PRICE TRANSMISSION BETWEEN DJIA, S&P 500 INDEX, PPI AND CPI

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Abstract

Our previous work on month effect in the DJIA, CPI and PPI led us to hypothesize that significant negative September effect that we found for the DJIA might have been caused by changes in the CPI and PPI. This led us to explore the nature of price transmission between the three (we add S&P 500 Index as well). Using VAR analysis and Granger causality analysis we find that the DJIA had a 2-month lagged impact on the CPI in the first two periods (1926-1945 and 1946-1972), and on the PPI in the second period (1946-1972); but in none of the three periods was the DJIA significantly impacted by the PPI or the CPI. For the period 1972-2003, the CPI and PPI were significantly unaffected by the DJIA and the S&P500 Index and also the DJIA and the S&P500 were also not affected significantly by the CPI and PPI. These results follow from both the VAR analysis and Granger causality tests.

(This is an on-going research. We want to add S&P 500 Index data from 1926 [this paper uses monthly S&P 500 Index data from 1973 to 2003]. We will add VAR analysis and Granger causality analysis for our entire data set [this paper does the analysis for three sub-periods: 1926-1945, 1946-1972, 1973-2003]. We will analyze for a bear market and a bull market period. We may add NASDAQ Index to the analysis.)

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INTRODUCTION

Stock prices are affected by a host of factors, including macro-economic indicators and variables. Many studies have analyzed the impact of the major macro-economic variables on stock price. Many studies have also looked at the relationship between inflation and stock prices. But no study looks at the Dow Jones as a measure of stock prices, and PPI as a measure of inflation. To these measures we also add S&P 500 Index and CPI. So we look at two indicators of the stock markets and two indicators of inflation. An understanding of the nature of price transmission between Dow Jones Industrial Average (DJIA), S&P 500 Index, PPI and CPI would be helpful for policy makers, analysts, traders and investors. Most money is invested in the DJIA and S&P 500 stocks, and the understanding of the nature of price transmission would benefit most investor classes.

A number of researchers have documented that *ex post* nominal stock returns and inflation are negatively correlated. This result is surprising since stocks which represent claims on real assets should compensate their holders for inflation. Researchers have also found a negative relation between *ex ante* nominal stock returns and *ex* ante inflation. This finding is also surprising since the Fisher model implies that expected nominal rates should have a one-to-one relationship with expected inflation. Researchers documenting these streams of findings include Lintner (1975), Bodie (1976), Nelson (1976), Fama and Schwert (1977), Jaffe and Mandelker (1977), Gultekin (1983), and Kaul (1987). These studies have mostly focused on asset returns with horizons of a year or less.[1] These findings seek to refute the Fisher's theory of the nominal interest rate or the theory of rational expectations. About 27 years back, Fama and Schwert (1977) suggested that it may be due to some "as yet unidentified phenomena" or to markets that are "inefficient in impounding available information about future inflation . . ."

Modigliani and Cohn (1979) argue that when inflation trend increases real stock prices decrease because economic agents suffer from "inflation illusion." It has been claimed that the illusion arises from not being able to properly differentiate between real and nominal interest rates when the nominal rate includes an inflation premium. It is then claimed that the agents use wrong discount rate when valuing shares. To Modigliani and Cohn (1979), the increase in inflation trend in the 1960s and 1970s credibly explains the prolonged decrease in real stock prices in the U.S. during that period. Ritter and Warr (2000) present evidence that support Modigliani and Cohn (1979) hypothesis. They use firm-level data and find valuation errors leading to depressed stock prices during inflationary periods. Feldstein (1980) develops a market equilibrium model of share valuation that shows that an increase in steady state inflation lowers as a result of the way depreciation costs and capital gains are treated in the tax codes. So Modigliani and Cohn (1979), and Feldstein (1980) hypothesize, inflation wears away the long-run real value of stocks.

Fama (1981) forwards the proxy hypothesis according to which the negative correlation is presumed to be induced by inflation and real equity returns reacting

[1] Fama and Schwert (1977) find a negative relation between *ex ante* stock returns and expected inflation using monthly, quarterly, and semiannual data.

oppositely to news about future real output growth.[2] Similar argument is put forward by Geske and Roll (1983), Benderly and Zwick (1985) and Kaul: that the inverse relationship is a spurious result of the dual effect that revisions in expected future output growth bear on expected future dividends and current inflation.

Santoni and Moehring (1994) argue that the "puzzling observation" may be a result of inappropriate use of currently published inflation indices.

