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## Time Variation in the Inflation Passthrough of Energy Prices

From Bayesian estimates of a vector autoregression that 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 that became less responsive to energy inflation starting around 1985.

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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%, from around \$20 to almost \$150. Over the same period, average quarterly core inflation in the United States was 2%. 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 primarily focused on oil prices, including Hooker (2002), De Gregorio, Landerretche, and Neilson (2007), van den Noord and André (2007), Chen (2009), and Blanchard and Galí (2010), 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

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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 vector autoregression (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 (e.g., Gordon 1997, 1998, Brayton, Roberts, and Williams 1999, Chen 2009). Baumeister and Peersman (2009) use a very similar approach to disentangle the effects of oil supply and demand shocks on GDP growth and total (not core) inflation, in a model including oil's price and production level, GDP, and the CPI (but not a short-term interest rate).<sup>1</sup>

Our model estimates yield a pronounced reduction from approximately 1975 onward 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 U.S. economy and a recent prolonged increase in the volatility of energy price shocks. On the basis of structural VAR evidence, we also find that monetary policy has been less responsive to energy price shocks since approximately 1985.

The paper is organized as follows: Section 1 outlines the Bayesian methodology and data used in this study, Section 2 describes the reduction in passthrough of energy price inflation, Section 3 presents evidence of changing volatility in each equation of our VAR, Section 4 discusses changing monetary policy responsiveness to energy prices, and Section 5 concludes.

1. Peersman and Van Robays (2009) report estimates of the effects of oil supply shocks on overall, energy, and core CPI prices, based on a constant-parameter model estimated with data for 1986–2008.

## 1. 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:

$$y_t = X_t' B_t + A_t^{-1} \Sigma_t^{1/2} \epsilon_t,$$

$$X_t' = I_n \otimes [1, y_{t-1}', \dots, y_{t-k}'],$$

$$\text{Var}(\epsilon_t) = I_n,$$

with

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

The parameter vectors  $B_t$ ,  $a_{i,t}$ , and  $\sigma_t^2$  follow random walk processes:

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

$$a_{i,t} = a_{i,t-1} + v_{i,t},$$

$$\log \sigma_t^2 = \log \sigma_{t-1}^2 + e_{i,t},$$

where  $a_{i,t}$  is a vector containing all nonzero, non-one elements of the  $i$ th row of  $A_t$  and  $\log \sigma_t^2$  denotes a vector containing the logs of the diagonal elements of  $\Sigma_t$ . 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 U.S. 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 the specification for which we report results, real activity is measured by (annualized) real GDP growth. However, using an output gap, the unemployment rate, or an unemployment gap (as is more

typically the case in the Phillips curve literature mentioned earlier) yields the same basic results. Similarly, for simplicity in the presented results, we define the energy inflation variable in the model as just PCE energy inflation. But estimates based on the growth rate of the relative price of energy (energy price index/core PCE price index), weighted by energy's share in consumer spending, yields the same results.<sup>2</sup> In the literature on the economic effects of oil prices, transformations of energy price series incorporating a large number of possible nonlinearities have generated considerable interest (e.g., Hooker 1999, Hamilton 2003). However, the results in Edelstein and Kilian (2007, 2009) suggest that a linear and symmetric specification of energy price inflation is appropriate, and Hooker (2002) presents evidence that the declining impact of oil prices on core inflation in later years is robust to alternative nonlinear 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  $A_t$ ,  $\Sigma_t$ , and other parameters, the linearity of the model and Gaussian distribution 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.<sup>3</sup> Similarly, conditional on  $B_t$ ,  $\Sigma_t$ , 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  $\Sigma_t$  are then drawn conditionally for each variable using the Cogley-Sargent version of the Metropolis algorithm developed by Jacquier, Polson, and Rossi (1994). With  $B_t$ ,  $A_t$ , and  $\Sigma_t$  in hand, weakly informative natural conjugate priors yield tractable inverse-Wishart conditional posterior distributions for  $Q$ ,  $S_t$ , and  $Z$ . Prior distributions for the initial values  $B_0$ ,  $A_0$ , and  $\Sigma_0$  are specified using information from a training sample of 1950:Q1–1964:Q4.

