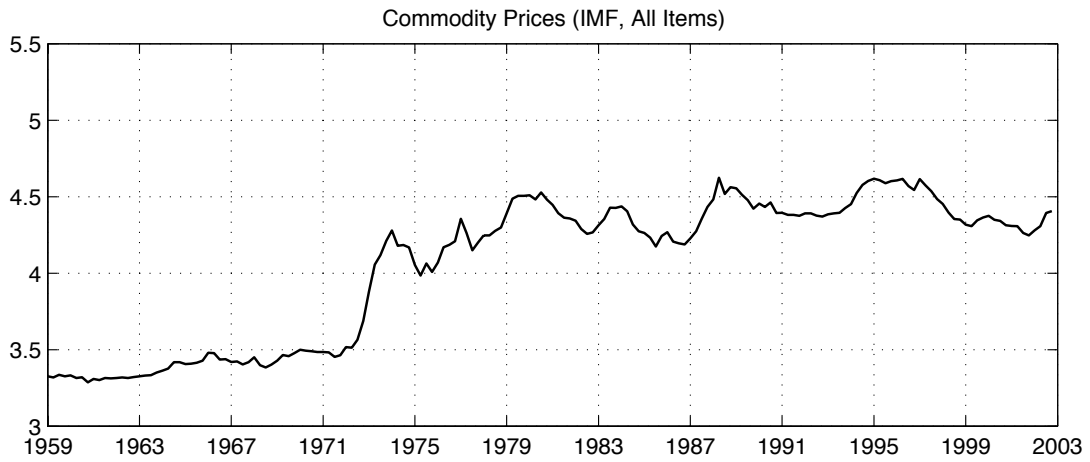
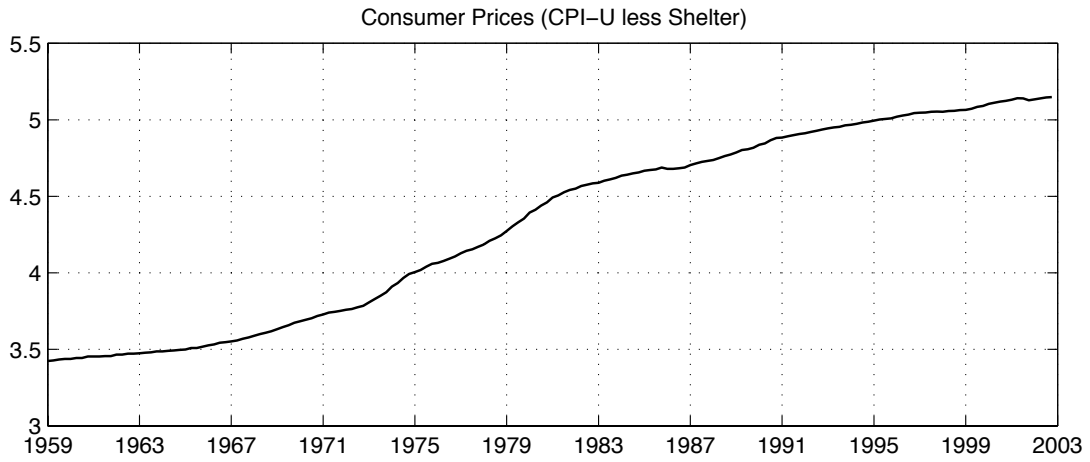