Fisher and Seater (1993) and King and Watson (1997) develop rigorous time series tests of long-run neutrality propositions to analyze the integration properties of economic variable and the identification of structural shocks. King and Watson (1997) use their method to measure the long-run nominal interest rate response to a permanent inflation shock in the U.S. Koustas and Serletis (1999) use the King and Watson methodology to measure the long-run interest rate response to a permanent inflation shock in 11 industrialized countries. Bullard and Keating (1995) and Crosby and Otto (2000) use the long-run neutrality tests to measure the long-run response of real-output and the capital stock, respectively in a large number of developed countries and developing countries. Thus, the recently developed time-series tests have been used to measure long-run response of interest rates, real output, and the capital stock to a permanent inflation shock. Rapach (2001) uses the econometric methodology of King and Watson (1997) to measure the long-run response of real stock-prices – measured using the real value of a broad index of share prices – to a permanent inflation shock in

[2] Aarstol (2000) argues that controlling for output growth does not fully eliminate this negative correlation. He argues that agency costs increase with the relative price variability (RPV) that tends to accompany inflation, and find evidence that variations in RPV explain much of the negative relationship between inflation and real equity returns that persists after controlling for output growth.

16 industrial countries. He finds considerable support for long-run inflation neutrality with respect to real stock prices. The overall finding is that inflation does not erode the long-run real value of stocks.

In a rebuttal of the large body of evidence that document that the stock market tends to perform poorly during inflationary time periods Ely and Robinson (1998) contend that these results are mostly obtained from models structured to estimate the short-run relationships between stock returns and inflation. Ely and Robinson (1998) use a reduced-form approach and recent advances in the theory of cointegration to explore the international evidence on the relationship between stock prices and goods prices. They claim this approach allows them to test if stocks maintain their value relative to goods prices and whether these response patterns depend on the source of inflation shocks. For most of the countries analyzed, their results indicate that stocks do maintain their value relative to movements in overall price indexes. One notable exception is that stocks do not maintain their value relative to goods price following real output shocks in the US.

Using data covering 1802-1990 for the United Kingdom and the United States, Boudoukh and Richardson (1993) regress total nominal stock returns on total inflation and find that stocks are a better hedge against inflation at a five-year as opposed to a one-year horizon. Solnik and Solnik (1997) using multi-country panel data provides evidence supporting the finding of Boudoukh and Richardson (1993) that the Fisher model holds at long horizons (5 years), using 180 years of US data.

Olesen, Jan Overgaard (2000), based on a simple structural model find for the Danish stock market over the post-World War II-period support for the hedge property.

Taufiq Choudhry (2000) analyzes the relationship between stock returns and inflation in four high inflation (Latin and Central American) countries and find evidence of a positive relationship between current stock market returns and current inflation; the result substantiate that stock returns act as a hedge against inflation. The study also finds some evidence of an inverse relationship between current real returns and current and one-period lagged inflation. Mohammed Omran and John Pointon (2001) examine the impact of the inflation rate on the performance of the Egyptian stock market and find inflation rate has had an impact upon the Egyptian stock market performance generally.

Authors have also looked at effect of inflation on capital stock and output. Crosby and Otto (2000) use a structural vector autoregression (VAR) model for thirty-four countries and discover that for the majority of these countries there is no statistically significant long run effect of inflation on the capital stock. Bullard and Keating (1995) also use a structural VAR model and find that a permanent shock to inflation is not associated with a permanent movement in the level of real output for most countries in their large sample of post-war economies. The main exceptions are certain low inflation countries, in which permanent inflation shocks permanently increase the level of output**.** Also using VAR model, Rapach (2003) finds that a permanent increase in inflation lowers the long-run real interest rate in each of 14 industrialized countries; a permanent increase in inflation also increases the long-run real output level in a number of countries.

Hamid and Dhakar (2002) using DJIA data for the last century, find that the mean as well as the median of monthly percentage changes of September were negative and significanly lower than for the other eleven months. Hamid and Dhakar (2003) find that the mean of monthly percentage CPI changes of June and September for the period 1926 to 2003 was significantly higher than the other months. Hamid and Thirunavukkarasu (2004) find that the mean of monthly percentage changes of the all commodities index (part of PPI) for July over the period 1913 to 2004 was the highest of all the months (the median for July was the second highest). Hamid and Habib (2006) find negative September effect for S&P 500 Index for the period 1926 to 2004. In this study, we want to analyze if there might be a causal relationship between the three (we as well add S&P 500 Index). We want to explore if the high mean CPI changes in June and September and the high mean PPI changes in July may have caused the negative September effect on the DJIA and the S&P 500 Index.