Following Cogley and Sargent (2005), our posterior sampler truncates explosive draws of  $B_t$ . In the case of an explosive draw, we “backstep” until drawing stable coefficients. Following Del Negro (2003), to reduce the frequency of explosive draws we use an informative prior in obtaining estimates of  $B_0$  from the training sample.<sup>4</sup> We first perform 50,000 burn-in draws, followed by an additional 50,000

2. This weighted relative inflation rate corresponds to the percent change in purchasing power developed in Edelstein and Kilian (2009) and to the energy variable used in some Phillips curve work (e.g., Gordon 1997, 1998, Brayton, Roberts, and Williams 1999).

3. 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 0.001 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).

4. 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:Q1). Due to small market volume early in the training sample, for the 1949–59 period we measure the short-term interest rate with the 3-month Treasury bill rate instead of the federal funds rate.

draws, retaining every 10th draw. More complete details describing the methodology used here can be found in Cogley and Sargent (2005) and Primiceri (2005).

To determine the responses of core inflation to energy prices, we identify shocks to each variable recursively through the VAR ordering. Following studies, such as Burbidge and Harrison (1984), Davis and Haltiwanger (2001), Leduc and Sill (2004), and Blanchard and Galí (2010), energy price inflation is ordered first. This model treats energy prices as endogenous, consistent with the arguments of papers including Rotemberg and Woodford (1996) and Barsky and Kilian (2001). Following the guidance provided in Kilian (2008, p. 875), identification of energy price impacts is achieved by assuming the innovations to energy prices in the VAR to be exogenous with respect to all other current-period shocks (by placing energy prices first in the VAR). Evidence in Kilian and Vega (Forthcoming) on the responses of energy prices to economic news supports the identification of energy price shocks on the basis of such timing restrictions. However, in light of potential sensitivity of our results to the identification scheme, we have verified that our main results hold up under two alternative schemes: (i) a recursive ordering with energy inflation after core inflation and economic activity and (ii) sign-restriction-based identification.<sup>5</sup> As to the remainder of the model, core inflation is ordered before the economic activity variable (GDP growth in our presented results), following Primiceri (2005). Finally, incorporating a common 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.

In order to account for the uncertainty associated with parameter movements after a shock, our estimates of impulse responses are based on the generalized impulse response methodology of Koop, Pesaran, and Potter (1996). Following Benati (2008) and Baumeister and Peersman (2009), we obtain a sample of 1,000 generalized impulse response estimates—for which we report medians and 70% credible sets—from the following steps. (i) From the 5,000 retained posterior draws of time series of VAR coefficients and error variances (the elements of  $A_t$  and the structural shock variances), we sample a time series of coefficients and error variances. (ii) For each period for which we compute impulse responses, we simulate the future paths (16 periods total) of coefficients, error variances, and shocks—100 times. (iii) From these paths, we compute paths of the model variables, with a baseline path based on all shocks and an alternative path that adds in a base period structural shock of 1 to the variable of interest. The  $i$ th period response to the shock of interest is the average, across the 100 samples, of the difference between the alternative and baseline paths in period  $i$ .

5. For computational tractability, we simply generated sign-restriction-based results for constant-parameter VARs estimated separately for 1965–1984 and 1985–2008. The energy price shock was defined as one increasing the price of energy and reducing GDP. In these split-sample estimates, an energy price shock caused a significant rise in core inflation in the first sample and no significant change in core inflation in the second sample. Split-sample estimates based on a recursive ordering are qualitatively similar.

