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Todd E. Clark and Stephen J. Terry

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Abstract

From Bayesian estimates of a vector autoregression (VAR) which allows for both coefficient drift and stochastic volatility, we obtain the following three results. First, beginning in approximately 1975, the responsiveness of core inflation to changes in energy prices in the United States fell rapidly and remains muted. Second, this decline in the passthrough of energy inflation to core prices has been sustained through a recent period of markedly higher volatility of shocks to energy inflation. Finally, reduced energy inflation passthrough has persisted in the face of monetary policy which quickly became less responsive to energy inflation starting around 1985.

Keywords: oil price shocks, monetary policy, time-varying parameters

JEL Classification: C11, E31, E52

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* Economic Research Department, 1 Memorial Drive Kansas City, Missouri 64198. Telephone: (816) 881-2575 (Clark) and (816) 881-2795 (Terry). Fax: (816) 881-2199. Email: todd.e.clark@kc.frb.org and stephen.terry@kc.frb.org. The views expressed in this paper are those of the authors and do not necessarily reflect the views of the Federal Reserve Bank of Kansas City or the Federal Reserve System.

1 Introduction

From the end of 2001 to the summer of 2008, the spot price of a barrel of crude oil (West Texas Intermediate) rose over 600 percent, from around \$20 to almost \$150. Over the same period, average quarterly core inflation in the United States was two percent. This sustained increase in oil prices, and its evidently tame impact on core price pressures, ignited a great deal of interest in the economic effects of energy price shocks. A growing body of recent studies, including Hooker (2002), De Gregorio et al. (2007), van den Noord and André (2007), Blanchard and Galí (2008), and Chen (2009) and primarily focusing on oil prices, reports a pronounced reduction over past decades in the passthrough of energy prices to broader inflation in the United States and elsewhere. This recent work is related to a large literature on the broader economic effects of oil price shocks dating back to at least Hamilton (1983), which documents a statistical link between oil price shocks and postwar recessions in the United States. Subsequent studies and rejoinders, including Hooker (1996), Barsky and Kilian (2001), and Hamilton (2003), debate the robustness of important real and inflationary effects of oil price shocks to different price specifications, assumptions of exogeneity, and, importantly, evidence of a weakening of the effects of oil prices in more recent data. Useful surveys of the economic literature on energy prices can be found in Segal (2007) and Kilian (2008).

The extant evidence on declining passthrough has generally relied on relatively simple methods of assessing changes over time — such as split samples or estimates over rolling samples of data. This paper seeks to more formally assess the evidence of changes over time by examining estimates of models with time-varying parameters and stochastic volatility. More specifically, we use Bayesian VAR methodology similar to that in Cogley and Sargent (2005), Primiceri (2005), and Benati (2008) to examine the passthrough of energy price inflation to core inflation in the United States. Our VAR in energy price inflation, core (excluding food and energy) inflation, a measure of economic activity, and the federal funds rate also generalizes — by modestly expanding the set of variables and allowing time variation in parameters and volatilities — some common reduced form Phillips curves that include energy inflation (see, e.g., Gordon (1997, 1998), Brayton, et al. (1999), and Chen (2009)).

Our model estimates yield a pronounced reduction from approximately 1975 onwards in the passthrough of energy price inflation to core inflation in the United States. Declining passthrough is evident in both smaller reduced form coefficients on energy price inflation and smaller impulse responses of core inflation to identified energy price shocks. We supplement these results with historical decompositions of core inflation during three periods ranging from the mid-1970s to this century. The rapid reduction in the inflationary passthrough of energy prices accompanies both declining energy consumption shares in the United States economy and a recent prolonged increase in the volatility of energy price shocks. On the basis of reduced form and structural VAR evidence, we also find that monetary policy has been less responsive to energy price inflation since approximately 1985.

The paper is organized as follows: Section 2 outlines the Bayesian methodology and data used in this study, Section 3 describes the reduction in passthrough of energy price inflation, Section 4 presents evidence of changing volatility in each equation of our VAR, Section 5 discusses changing monetary policy responsiveness to energy prices, and Section 6 concludes.

2 Time-Varying Parameters VAR with Stochastic Volatility

Based on a Phillips curve, Chen (2009) finds evidence of significant time variation in the coefficients on oil prices (holding other coefficients constant). However, as emphasized by Cogley and Sargent (2005), the evolution over time of macroeconomic relationships in the context of a statistical model may be reflected by both drift in coefficients and changes in the volatility of innovations. To capture both sources of time variation, in a model more general than a Phillips curve, we examine the following VAR:

$$
\begin{aligned}
y_t &= X_t' B_t + A_t^{-1} \Sigma_t \epsilon_t, \\
X_t' &= I_n \otimes [1, y_{t-1}', \ldots, y_{t-k}'], \\
Var(\epsilon_t) &= I_n,
\end{aligned}
$$

with

$$
A_t = \begin{bmatrix} 1 & 0 & \cdots & 0 \\ a_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ a_{n1,t} & \cdots & a_{nn-1,t} & 1 \end{bmatrix} \quad \text{and} \quad \Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \cdots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{n,t} \end{bmatrix}.
$$

The parameter vectors B_t, $a_{i,t}$, and σ_t follow the random walk processes:

$$
B_t = B_{t-1} + u_t,
$$

$$a_{i,t} = a_{i,t-1} + v_{i,t},$$
$$\log \sigma_t = \log \sigma_{t-1} + e_t,$$

where $a_{i,t}$ is a vector containing all non-zero, non-one elements of the i-th row of A_t and σ_t is the diagonal of Σ_t. The random walk disturbances are normally distributed: $u_t \sim N(0, Q)$, $v_{i,t} \sim N(0, S_i)$, and $e_t \sim N(0, Z)$, where Z is diagonal.