DATA AND METHODOLOGY

We use monthly change data for the all commodities index as a measure of Producer Price Index (PPI), monthly change data for Consumer Price Index (CPI), and change for the Dow-Jones Industrial Average (DJIA) from 1926 to 2003. We obtained the data from the Stocks, Bonds, Bills and Inflation (SBBI) Yearbook. We obtained monthly change data for S&P 500 Index for the years 1965 to 2003 from Data Stream. (We want to use monthly S&P 500 Index data from the SBBI Yearbook for 1926-2003 in the final version of this paper.) To gain deeper insight, we sub-divided the data period based on presumed structural changes in the economy as follows:

- 1926 to 1945: which includes the Great Depression years, and the Second World War;
- 1946 to 1972: which includes the Breton Woods fixed exchange rate era, and the break down of that era in 1972;

- 1973 to 2003: which includes the volatile world we live in since the first oil crisis of 1973.

Descriptive statistics of entire data

The Producer Price Index (PPI) data in the following table refers to the monthly all commodities PPI between 1926 and 2003. The Consumer Price Index (CPI) data refers to the urban seasonally unadjusted CPI between 1926 and 2003. The DJIA refers to the monthly change in the Dow Jones Industrial Average between 1926 and 2003. The SP refers to the monthly change in the S&P500 Index values from 1965 to 2003.

	PPI	CPI	DJIA	SP
Mean	0.001755	0.001122	0.004498	0.001685
Median	0.000000	0.000000	0.008557	0.000200
Maximum	0.108200	0.059000	0.401805	0.169500
Minimum	-0.036000	-0.056000	-0.307008	-0.122300
Std. Dev.	0.014707	0.008671	0.074243	0.043724
Skewness	2.026761	0.250128	0.237739	0.257547
Kurtosis	14.45085	14.57700	8.568647	4.518896
Jarque-Bera	1918.189	1745.604	406.0668	16.72039
Probability	0.000000	0.000000	0.000000	0.000234
Sum	0.547600	0.350200	1.403518	0.262800
Sum Sq. Dev.	0.067272	0.023384	1.714246	0.296331
Observations	312	312	312	156

VAR Method

A vector autoregression (VAR) analysis was performed to test whether there is a causal relationship between CPI, PPI and DJIA and S&P500 and vice versa. The tests were performed for three different time intervals (1926-1945, 1946-1972, and 1973-2003) to avoid the bias of watershed events. The relationship between DJIA and S&P500 was performed only for 1973-2003.

The vector autoregression is commonly used for forecasting systems of interrelated time series and for analyzing the dynamic impact of random disturbances on the system of variables. The VAR approach sidesteps the need for structural modeling by treating every endogenous variable in the system as function of the lagged values of all the endogenous variables in the system. The mathematical expression of a VAR is:

$$Y_t = A_1\, y_{t-1} + \ldots. + A_p y_{t-p} + \ldots\ldots. + Bx_t + \varepsilon_t$$

Where y_t is a k-vector of endogenous variables, x_t is a d vector of exogenous variables, A1,……A_p and B are matrices of coefficients to be estimated, and ε_t is a vector of innovations that may be contemporaneously correlated but are uncorrelated with their own lagged values and uncorrelated with all the right-hand side variables.

ANALYSIS OF RESULTS

VAR Analysis

The VAR analysis (tables shown below) indicates that in the interval 1926-1945 (Table 1), the change in CPI was significantly (t=2.15) affected by the changes in the DJIA lagged by two time periods. But the changes in the PPI were not affected by the changes in the DJIA. Also the changes in the PPI and CPI did not significantly affect the changes in the DJIA. The change in PPI was also significantly (t=2.85) affected by the changes in CPI lagged by two time periods. But PPI did not significantly affect CPI.

Table 2 shows that for the period 1946-1972 CPI and PPI were significantly affected by DJIA lagged by two time periods (t=3.18; t=2.89 respectively). Again, the

DJIA was not significantly influenced by the lagged values of PPI and CPI. Also CPI lagged by one period and two periods affected PPI (t=2.75 and t=7.42). Once again, we do not find impact of PPI on CPI.

Finally, for the period 1972-2003, Table 3 shows that the CPI and PPI were significantly unaffected by the DJIA and the S&P500 Index and also the DJIA and the S&P500 Index were not affected significantly by the CPI and PPI. However, one and two period lagged values of CPI affected PPI (t=2.52 and t= 10.60 respectively). One period lagged values of CPI also affected CPI (t=4.28). Two period lagged values of PPI affected CPI (t=2.44) and PPI to a lesser extent (t=-2.00). These trends can also be seen in the impulse response functions that follow the tables.

Table 1: VAR 1926 – 1945
The Producer Price Index (PPI) data in the following table refers to the monthly all commodities PPI. The Consumer Price Index (CPI) data refers to the urban seasonally unadjusted CPI. The DJIA refers to the Dow Jones Industrial Average.