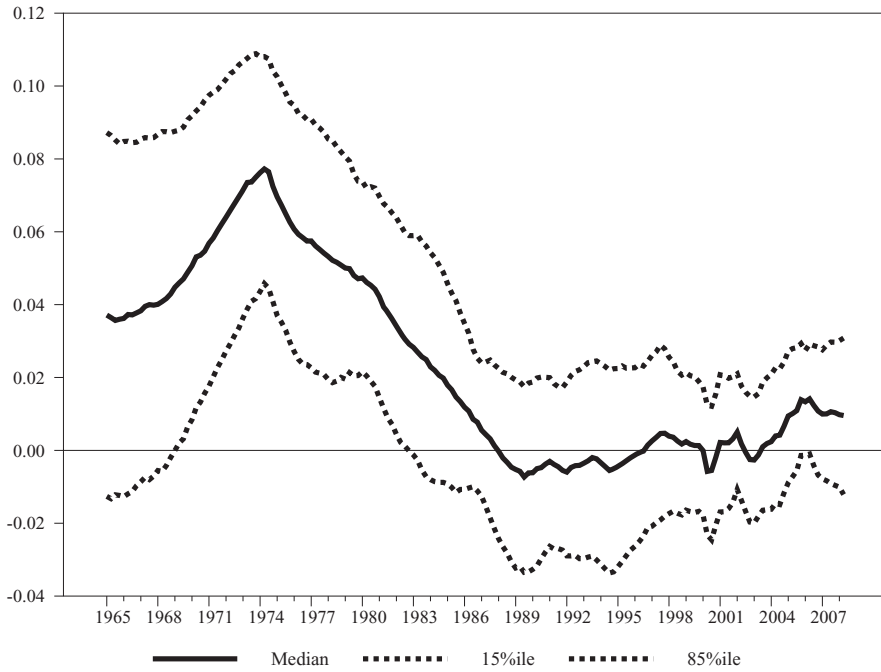


FIG. 1. Sum of Coefficients on Energy Inflation in the Core Inflation Equation.

## 2. 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 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 differences among historical decompositions of core inflation during three periods ranging from the early 1970s to the end of our sample.

### 2.1 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% 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 1982:Q4 the posterior bands include zero. From this date throughout the remainder of the sample, it is impossible to

differentiate the coefficient sum from zero based on the 70% credible sets. Given our Bayesian methodology, the posterior probability of a decline in the coefficient sum is simple to calculate and interpret (as in Cogley, Primiceri, and Sargent 2010). Evidence for a decline in passthrough during the 11 years from 1974:Q1 to 1985:Q1 is strong, with a posterior probability of 93.5%.<sup>6</sup> The same decline is evident in unreported estimates based on the relative price of energy weighted by energy's share in consumer spending, confirming that a large part of the estimated reduction in energy price passthrough occurs during a period of reduced energy consumption shares in the U.S. economy.<sup>7</sup>

## 2.2 Impulse Responses

Based on our recursive identification scheme and posterior samples, we calculate generalized impulse responses for an energy inflation shock at six dates (1975:Q1, 1980:Q1, 1985:Q1, 1990:Q1, 1995:Q1, and 2008:Q2). 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 12.2%. Figure 2 plots the median point estimates of the responses of both energy and core inflation, along with 70% 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 to 2008:Q2 the 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 (2010) describe a decline in the inflationary impact of oil price shocks in the United States and some European countries, and Herrera and Pesavento (2009) also uncover evidence of smaller impulse responses to oil price shocks in U.S. data.

## 2.3 Historical Decompositions

On the basis of our identification scheme for shocks to each variable, we estimate an historical decomposition of core inflation. Specifically, for the full sample, we decompose movements in inflation into contributions from each of the (orthogonalized) shocks in the model plus a baseline component that reflects initial conditions at the start of the sample and movement over time in the model's coefficients.<sup>8</sup> In the interest

6. 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, Landerretche, and Neilson (2007), van den Noord and André (2007), and Chen (2009) extend the result to a broad range of other countries.

7. As a nominal percentage of PCE, after peaking in 1981:Q2 at 9.3%, spending on energy goods and services fell quickly to a share of 7.6% only 3½ years later in 1985:Q1.