Here y_t is four-dimensional, with quarterly United States data. We include three lags of core inflation (inflation excluding food and energy components), energy inflation, a measure of real economic activity, and the effective federal funds rate. This parsimonious specification allows for tractability in our posterior sampling scheme, discussed below. Core and energy prices are measured by the personal consumption expenditures (PCE) price indexes. Inflation rates are computed as annualized log percent changes ($400 \ln(P_t/P_{t-1})$). In our benchmark specification, real activity is measured by the unemployment gap, specifically the difference between the quarterly unemployment rate and the Congressional Budget Office's (CBO's) estimate of the natural rate of unemployment. We use the unemployment gap for consistency with the Phillips curve literature mentioned above. However, our results are robust to instead measuring economic activity with the level of the unemployment rate or with an output gap computed as the percent difference between real GDP and the CBO's estimate of potential output.

From estimates of a model relating actual inflation and survey measures of inflation expectations to a time-varying unobserved trend in inflation, Clark and Davig (2008) conclude that survey-based measures of long-run inflation expectations are essentially trend inflation. Therefore, in our benchmark specification, we make use of the survey-based long-run (5- to 10-year-ahead) PCE inflation expectations series used in the Federal Reserve Board of Governor's FRB/US econometric model to detrend core PCE inflation and detrend and convert the effective federal funds rate into real terms.[1] However, our basic results are robust to instead leaving inflation and the interest rate in levels terms.[2]

We define the energy inflation variable in the model as PCE energy inflation less core

[1] The FRB/US measure splices econometric estimates of inflation expectations from Kozicki and Tinsley (2001) early in the sample to 5- to 10-year-ahead survey measures compiled by Richard Hoey and, later in the sample, to 10-year-ahead values from the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters.

[2] Although we do include a constant term in the baseline VAR results reported in the paper, the posterior implied mean values of detrended core inflation and the unemployment gap fluctuate closely around zero for most of our sample, confirming that long-term inflation expectations and the CBO's natural rate are adequate measures of underlying statistical trends.

PCE inflation, weighted by energy's share in nominal consumer spending. This specification of energy inflation follows the approach used in Phillips curve work (e.g., Gordon (1997, 1998) and Brayton, et al. (1999)).[3] In the literature on the economic effects of oil prices, transformations of energy price series incorporating a large number of possible non-linearities have generated considerable interest.[4] However, the results in Edelstein and Kilian (2007a, 2007b) suggest that a linear and symmetric specification of energy price inflation such as ours may be appropriate, and Hooker (2002) presents evidence that the declining impact of oil prices on core inflation in later years is robust to alternative non-linear specifications.

We estimate the model over a sample of 1965:Q1-2008:Q2 using Bayesian methods, specifically a Metropolis-within-Gibbs posterior sampler. Following Benati (2008), we generally use the methodology of Cogley and Sargent (2005), with the exception that A is allowed to vary over time, as in Primiceri (2005). More precisely, in our Markov chain Monte Carlo algorithm, conditional on prior information and draws for A_t, Σ_t, and other parameters, the linearity of the model and Gaussian nature of the error term allow a draw of the coefficients B_t to be obtained from a standard application of the Carter and Kohn (1994) smoother.[5] Similarly, conditional on B_t, Σ_t, and the supplemental parameters, draws of the $a_{i,t}$ come from independent applications of the Carter and Kohn (1994) smoother. The log stochastic volatilities in Σ_t are then drawn conditionally for each variable using the Cogley-Sargent version of the Metropolis algorithm developed by Jacquier, et al. (1994). With B_t, A_t, and Σ_t in hand, weakly informative natural conjugate priors yield tractable inverse-Wishart conditional posterior distributions for Q, S_i, and Z. We first perform 25,000 burn-in draws, followed by an additional 25,000 draws, retaining every fifth draw. Prior distributions for the initial values B_0, A_0, and Σ_0 are specified using information from a training sample of 1950:Q1-1964:Q4. Following Cogley and Sargent (2005), our posterior

[3] Relative energy inflation is weighted by the two-period moving average (the two-period average maintains consistency with the chain weighting used in the construction of all NIPA data) of the nominal share of energy in total personal consumption expenditures. In order to preserve the actual volatility of energy inflation (the fact that energy inflation is much more volatile than core inflation), we then normalize by the full-sample mean of energy's share. Studies such as Gordon (1998) compute a relative inflation rate by subtracting core inflation from total. With aggregate inflation a weighted average of components, it is easy to show that our relative measure is conceptually the same as the measure used in Gordon (1998) and other Phillips curve work.

[4] See, for instance, Hooker (1999) or Hamilton (2003).