Vector Autoregression Estimates; Sample (adjusted): 1926Q3 1945Q4
Included observations: 78 after adjustments; Standard errors in ( ); t-statistics in [ ]

	CPI	PPI	DJIA
CPI(-1)	0.022570	0.162959	-0.514959
	(0.11671)	(0.12933)	(1.42525)
	[0.19338]	[1.26004]	[-0.36131]
CPI(-2)	0.020898	0.366578	-0.099650
	(0.11587)	(0.12840)	(1.41500)
	[0.18036]	[2.85498]	[-0.07042]
PPI(-1)	0.171381	0.182406	1.741649
	(0.10236)	(0.11343)	(1.25002)
	[1.67428]	[1.60811]	[1.39330]
PPI(-2)	0.083026	-0.076665	0.678708
	(0.10316)	(0.11432)	(1.25983)
	[0.80479]	[-0.67062]	[0.53873]
DJIA(-1)	0.009561	0.002316	0.157689
	(0.00971)	(0.01076)	(0.11856)
	[0.98480]	[0.21526]	[1.33008]
DJIA(-2)	0.022662	0.020511	-0.060872
	(0.01055)	(0.01169)	(0.12885)
	[2.14790]	[1.75432]	[-0.47244]
C	-0.002269	-0.002943	0.009210
	(0.00127)	(0.00140)	(0.01546)
	[-1.79255]	[-2.09851]	[0.59586]
R-squared	0.154379	0.238765	0.067398
Adj. R-squared	0.082918	0.174436	-0.011413
Sum sq. resids	0.005288	0.006493	0.788520
S.E. equation	0.008630	0.009563	0.105385
F-statistic	2.160332	3.711590	0.855181
Log likelihood	263.6883	255.6804	68.50073
Akaike AIC	-6.581751	-6.376420	-1.576942
Schwarz SC	-6.370251	-6.164921	-1.365442
Mean dependent	-0.004238	-0.005981	-0.003062
S.D. dependent	0.009011	0.010525	0.104788
Log likelihood		588.4693	
Akaike information criterion		-14.55049	
Schwarz criterion		-13.91600	

Table 2: VAR Output: 1946 – 1972
The Producer Price Index (PPI) data in the following table refers to the monthly all commodities PPI between 1946 and 1972. The Consumer Price Index (CPI) data refers to the urban seasonally unadjusted CPI between 1946 and 1972. The DJIA refers to the Dow Jones Industrial Average between 1946 and 1972.

Vector Autoregression Estimates; Sample: 1946Q1 1972Q4
Included observations: 108; Standard errors in ( ); t-statistics in [ ]

	CPI	PPI	DJIA
CPI(-1)	0.393017	0.437969	0.190167
	(0.10068)	(0.15906)	(1.35325)
	[3.90367]	[2.75349]	[0.14053]
CPI(-2)	0.104961	1.232147	0.003306
	(0.10515)	(0.16613)	(1.41338)
	[0.99818]	[7.41689]	[0.00234]
PPI(-1)	-0.045448	0.030082	-0.118218
	(0.05026)	(0.07940)	(0.67550)
	[-0.90434]	[0.37888]	[-0.17501]
PPI(-2)	0.041676	0.098933	-0.792968
	(0.04727)	(0.07469)	(0.63541)
	[0.88160]	[1.32465]	[-1.24795]
DJIA(-1)	0.001388	0.017113	0.001476
	(0.00747)	(0.01179)	(0.10034)
	[0.18586]	[1.45093]	[0.01471]
DJIA(-2)	0.022186	0.031847	-0.017357
	(0.00698)	(0.01102)	(0.09379)
	[3.17947]	[2.88879]	[-0.18505]
C	0.000462	0.000893	0.010896
	(0.00061)	(0.00097)	(0.00825)
	[0.75290]	[0.92073]	[1.31992]
R-squared	0.252695	0.586238	0.021784
Adj. R-squared	0.208301	0.561659	-0.036328
Sum sq. resides	0.003753	0.009366	0.677965
S.E. equation	0.006095	0.009630	0.081930
F-statistic	5.692059	23.85033	0.374865
Log likelihood	401.1970	351.8037	120.5773
Akaike AIC	-7.299944	-6.385253	-2.103284
Schwarz SC	-7.126102	-6.211411	-1.929442
Mean dependent	0.001154	0.003083	0.008438
S.D. dependent	0.006851	0.014545	0.080481
Log likelihood		877.0634	
Akaike information criterion		-15.85303	
Schwarz criterion		-15.33150	

Table 3: VAR Output: 1973 - 2003

The Producer Price Index (PPI) data in the following table refers to the monthly all commodities PPI. The Consumer Price Index (CPI) data refers to the urban seasonally unadjusted CPI. The DJIA refers to the Dow Jones Industrial Average. The SP refers to the S&P500 Index.