8. The basic methodology, for the simpler case of constant coefficients, is described in Burbidge and Harrison (1984).

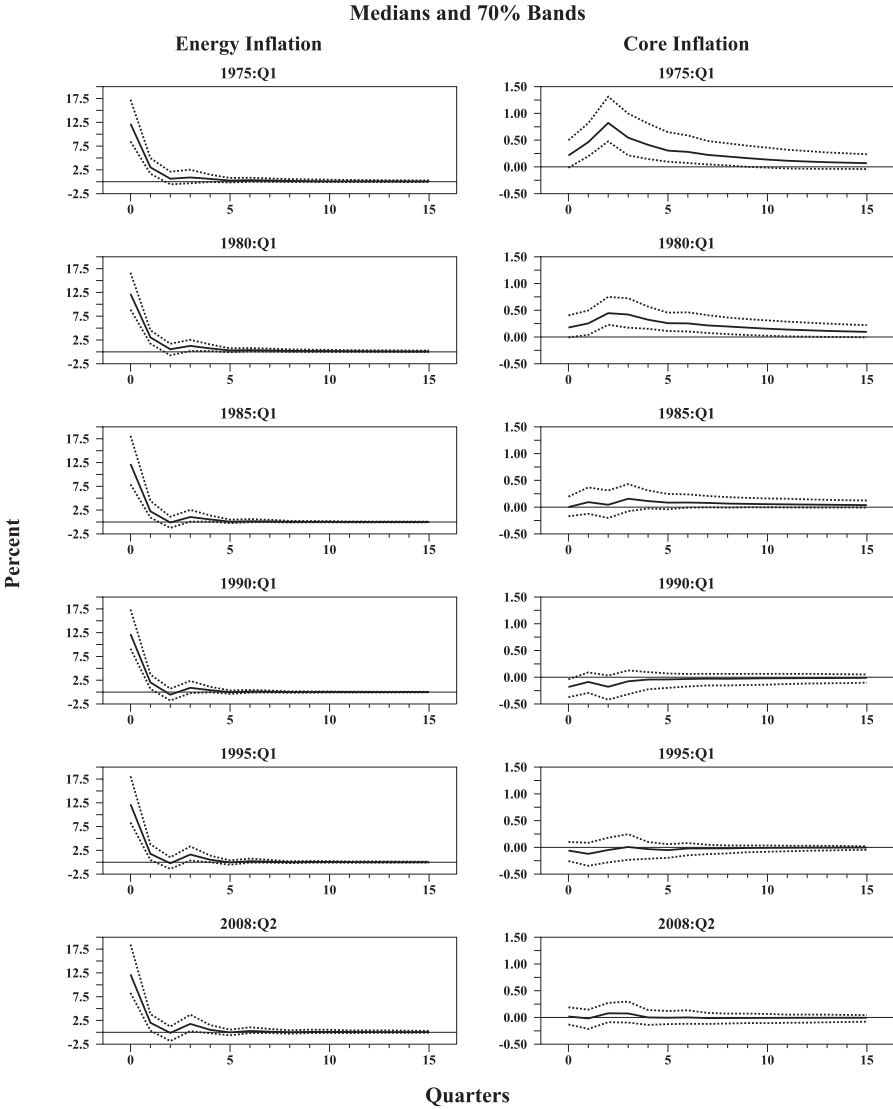


FIG. 2. Impulse Responses to a Positive Energy Inflation Shock.

of brevity and chart readability, we report in Figure 3 the estimated decomposition—actual core inflation against posterior means of the contributions from shocks to energy and core inflation—for the periods 1970:Q1–1976:Q4, 1978:Q1–1984:Q4, and 2000:Q1–2008:Q2.<sup>9</sup> These intervals were roughly chosen to include the October

9. Estimates of decompositions performed separately for each of these periods, based on initial conditions at the start of each subsample instead of at the start of the full sample, yield the same conclusions.