[5] As in Cogley and Sargent (2005), our prior on the amount of time variation in the VAR coefficients is relatively uninformative. The prior variance is set to .001 times the variance-covariance matrix of coefficients estimated in a training sample, with degrees of freedom equal to the total number of coefficients plus one (the minimum allowable).

sampler truncates explosive draws of B_t. In the case of an explosive draw, we "backstep" until drawing stable coefficients. To reduce the frequency of explosive draws, following Del Negro (2003) we use an informative prior in obtaining estimates of B_0 from the training sample.[6] More complete details describing the methodology used here can be found in Cogley and Sargent (2005) and Primiceri (2005).

3 Declining Passthrough

We document a reduction in recent decades in the passthrough of energy price inflation to core inflation based on three results from our VAR analysis: evidence of a downward drift in the sum of the reduced form coefficients on energy inflation in the core inflation equation, a decline in the impulse response of core inflation to identified energy inflation shocks, and qualitative differences among historical decompositions of core inflation during three periods ranging from the early 1970s to the end of our sample.

3.1 Drifting Coefficients

As a simple measure of the effect of energy price changes on core inflation, we present in Figure 1 posterior medians over time and the 70 percent credible sets for the sum of the reduced form coefficients on energy inflation in the equation for core prices in our time-varying parameters VAR. In the early 1970s, both the posterior medians and all credible-set values are positive, peaking just before 1975, in 1974:Q2. The medians then quickly move downward, and in 1984:Q2 the posterior bands include zero. From this date throughout most of the remainder of the sample, it is impossible to differentiate the coefficient sum from zero based on the 70 percent intervals. Given our Bayesian methodology, the posterior probability of a decline in the coefficient sum is simple to calculate and interpret.[7] Evidence for a decline in passthrough during the ten years from 1975:Q1 to 1985:Q1 is strong, with a posterior probability of 96.2 percent. It is notable that, although we incorporate a measure of energy inflation adjusted for changes in energy expenditures, a large part of the estimated

[6]In our training sample, before PCE core and energy inflation become available in 1959:Q2, we use measures of core and energy inflation calculated from a longer available series of prices for energy goods (which excludes the energy services component incorporated after 1959:Q2) and hold the energy expenditure weights constant at their value in 1959:Q2. Similarly, our long-run inflation expectations series becomes available in 1960:Q1, so before this date expectations are held constant at the mean value of core PCE inflation from 1950:Q1-1959:Q4. Due to small market volume early in the training sample, for the 1949-1959 period we measure the short-term interest rate with the 3-month Treasury bill rate instead of the federal funds rate.

[7]As in Cogley, Primiceri, and Sargent (2007).

reduction in energy price passthrough occurs during a period of reduced energy consumption shares in the United States economy.[8] Also, even though the time-varying parameters framework is normally thought of as being best suited for capturing gradual coefficient change, the speed of the decline is striking, with a 91.9 percent probability of decline in the sum of energy coefficients during the five years from 1975:Q1 to 1980:Q1.[9]

3.2 Impulse Responses

We identify shocks to each variable recursively through their ordering in the VAR. Following studies such as Davis and Haltiwanger (2001) and Blanchard and Galí (2008), energy price inflation is ordered first, so that innovations to that variable are assumed exogenous with respect to all other current-period shocks. Admittedly, papers including Rotemberg and Woodford (1996) and Barsky and Kilian (2001) challenge the appropriateness of this identifying assumption. However, in our analysis, ordering energy inflation after core inflation and the unemployment gap yields results qualitatively identical to those we report for the energy-first identification. As to the remainder of the model, core inflation is ordered before the unemployment gap, following Primiceri (2005). Finally, incorporating a standard assumption in the structural VAR literature, interest rates appear last, reflecting the presumption that monetary policy is able to respond immediately to developments in the economy while the economy responds to policy with at least a one-quarter lag.

Based on our identification scheme and posterior draws at six dates (1975:Q1, 1980:Q1, 1985:Q1, 1990:Q1, 1995:Q1, and 2008:Q2), we calculate impulse responses to an energy inflation shock. To facilitate comparison, the size of the shock at each date is normalized to the standard deviation of the full-sample OLS residual of the energy inflation equation, equal to 11.6 percent. Figure 2 plots the median point estimates of the responses of both energy and core inflation, along with 70 percent credible sets. Two quarters after the shock the response of energy inflation drops close to zero, in a pattern that is consistent over time. In 1975:Q1 and 1980:Q1, we see a corresponding increase in core inflation, which peaks after two or three quarters and then gradually declines. However, from 1985:Q1 through 2008:Q2

[8] As a nominal percentage of personal consumption expenditures, after peaking in 1981:Q2 at 9.3 percent, spending on energy goods and services fell quickly to a share of 7.6 percent only three and a half years later in 1985:Q1.

[9] Based on Phillips curve analysis of the inflationary effects of oil price changes, Hooker (2002) presents similar evidence of a fall in the passthrough of energy prices to inflation for the United States, and De Gregorio et al. (2007), van den Noord and André (2007), and Chen (2009) extend the result to a broad range of other countries.

the impulse response of core prices to energy inflation shocks is either indistinguishable from zero or briefly negative. In this sense, our results corroborate some prior evidence generally based on cruder methodology (split sample or rolling window estimation) for assessing changes over time. In particular, Blanchard and Galí (2008) describe a decline in the inflationary impact of oil price shocks in the United States and some European countries, and Herrera and Pesavento (2007) also uncover evidence of smaller impulse responses to oil price shocks in United States data.