Vector Autoregression Estimates
Sample: 1973Q1 2003Q4
Included observations: 124; Std. errors in ( ); t-statistics in [ ]

	CPI	PPI	SP	DJIA
CPI(-1)	0.388852	0.302550	0.096190	-0.818250
	(0.09095)	(0.12003)	(0.58622)	(0.43716)
	[4.27568]	[2.52068]	[0.16409]	[-1.87173]
CPI(-2)	-0.061376	1.394239	-0.121546	-0.582358
	(0.09968)	(0.13156)	(0.64255)	(0.47917)
	[-0.61570]	[10.5976]	[-0.18916]	[-1.21534]
PPI(-1)	0.067229	-0.047964	-0.072372	-0.116607
	(0.04959)	(0.06545)	(0.31964)	(0.23837)
	[1.35571]	[-0.73286]	[-0.22641]	[-0.48919]
PPI(-2)	0.111360	-0.120478	-0.266769	-0.004488
	(0.04559)	(0.06017)	(0.29385)	(0.21913)
	[2.44278]	[-2.00245]	[-0.90784]	[-0.02048]
SP(-1)	0.007592	-0.011737	-0.003524	0.076647
	(0.01450)	(0.01914)	(0.09349)	(0.06972)
	[0.52347]	[-0.61318]	[-0.03770]	[1.09939]
SP(-2)	0.010255	0.009534	-0.065394	0.012722
	(0.01451)	(0.01915)	(0.09354)	(0.06976)
	[0.70666]	[0.49777]	[-0.69907]	[0.18238]
DJIA(-1)	-0.013660	0.004894	0.075704	0.011242
	(0.01931)	(0.02548)	(0.12444)	(0.09280)
	[-0.70757]	[0.19209]	[0.60836]	[0.12115]
DJIA(-2)	0.016094	0.025294	-0.119888	-0.071780
	(0.01903)	(0.02511)	(0.12265)	(0.09146)
	[0.84583]	[1.00726]	[-0.97752]	[-0.78482]
C	0.001940	-0.001661	0.003652	0.014376
	(0.00088)	(0.00116)	(0.00567)	(0.00422)
	[2.20722]	[-1.43181]	[0.64472]	[3.40300]
R-squared	0.257112	0.631499	0.022609	0.102056
Adj. R-squared	0.205433	0.605864	-0.045383	0.039590
Sum sq. resides	0.006191	0.010783	0.257225	0.143048
S.E. equation	0.007337	0.009683	0.047294	0.035269
F-statistic	4.975166	24.63437	0.332524	1.633790
Log likelihood	438.1591	403.7537	207.0930	243.4728
Akaike AIC	-6.921921	-6.366995	-3.195048	-3.781819
Schwarz SC	-6.717223	-6.162297	-2.990350	-3.577121
Mean dependent	0.004560	0.005582	0.001179	0.006779
S.D. dependent	0.008231	0.015424	0.046256	0.035988
Log likelihood		1293.848		
Akaike information criterion		-20.28787		
Schwarz criterion		-19.46908		