Figure 1: Logarithmic Levels of Consumer and Commodity Prices, 1958 Q1 – 2002 Q4

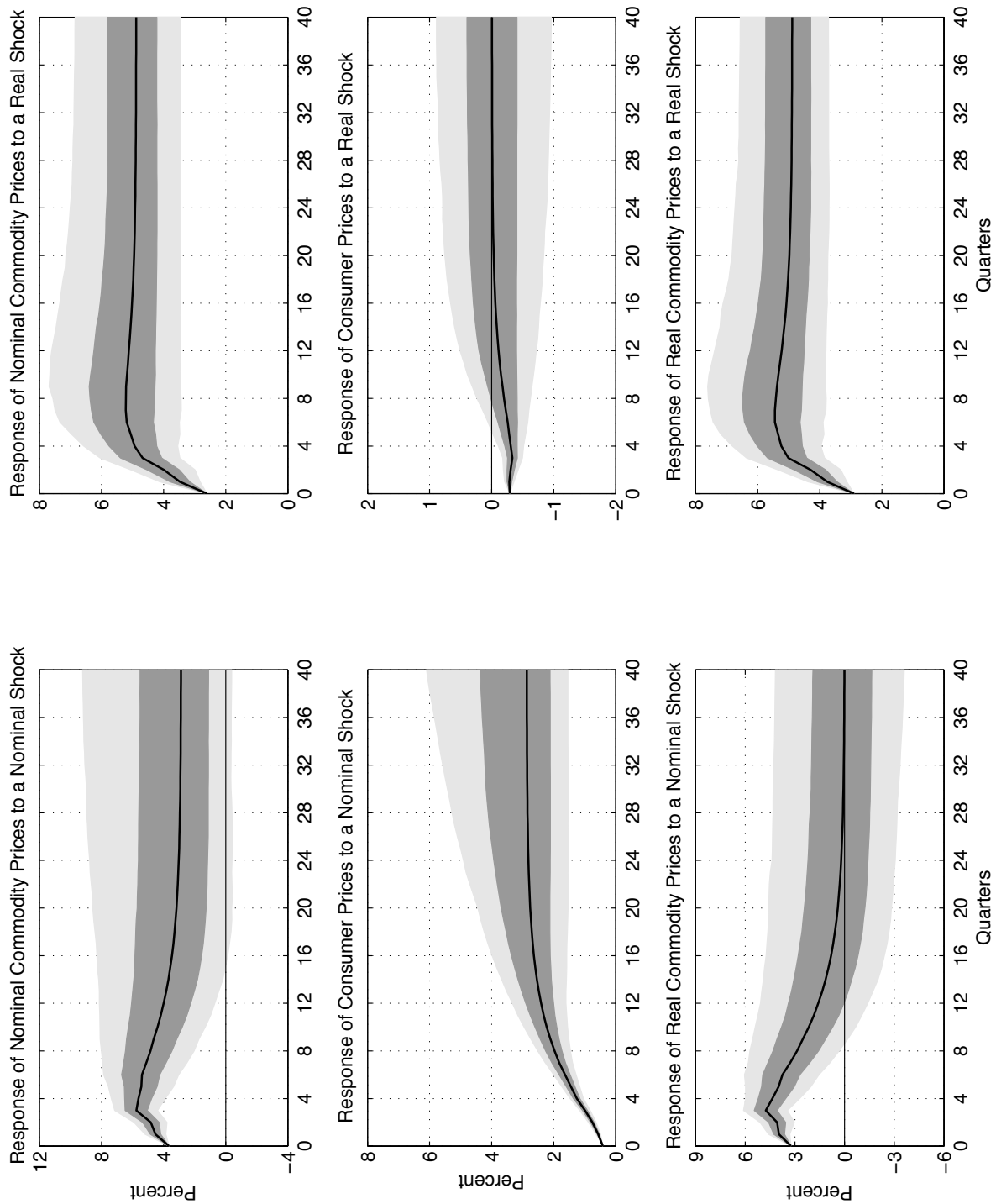


Figure 2: Impulse Responses of Consumer