3.3 Historical Decompositions

On the basis of our identification scheme for shocks to each variable, we perform three historical decompositions of core inflation, for the periods 1970:Q1-1976:Q4, 1978:Q1-1984:Q4, and 2000:Q1-2008:Q2.[10] For each historical interval and posterior draw, we fix the model's parameters at their values in the middle of the period. Then, we compare the realized values of core inflation (specifically, realized values of core inflation detrended by long-run expectations) to three series derived from the fixed model parameters: 1) the baseline (from the starting point of the interval) VAR forecast of core inflation over the interval, 2) the sum of this base forecast and the cumulative effects on core inflation of the identified shocks to energy prices, and 3) the base projection plus the cumulative effects of identified shocks to core inflation itself. Figure 3 presents the posterior mean values of these decompositions for each historical period, along with energy inflation in each period.

Our decomposition suggests that shocks to energy inflation contributed strongly to the pronounced increases (and subsequent decreases) of core inflation around 1974 and 1980, but had little impact on core inflation in the current decade. In the mid-1970s, shocks to energy prices accounted for a very large portion of the sharp rise in core inflation, as well as the subsequent decline. Similarly, in the late 1970s and early 1980s, shocks to energy price inflation largely account for the rise and fall of core inflation. More recently, from 2000:Q1-2008:Q2, innovations to core inflation itself dominate the effects of energy price shocks. In this episode, nearly all of the variation in core inflation is accounted for by shocks to inflation; energy shocks have virtually no impact on core inflation. These changes in decompositions are consistent with the evidence of drifting coefficients and declining impulse responses presented above.

[10]These intervals were roughly chosen to include the October War and Arab oil embargo (1973-1974), the Iranian Revolution and subsequent outbreak of the Iran/Iraq war (1978-1980), and the recent protracted period of increasing oil prices, respectively.

4 Changing Volatility

With time-varying parameters methodology similar to that used in this paper, both Cogley and Sargent (2005) and Primiceri (2005) present evidence of declining volatility in innovations to United States macroeconomic variables beginning in the early 1980s. Figure 4 plots posterior medians for the standard deviation of the reduced form residuals and identified innovations of our VAR. We report marked declines in the volatility of shocks to the unemployment gap, core inflation, and the real federal funds rate from 1980 onwards, consistent with this earlier work. Energy inflation shocks exhibit a different pattern, however, with notable spikes in volatility corresponding to the periods of the October War and Arab oil embargo, the Iranian Revolution and the outbreak of the Iran/Iraq War, the oil price collapse of 1986, and the first Gulf War.[11] From the mid-1990s onward, the volatility of innovations to energy inflation has trended sharply upwards.[12]

5 Monetary Policy

A monetary policy less responsive to energy-specific price shocks might be expected to amplify the passthrough of energy price changes to core inflation. Below we present evidence that, from approximately 1985 onwards, monetary policy has indeed become less responsive to energy prices. However, combined with the reduced energy passthrough documented above, this result indicates that the effects of less responsive monetary policy in response to energy price changes have not been sufficient to generate a stable or increased transmission of energy inflation to core inflation.

5.1 Drifting Coefficients

Information about changes in the stance of monetary policy towards energy inflation might be reflected in either time variation in the VAR coefficients or the evolving impulse response of the federal funds rate to energy price shocks. As a simple measure of the responsiveness of monetary policy, we consider the sum of the reduced form coefficients on energy price inflation in the equation for the real federal funds rate, plotted in Figure 5.[13] This

[11] For a description of the historical circumstances surrounding each episode, see, for instance, Barsky and Kilian (2004).

[12] Using a split sample estimation, Edelstein and Kilian (2007a) report increasing volatility in the innovations of energy price shocks from 1988 onwards.

[13] This reaction function allows the policy interest rate to respond separately to core inflation and energy inflation. In other models, such as those of Cogley and Sargent (2005), Primiceri (2005), and Chen (2009), the policy interest rate simply responds to total inflation, with no distinction between core and energy

sum drifts upwards throughout the 1970s, peaking in 1983:Q3. Afterwards, the posterior distribution shifts quickly downwards, and in 1987:Q1 the 70 percent credible sets include zero. Later in the sample the posterior median fluctuates around zero with substantial uncertainty. Although a decline in the reduced form response of the real federal funds rate to energy inflation begins approximately ten years later than the corresponding change in the core inflation equation, evidence of a rapid downward shift is just as strong. From 1985:Q1 to 1990:Q1, the energy inflation coefficient sum declined with 96.3 percent posterior probability.