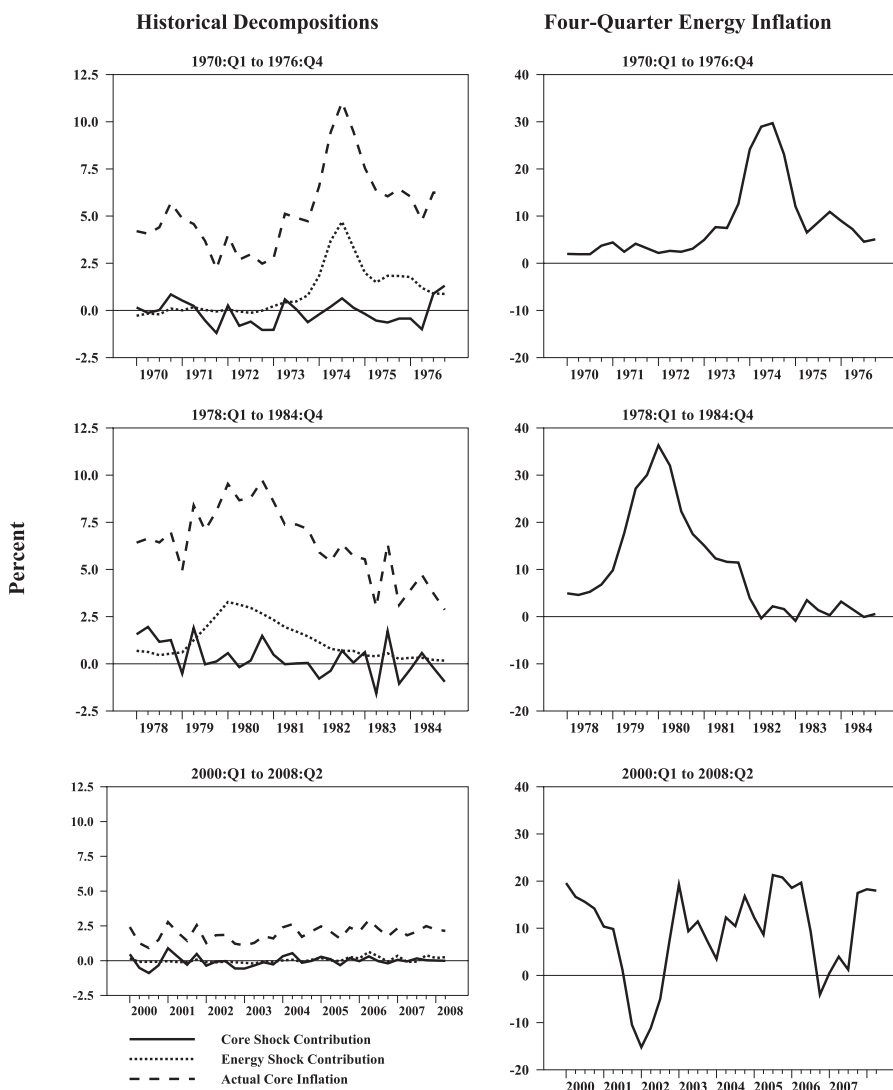


FIG. 3. Historical Decompositions and Energy Inflation.

War and Arab oil embargo (1973–74), the Iranian Revolution and subsequent outbreak of the Iran/Iraq war (1978–80), and the recent protracted period of increasing oil prices, respectively. Figure 3 also reports energy inflation in each of these samples.

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 relatively 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 to 2008:Q2, innovations to core inflation itself dominate the effects of energy price shocks. The sharp swings in energy prices in the last couple of years of the sample are estimated to have resulted in small movements in core inflation. Overall, these changes in decompositions are consistent with the evidence of drifting coefficients and declining impulse responses presented earlier.

### 3. 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 U.S. macroeconomic variables beginning in the early 1980s. Figure 4 plots estimates of the standard deviations of the orthogonalized (structural) innovations of our VAR. We report marked declines in the volatility of shocks to GDP growth, core inflation, and the federal funds rate from 1980 onward, 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.<sup>10</sup> From the mid-1990s onward, the volatility of innovations to energy inflation has trended sharply upward.<sup>11</sup>

### 4. 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. This section presents evidence that, from approximately 1985 onward, monetary policy has indeed become less responsive to energy prices. However, combined with the reduced energy passthrough documented earlier, 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.