Figure 1: Impulse Response Functions: 1926 - 1945

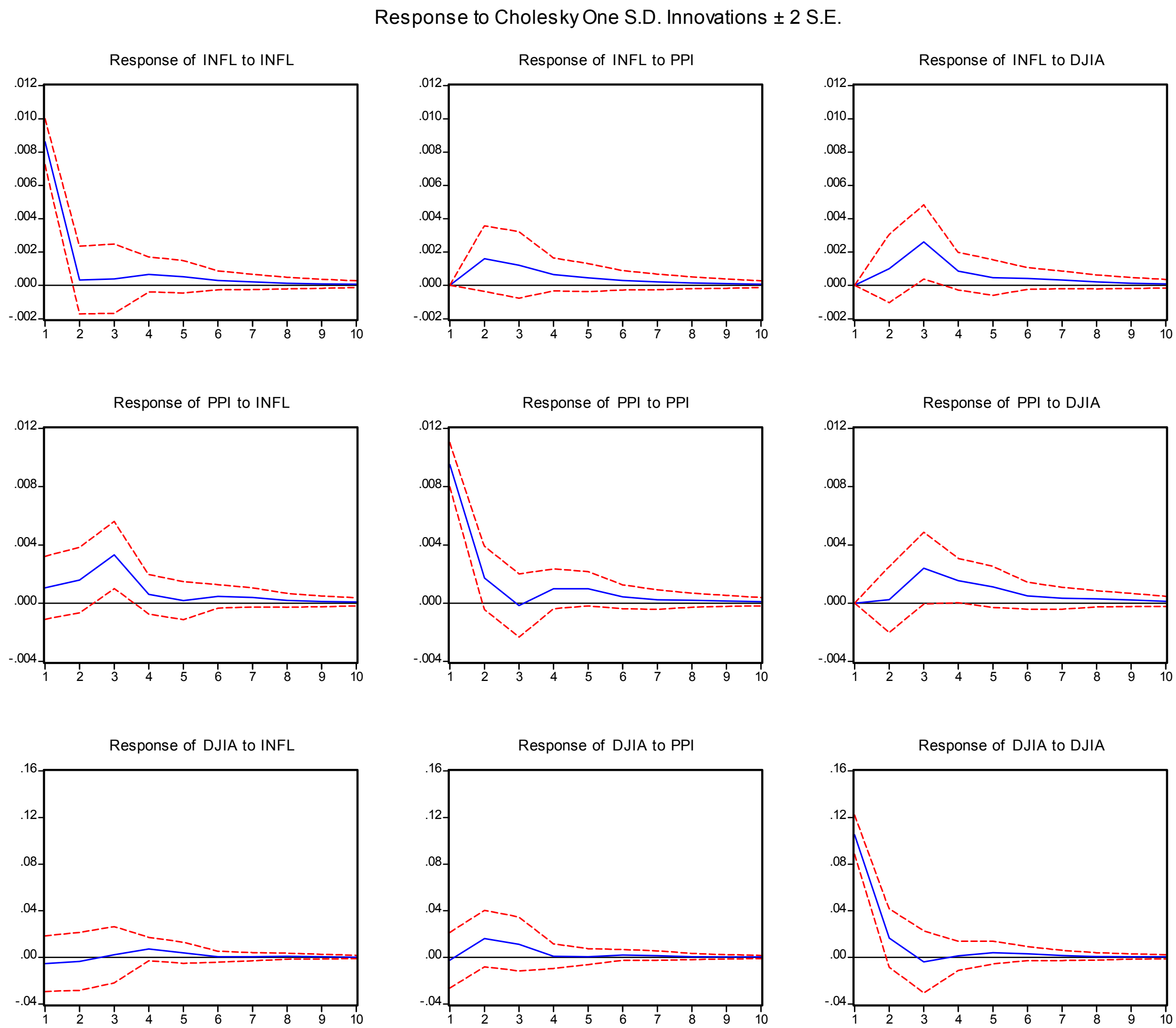


Figure 2: Impulse