5.2 Impulse Responses

In addition to the reduced form evidence presented above, the response of the federal funds rate to an identified energy inflation shock also changes rapidly around 1985.[14] Figure 6 plots posterior medians and 70 percent credible sets for the responses of the unemployment gap and the real federal funds rate to a shock to energy inflation in 1975:Q1, 1980:Q1, 1985:Q1, 1990:Q1, 1995:Q1, and 2008:Q2. As in Figure 2, for comparability the size of the energy price shock is normalized to the full sample estimate of 11.6 percent. In 1980:Q1 and 1985:Q1, the real rate rises for three or four quarters on impact, then gradually declines. This path is accompanied by a pronounced increase in the unemployment gap one to two years after the energy price shock. From 1990:Q1 onwards, however, the response of the real federal funds rate is either indistinguishable from zero or negative in the face of an identified energy inflation shock, and the response of the unemployment gap essentially disappears. Our results are consistent with those in Bernanke, et al. (1997), Hooker (2002), and Herrera and Pesavento (2007), which document a more muted response of monetary policy to oil price shocks in recent decades.[15]

6 Conclusion

With extant evidence on declining passthrough of energy prices to core inflation typically based on simple methods, in this paper we use more general techniques for assessing the evidence of change over time. Specifically, we use Bayesian methodology to estimate a VAR

inflation.

[14] As in Section 3.2, an alternative identification scheme with energy inflation entering the VAR later did not qualitatively change the results presented here.

[15] Our reduced form and impulse response evidence do not, however, rule out the possibility that monetary policy became more responsive to inflation *as a whole* later in our sample, as argued in, for instance, Clarida, et al. (2000).

with time-varying parameters and stochastic volatility. Our VAR generalizes some common reduced form Phillips curves that include energy inflation (see, e.g., Gordon (1997, 1998), Brayton, et al. (1999), and Chen (2009)). According to our model estimates, beginning in approximately 1975, core inflation in the United States quickly became less responsive to changes in energy prices. Statistically speaking, by 1985, passthrough from energy to core inflation declined to zero. This conclusion is based on time variation in the estimated reduced form relationship between energy inflation and core prices, vanishing impulse responses to identified energy price shocks, and qualitative changes among historical decompositions of core inflation. The speed of the fall in passthrough is remarkable, as is the continuation of weakness in energy's inflationary effects through a recent period of pronounced increases in the volatility of shocks to energy prices. The responsiveness of monetary policy to energy inflation has also changed importantly over time. From approximately 1985 onwards, shifts in the reduced form link between real interest rates and energy prices, in addition to the elimination of a contractionary response to identified energy inflation shocks, indicate that monetary policy has been less responsive to energy inflation.

Needless to say, our statistical evidence of change begs the question of what might have caused the change — a topic we leave for future research. To this point, a wide range of explanations for weaker passthrough of energy prices to broader inflation has been put forward in previous work. Some of these proposed causes, such as a reduction in the consumption share of energy in the United States economy or changes in monetary policy, are related to our analysis. Other suggested explanations include a reduction in real wage rigidity and the beneficial effects of a low-inflation environment. The interested reader is referred to studies including Hooker (2002), De Gregorio, et al. (2007), Segal (2007), Blanchard and Galí (2008), Kilian (2008), and Chen (2009) for a more complete discussion of these alternative possibilities.

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Figure 1: Sum of Coefficients on Energy Inflation in the Core Inflation Equation