Our generalized impulse response estimates indicate the response of the federal funds rate to an identified energy inflation shock changed around 1985. Figure 5

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

11. Based on split-sample estimates, Edelstein and Kilian (2009) report increasing volatility in the innovations of energy price shocks from 1988 onward.

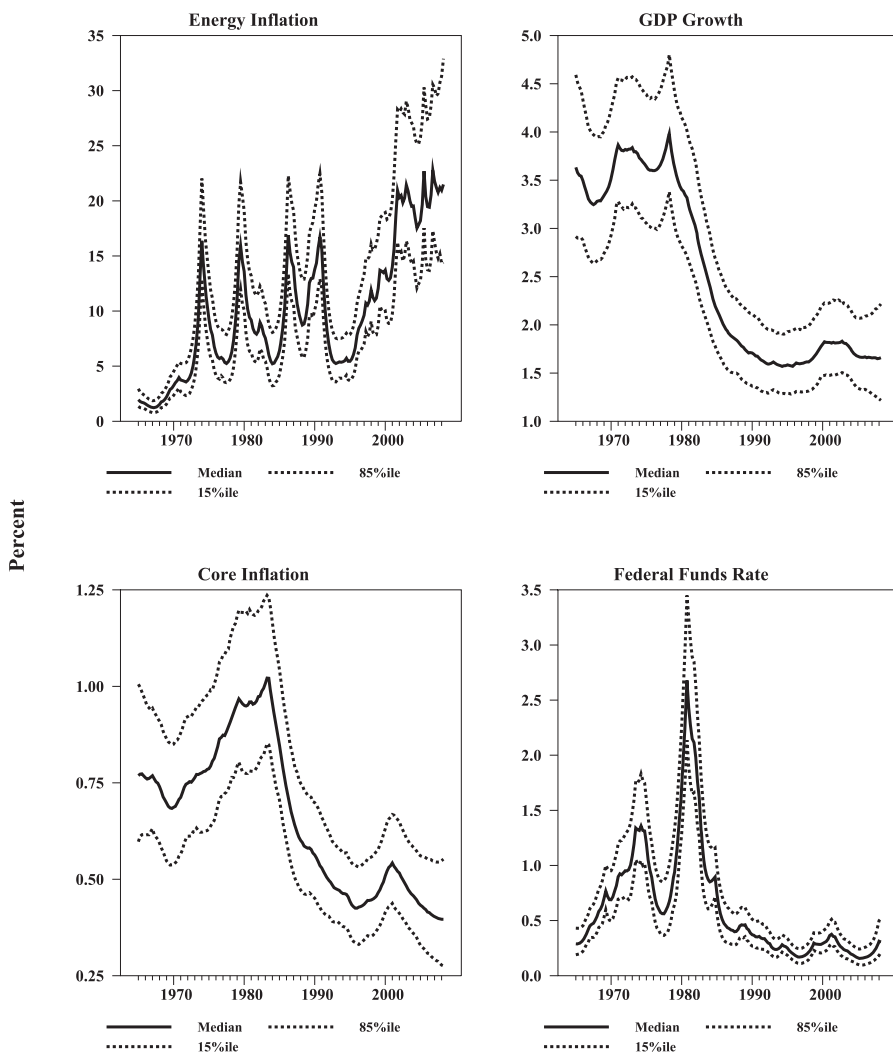


FIG. 4. Residual Standard Deviations (Structural Shocks).

plots posterior medians and 70% credible sets for the responses of the log of real GDP (obtained by cumulating each draw of the estimates of the response of GDP growth) and the 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 12.2%. In 1975:Q1 and 1980:Q1, the federal funds rate rises for three or four quarters after the shock, then gradually declines. This path is accompanied by a sharp, long-lasting decline in GDP. From 1990:Q1 onward, however, the response of the federal funds rate is either much smaller or negative (and indistinguishable from zero in either

