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# Commodity Price Shocks and the Business Cycle: Structural Evidence for the U.S.\*

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## Abstract

This paper evaluates the relative importance of commodity price shocks in the U.S. business cycle. Therefore, we extend the standard set of business cycle shocks to include unexpected changes in commodity prices. The resulting SVAR shows that commodity price shocks are a very important driving force of macroeconomic fluctuations — second only to investment-specific technology shocks — particularly with respect to inflation. Neutral technology shocks and monetary policy shocks, on the other hand, seem less relevant at business cycle frequencies. Neutral technology shocks rather play an important role at low frequencies.

**JEL Classifications:** C32, E32, E52, Q43

**Keywords:** business cycles, commodity price shocks, structural VAR

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# 1 Introduction

What are the sources of the U.S. business cycle? In recent years, the main body of business cycle research has focused on the impact of (neutral) technology shocks (Galí 1999), investment-specific technology shocks (Fisher 2006), and monetary policy shocks (Christiano et al. 1996). Following this strand of the literature, technology shocks, particularly investment-specific technology shocks, are considered to be a very important driving force of the business cycle.<sup>1</sup> On the other hand, Hamilton (2008) emphasizes that “nine out of ten” recessions in the postwar era are associated with a surge in oil prices. Given that oil is both a consumption good and an intermediate good, such a surge has a direct impact on the price level and changes the cost structure of firms (Medina & Soto 2005). Quite surprisingly, however, the macroeconomic effects of oil/commodity price shocks have been studied either in isolation (Edelstein & Kilian 2007, 2009, Blanchard & Galí 2010), or together with monetary policy shocks (Bernanke et al. 1997, Hamilton & Herrera 2004, Herrera & Pesavento 2009, Kilian & Lewis 2011), but never so far in conjunction with the standard set of business cycle shocks.

This paper bridges the gap between these two strands of the literature. Our main aim is to quantify the *relative importance* of commodity price shocks in the U.S. business cycle. Therefore, we develop a nine-dimensional SVAR, where the standard set of business cycle shocks (Altig et al. 2011) is extended to include unexpected changes in commodity prices.<sup>2</sup> The commodity price shock is identified by assuming that commodity prices are predetermined with respect to U.S. macroeconomic aggregates. This short-run restriction is based on the observation that energy prices do not respond immediately to macroeconomic news (Kilian & Vega 2011). As explained by Kilian (2008, p. 4), this assumption provides a “good approximation” when working with quarterly data. Furthermore, following Edelstein & Kilian (2007, 2009), we focus our analysis on the average effects of commodity price shocks, irrespective of whether these movements are driven by changes in supply or demand. We subject these identifying assumptions to a number of robustness checks (see below).

We find that that commodity price shocks are a very important driving force of the U.S. business cycle, second only to investment-specific technology shocks. In particular, we show that commodity price shocks explain a large share of cyclical movements in inflation. Results from a historical decomposition of shocks indicate that commodity price shocks have played a significant role especially during and after the the first OPEC oil crisis. Unexpected variations in the relative price of investment goods are the pri-

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<sup>1</sup>Evidence supporting this view includes SVAR studies (Ravn & Simonelli 2008, Canova et al. 2010, Altig et al. 2011) and estimated DSGE models (Gertler et al. 2008, Justiniano et al. 2010).

<sup>2</sup>We measure commodity prices using the index by the BLS (2012, see also Section 2.1). Given that this broad commodity price index is less prone to changes in institutional settings than the price of crude oil (see Figure 1), we are able to estimate the 9-dimensional SVAR model based on a long sample period.

mary determinant of business cycle fluctuations in output and per-capita hours. Neutral technology shocks and monetary policy shocks, on the other hand, seem less relevant at business cycle frequencies. Neutral technology shocks rather play an important role at low frequencies.

Furthermore, we demonstrate that an unexpected increase in *commodity prices* is characterized by significant U-shaped responses in output, consumption and per-capita hours. Most notably, the inflation rate displays a significant spike, followed by a rapid return to the initial level. The unexpected surge in the inflation rate prompts the Federal Reserve to elevate the nominal interest rate. Results of a counterfactual exercise indicate that the systematic contractionary response helped the Federal Reserve to achieve price stability in the long run, yet at the cost of a significant economic downturn in output and per-capita hours.<sup>3</sup>

Besides, we find that the estimated impulse response functions to *monetary policy* shocks, *neutral technology* shocks, and *investment-specific technology* shocks are very similar compared to those obtained by Altig et al. (2011) or Ravn & Simonelli (2008). In particular, the response of per-capita hours to neutral technology shocks is positive and marginally significant. This result is very robust, no matter whether the data are filtered or not, indicating that concerns about leaving (Fernald 2007, Francis & Ramey 2009, Canova et al. 2010) or removing (Gospodinov et al. 2011) low-frequency movements in the data are quantitatively not very important as long as the size of the information set is sufficiently large. On the other hand, if the information set is small, the impact response of per-capita hours is indeed extremely sensitive to the treatment of the data. This result, confirms our choice to estimate a large-scale SVAR.