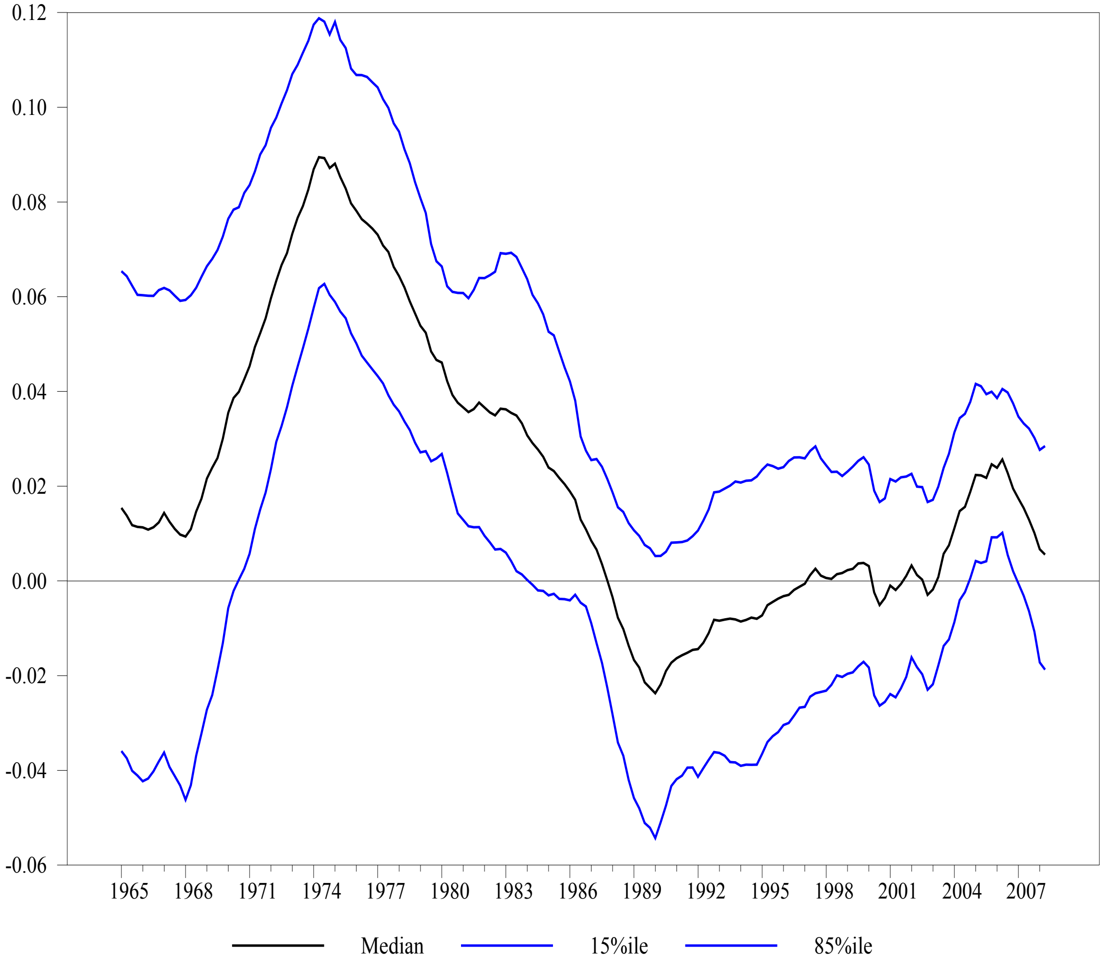

Figure 2: Impulse Responses to a Positive Energy Inflation Shock

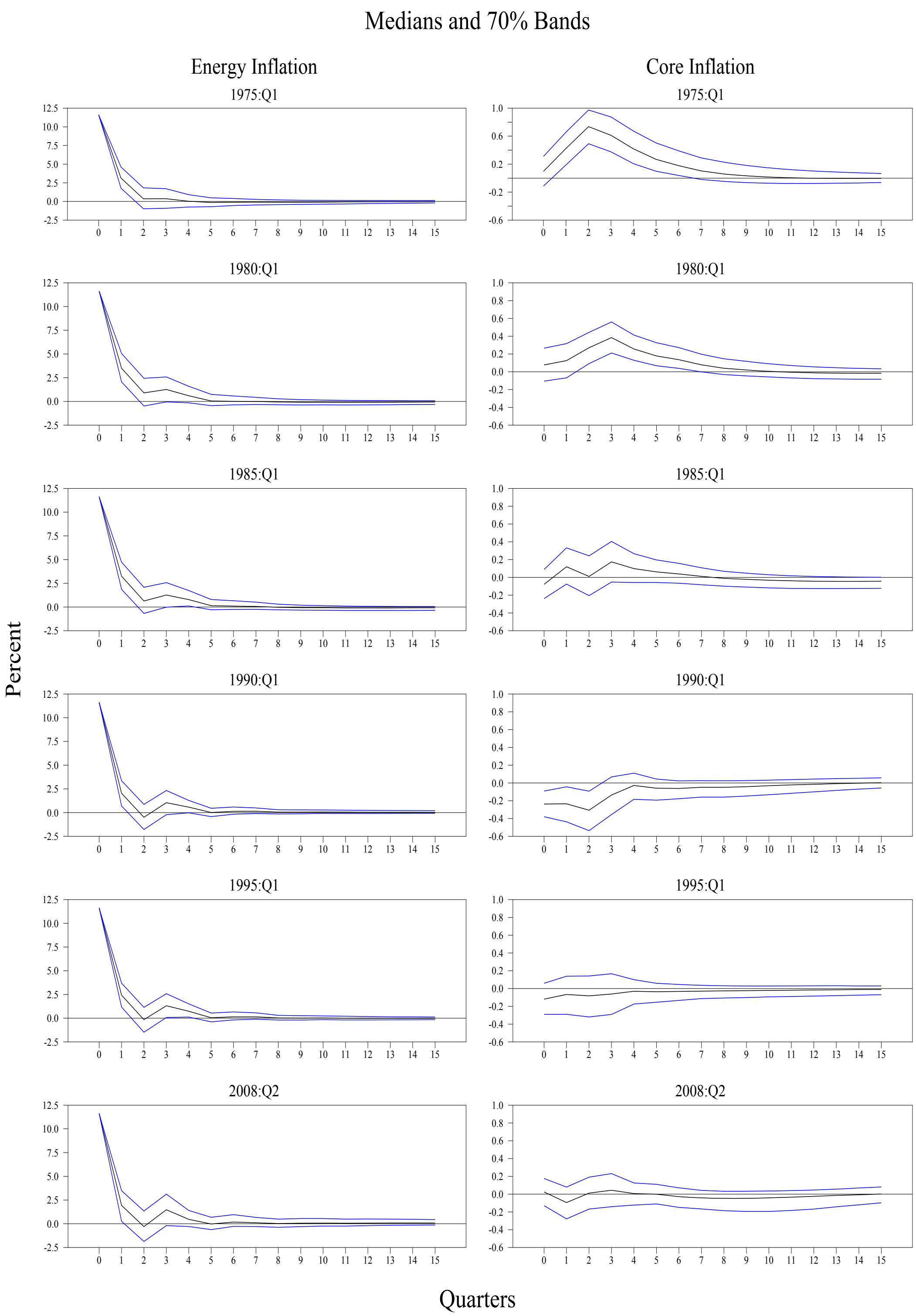