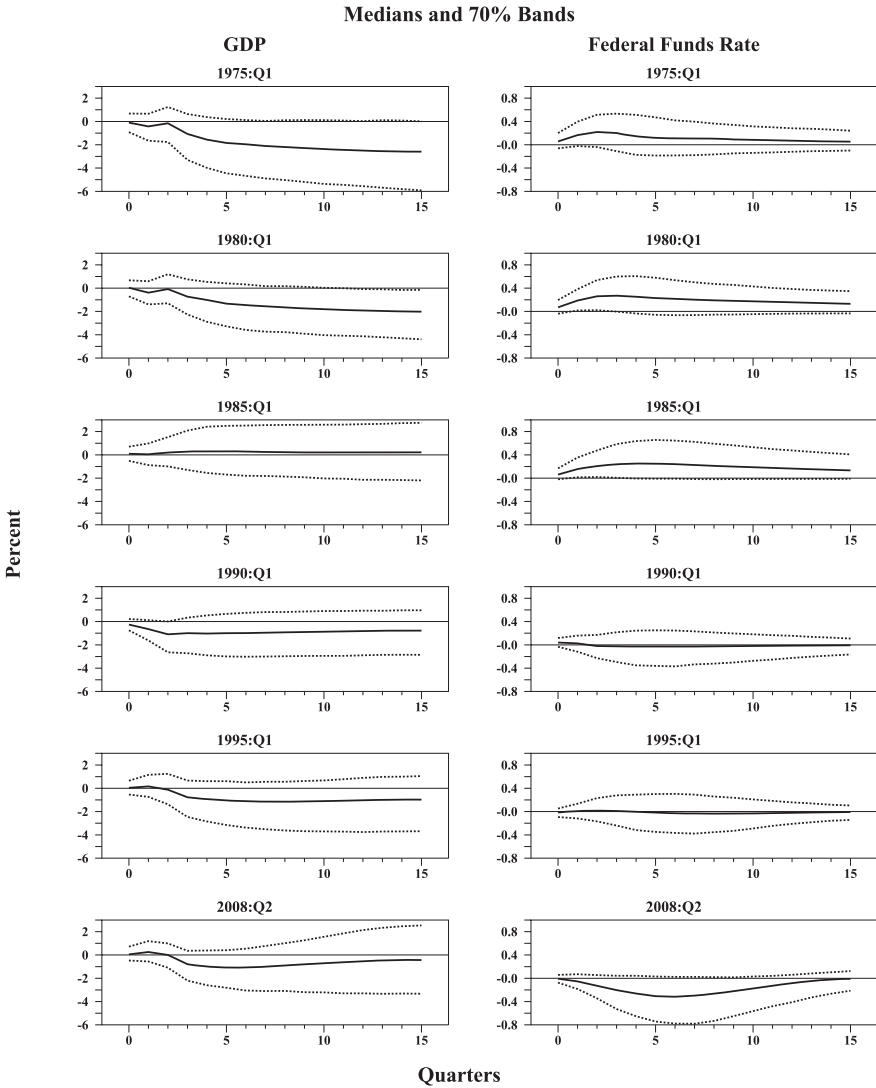


FIG. 5. Impulse Responses to a Positive Energy Inflation Shock.

case) in the face of an identified energy inflation shock. Our results are consistent with those in Bernanke, Gertler, and Watson (1997), Hooker (2002), and Herrera and Pesavento (2009), which document a more muted response of monetary policy to oil price shocks in recent decades.<sup>12</sup>

12. Our impulse response evidence does not 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, Gali, and Gertler (2000).

## 5. 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 with time-varying parameters and stochastic volatility. Our VAR generalizes some common reduced-form Phillips curves that include energy inflation (e.g., Gordon 1997, 1998, Brayton, Roberts, and Williams 1999, 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 changes among historical decompositions of core inflation. The responsiveness of monetary policy to energy inflation has also changed importantly over time. From approximately 1985 onward, identified shocks to energy prices have induced little movement in the federal funds rate.

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 U.S. 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, Landerretche, and Neilson (2007), Segal (2007), Kilian (2008), Chen (2009), and Blanchard and Galí (2010) for a more complete discussion of these alternative possibilities.

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