and Commodity Prices, 1959 Q1 – 2002 Q4

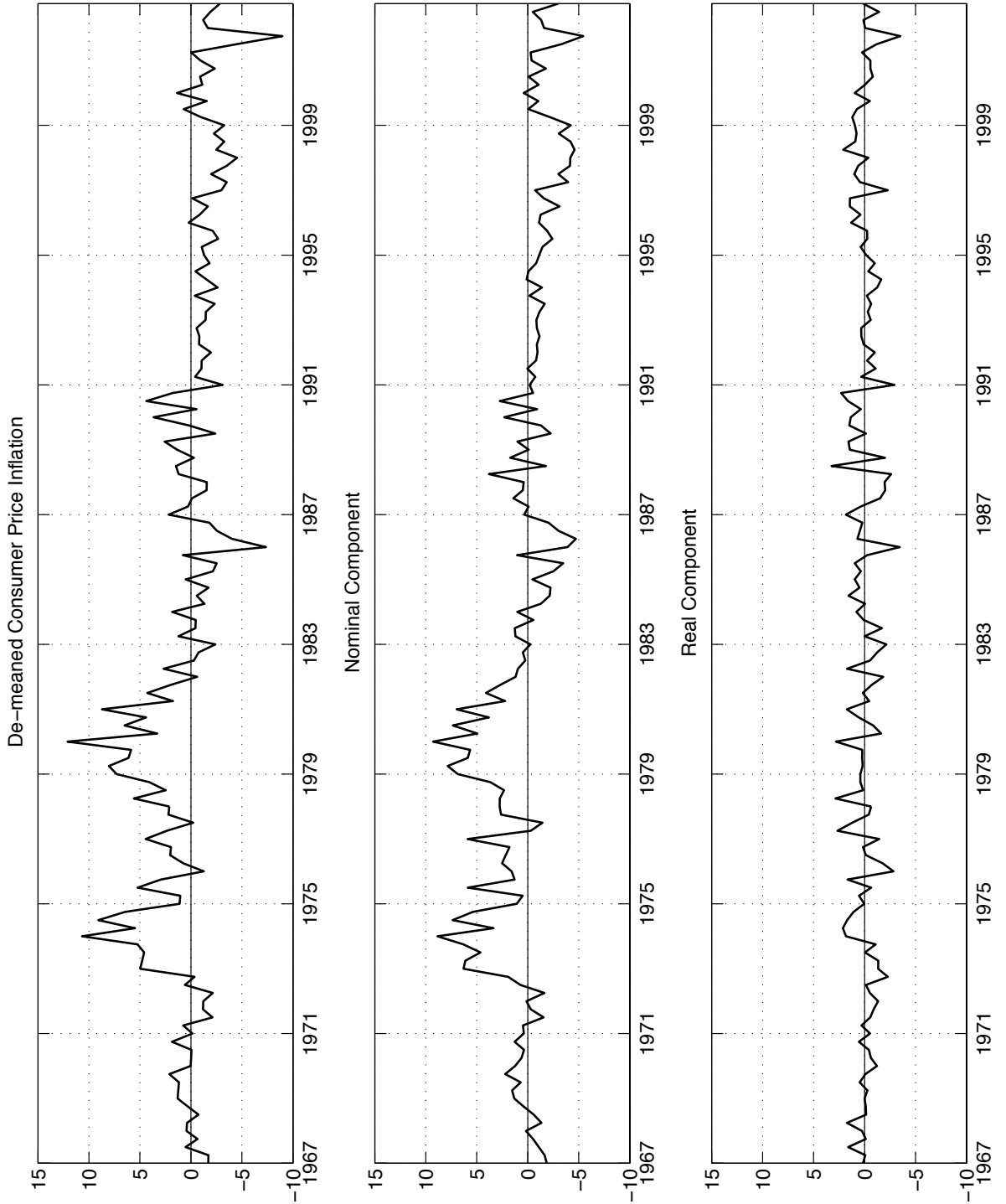


Figure 3: Historical Decomposition of Consumer Price Inflation, 1967 Q1 – 2002 Q4