Figure 3: Historical Decompositions and Energy Inflation

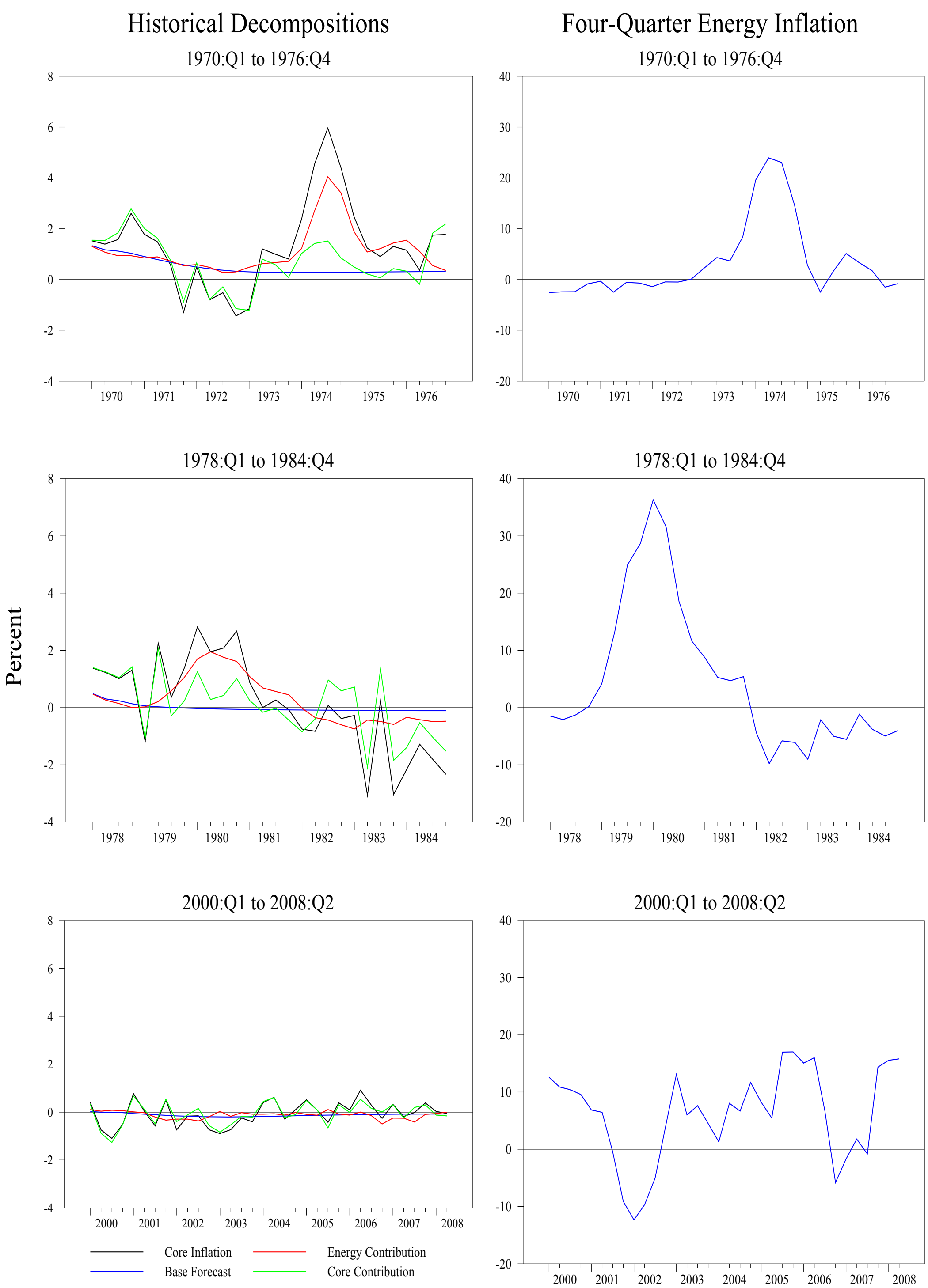

Figure 4: Residual Standard Deviations

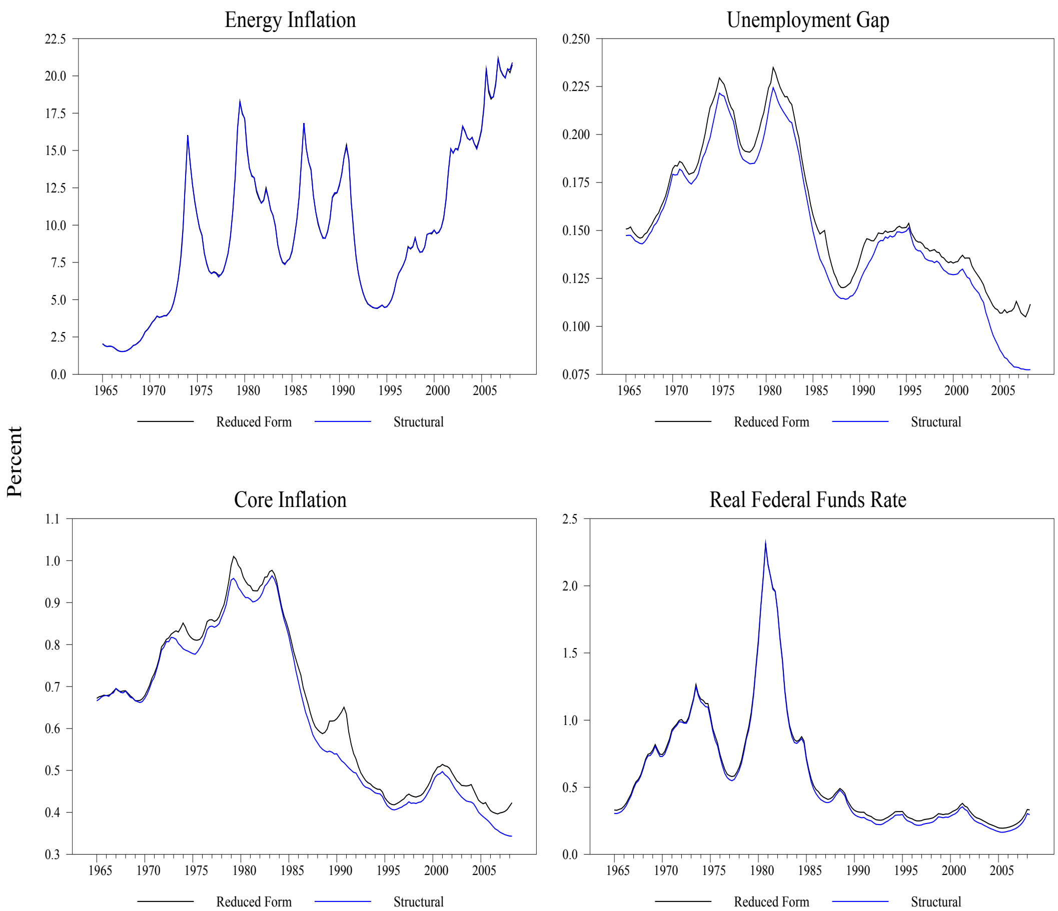