Response Functions: 1946-1972

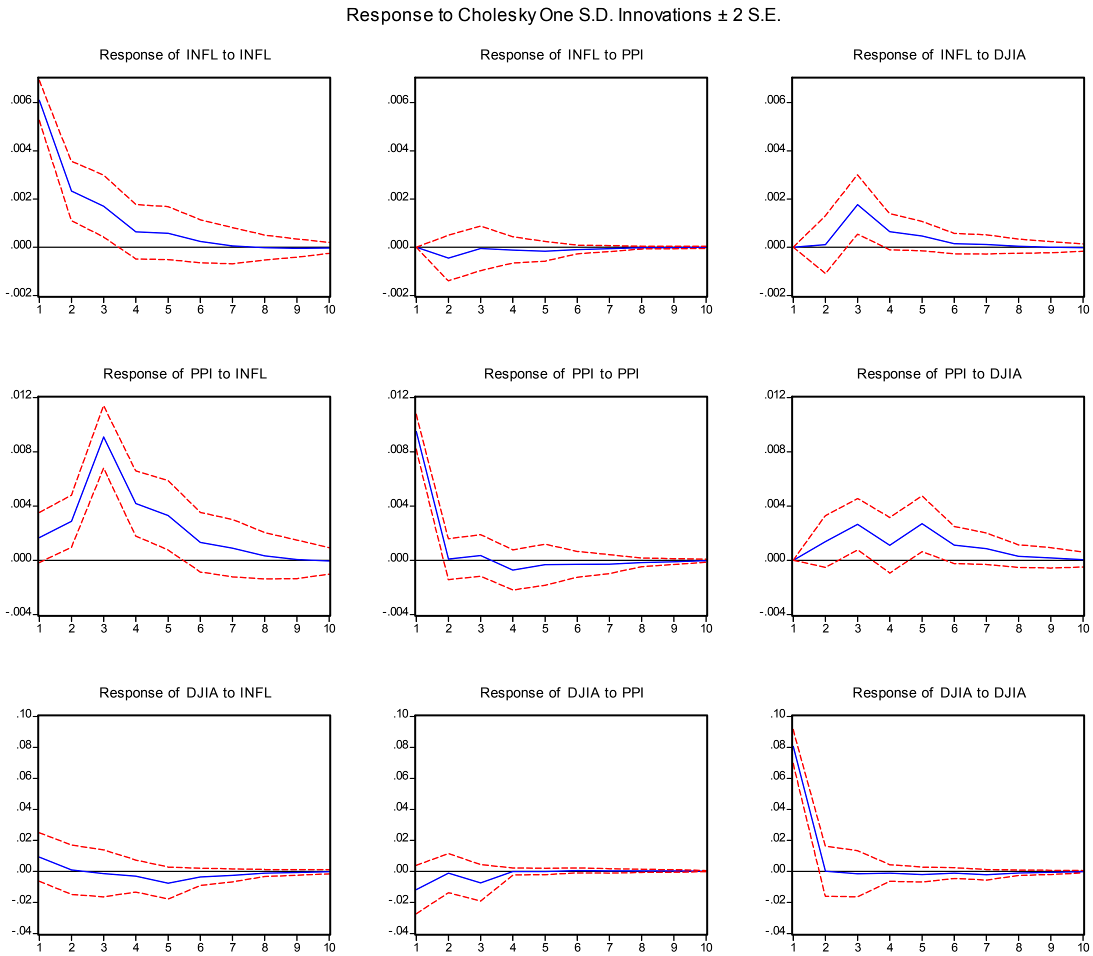
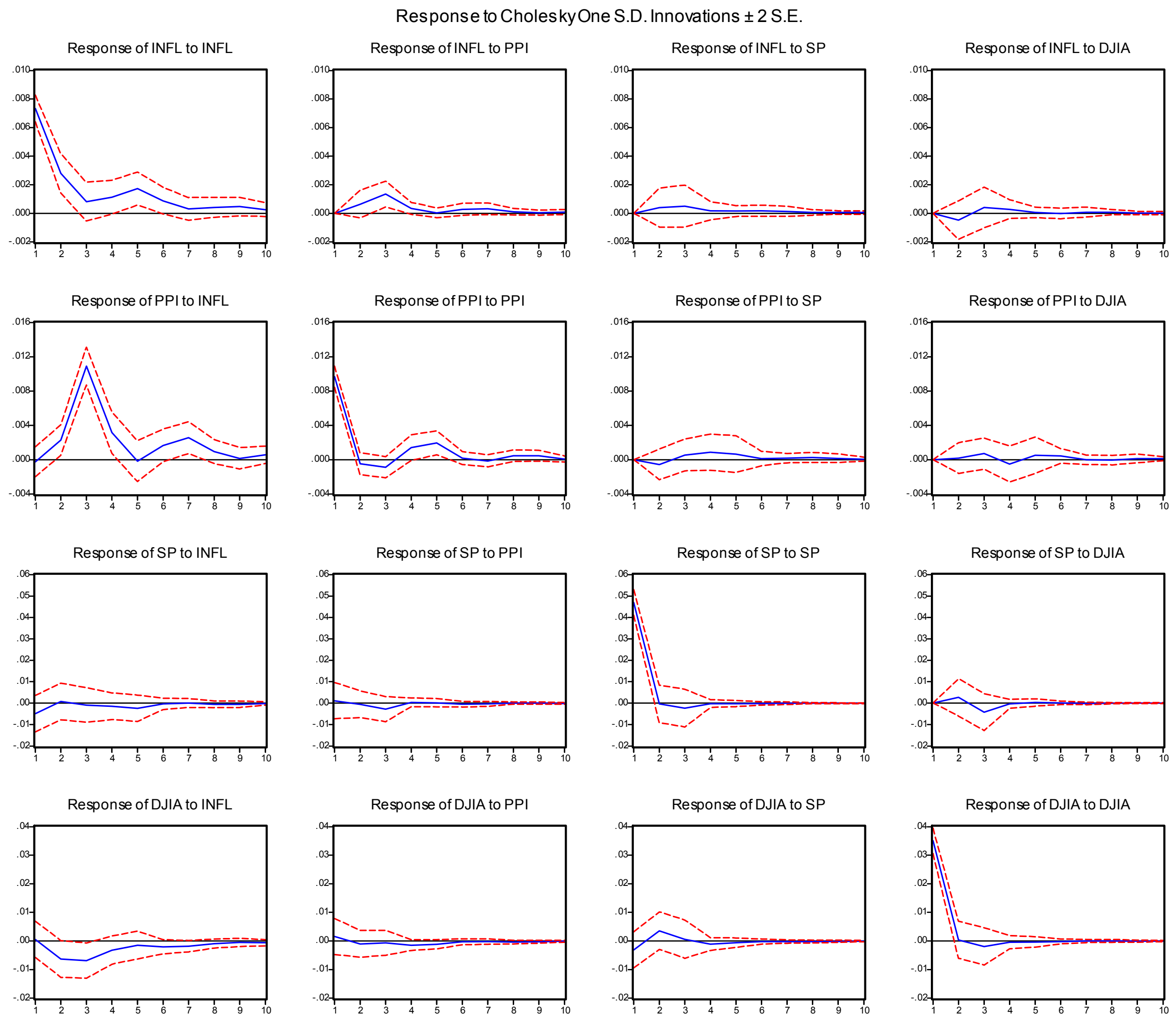