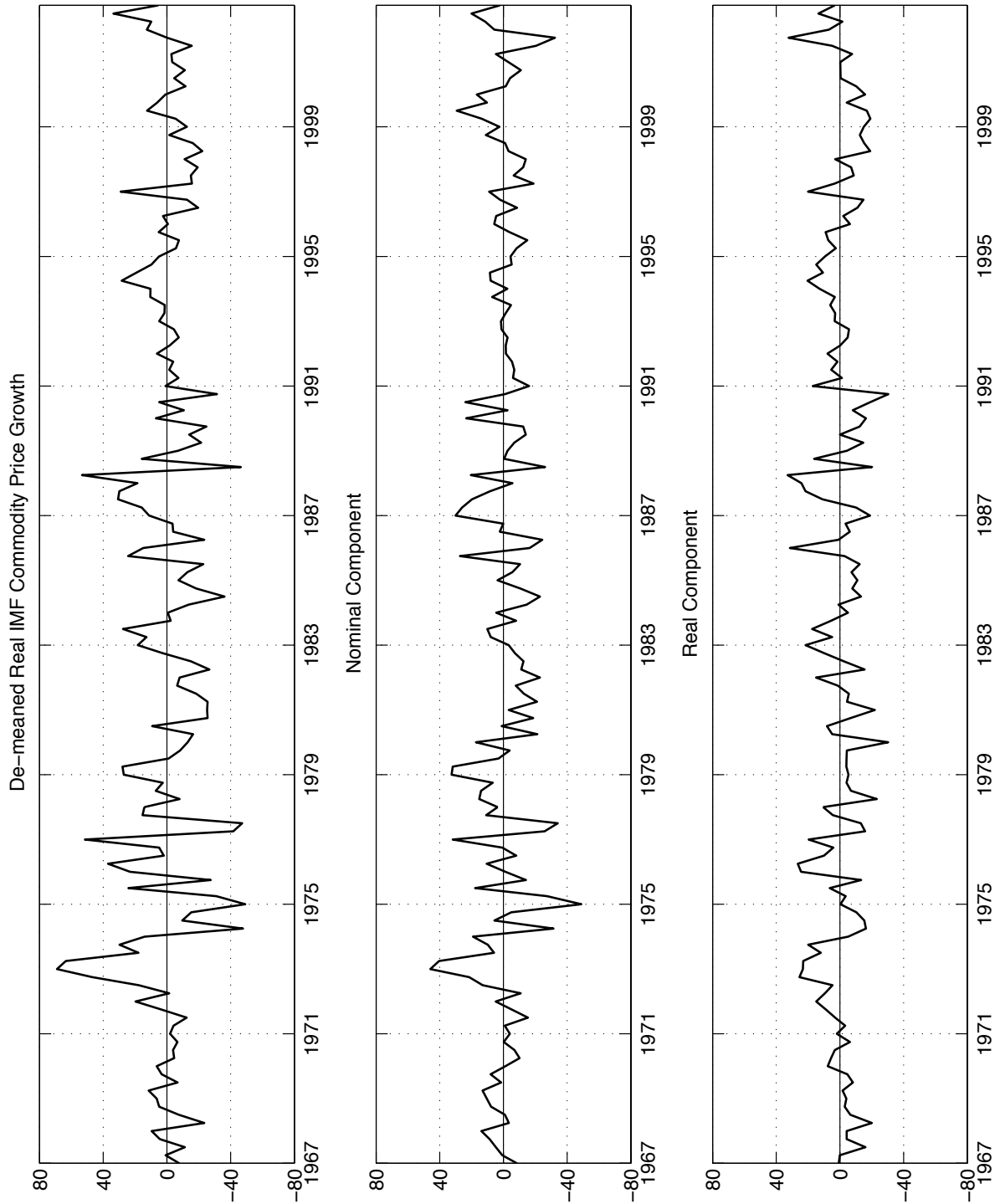


Figure 4: Historical Decomposition of Real Commodity Price Growth, 1967 Q1 – 2002 Q4

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