Figure 5: Sum of Coefficients on Energy Inflation in the Real Federal Funds Rate Equation

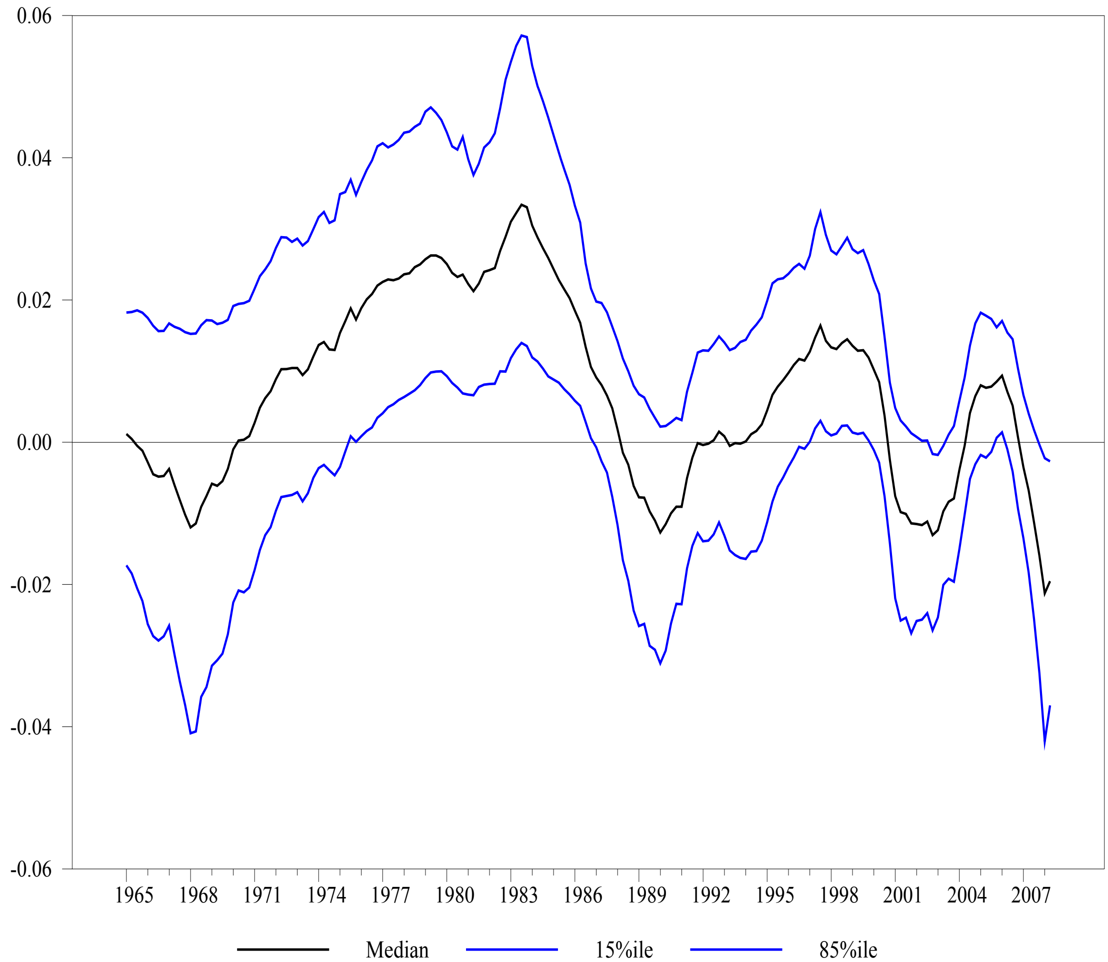

Figure 6: Impulse Responses to a Positive Energy Inflation Shock