Figure3: Impulse Response Functions: 1973-2003



Granger-causality tests

The Granger (1969) approach to the question of whether X causes Y is to analyze how much of the current value of Y can be explained by the past values of Y and then to see whether adding lagged values of X can improve the explanatory power. Y is said to be Granger-caused by X if X helps in the prediction of Y, or equivalently if the coefficients on the lagged X's are statistically significant. Two-way causation is frequently the case; X Granger causes Y and Y Granger-causes X. It is important to note that the statement "X Granger causes Y" does not imply that it is the effect or the result of X. Granger-causality measures precedence and information content but does not by itself indicate causality in the more common use of the term.

The Granger-causality test is used to determine if there is a "granger" casual relationship between CPI, PPI and DJIA, S&P500 indices. Monthly data from 1926 to 2003 for the all commodities PPI and urban seasonally unadjusted CPI were obtained from the Bureau of Labor Statistics. The monthly data fro DJIA was obtained from Data Stream (1926-2003) and the S&P500 index data from 1965 to 2003 were obtained from the DataStream. The data series were checked for stationarity using unit roots test. The Augmented Dickey fuller test for unit roots showed that the data series were stationary in the levels. The Granger causality test was performed to check if there was a bidirectional causal relationship between PPI, CPI and DJIA, S&P500. The tests revealed that while the DJIA Granger-causes the CPI and PPI, there is no causal relationship from CPI and PPI to the DJIA. Furthermore, there is no causal relationship either ways from the S&P 500 to the CPI and PPI.

Table 4: Output of Granger-causality tests

The Producer Price Index (PPI) data refers to the monthly all commodities PPI between 1926 and 2003. The Consumer Price Index (CPI) data refers to the urban seasonally unadjusted CPI between 1926 and 2003. The DJIA refers to the Dow Jones Industrial Average between 1926 and 2003. The SP refers to the S&P500 index values from 1965 to 2003.

Pairwise Granger Causality Tests
Sample: 1926Q1 2003Q4
Lags: 4

Null Hypothesis:	Obs	F-Statistic	Probability
PPI does not Granger Cause CPI	308	1.89637	0.11101
CPI does Granger Cause PPI		48.0073	3.7E-31
SP does not Granger Cause CPI	152	0.94911	0.43760
CPI does not Granger Cause SP		0.74932	0.56001
DJIA does Granger Cause CPI	308	3.85947	0.00449
CPI does not Granger Cause DJIA		0.06992	0.99104
SP does not Granger Cause PPI	152	0.31729	0.86605
PPI does not Granger Cause SP		0.09611	0.98356
DJIA does Granger Cause PPI	308	7.24064	1.4E-05
PPI does not Granger Cause DJIA		0.83692	0.50259
DJIA does not Granger Cause SP	152	0.57309	0.68259
SP does not Granger Cause DJIA		0.51693	0.72340

CONCLUSION

The DJIA had a 2-month lagged impact on the CPI in the first two periods (1926-1945 and 1946-1972), and on the PPI in the second period (1946-1972); but in none of the three periods was the DJIA significantly impacted by the PPI or the CPI. For the period 1972-2003, the CPI and PPI were significantly unaffected by the DJIA and the S&P500 Index and also the DJIA and the S&P500 Index were also not affected significantly by the CPI and PPI. So the negative September effect in the DJIA and S&P

500 Index does not appear to be caused by CPI or PPI. In the first period, the change in PPI was also significantly affected by the changes in CPI lagged by two time periods; but PPI did not significantly affect CPI. In the second period, CPI lagged by one period and two periods affected PPI; nce again, we do not find impact of PPI on CPI. In the third period one and two period lagged values of CPI affected PPI; one period lagged values of CPI also affected CPI; two period lagged values of PPI affected CPI and PPI to a lesser extent. These results follow from both the VAR analysis and the Granger causality tests. These trends can also be seen in the impulse response functions that follow the tables.

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