We perform the following robustness checks to test the sensitivity of our results. First, we relax the contemporaneous exogeneity assumption by allowing for immediate responses in the commodity price index to innovations in U.S. aggregate activity; i.e., labor productivity growth and per-capita hours. Second, in order to control for contemporaneous movements in the global demand for commodities, we include a global demand indicator, which is ordered first before the commodity price index (as in Kilian & Lewis 2011). Furthermore, we examine robustness to the specific commodity price index used, the choice of the lag length, and the sub-sample properties of our model. We find that the effects of a commodity price shock on output and the inflation rate are milder in the post-Volcker period, but the impulse responses remain statistically significant at the 10% level. This result is consistent with previous estimates by Edelstein & Kilian (2009), Herrera & Pesavento (2009) or Blanchard & Galí (2010).

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<sup>3</sup>The counterfactual exercise presumes that the Federal Funds rate is kept constant when the U.S. economy is hit by a commodity price shock (as in Bernanke et al. 1997). However, our approach differs in many aspects from theirs as we adopt several suggestions — e.g. a linear VAR model (Kilian & Vigfusson 2011) with a lag length beyond one year (Hamilton & Herrera 2004) — made in the subsequent debate.

The remainder of this paper is organized as follows. Section (2) presents the identification and estimation strategies. Section (3) presents the results. Section (4) performs several robustness checks. Section (5) concludes.

## 2 Identification and Estimation Strategy

This paper evaluates the relative importance of commodity price shocks in the U.S. business cycle. Therefore, we extend the standard set of business cycle shocks (i.e., monetary policy, neutral technology, and investment-specific technology [Altig et al. 2011](#)) to include unexpected changes in commodity prices.

### 2.1 Data

We develop a nine-dimensional VAR in order to capture the impact of the relevant factors for macroeconomic fluctuations. The sample period covers aggregate U.S. data between 1955Q3 and 2007Q4.<sup>4</sup> The following variables enter the SVAR: growth in the relative price of investment goods  $\Delta q_t$ , growth in labor productivity  $\Delta a_t$  (measured by the ratio of real output to hours per capita in the business sector), the CPI inflation rate  $\pi_t$ , hours per capita  $h_t$ , the consumption share in output  $c_t$ , the investment share in output  $i_t$ , the employment rate  $n_t$ , the Federal Funds rate  $r_t$ , and the commodity price index “PPI: crude materials for further processing”  $p_t$  (see also [Figure 1](#)). We prefer to use this particular commodity price index by the [BLS \(2012, p. 8\)](#) — also used by [Hanson \(2004\)](#) or [Sims & Zha \(2006\)](#) — as it appropriately captures the time-varying importance of different raw materials, based on input-output studies by the BEA.<sup>5</sup> In addition, [Figure \(1\)](#) shows that, compared with the price of crude oil in the pre-1973 period, the pattern of changes in the broad commodity price index is much less discrete. All time series are seasonally adjusted (where applicable). Precise definitions can be found in the [Appendix \(Tables 1 and 2\)](#).

Note that our information set explicitly considers the consumption and the investment share in output. As demonstrated by [Christiano et al. \(2003\)](#), omitting these two variables may be associated with a serious specification error. This effect is likely due to the fact

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<sup>4</sup>The endpoint of our sample marks the start of the Great Recession when the Federal Reserve adopted several unconventional monetary policy measures, which are unlikely to be appropriately captured by our identification procedure.

<sup>5</sup>In contrast, the [Commodity Research Bureau \(2013\)](#) BLS Spot Index (available from 1951Q1) does not capture petroleum based products. The Continuous Commodity Index by [Thomson Reuters \(2013\)](#), calculated backwards until 1956Q4; i.e., the “old CRB”, captures petroleum based products, but continuously rebalances the different commodity categories to maintain an equal and time-invariant weight (we use this index for robustness checks presented in [Section 4.3](#) and [Figure 18](#)). The “new” [Thomson Reuters/Jefferies \(2013\)](#) CRB Index, introduced in 2005, captures petroleum based products and uses time-varying weights, but was calculated backwards only until 1994Q1 (see the link in the references). [Figure \(1\)](#) shows that the [BLS \(2012\)](#) index and the “new” [Thomson Reuters/Jefferies \(2013\)](#) CRB index behave remarkably similar in the overlapping sample.

that these two variables reflect “news that agents see but we do not” (Cochrane 1994, p. 295). In addition, we include the employment rate in order to analyze labor adjustment along the extensive and the intensive margin.

Recently, Fernald (2007) has shown that SVAR models with long-run restrictions may be biased by “low-frequency movements” in hours per capita. This effect can be attributed to sectoral changes involving government and non-profit employment or the movement of the baby boom generation through the labor market (Francis & Ramey 2009). Differencing removes the low-frequency movements from the data. However, differencing a bounded series (like per-capita hours) may involve misspecification issues (Hamilton 1994, p. 652). Therefore, Canova et al. (2010) suggest to filter the data prior to estimation. On the other hand, Gospodinov et al. (2011) argues that filtering the data prior to estimation removes information necessary to identify these shocks using long-run restrictions.

With this in mind, we estimate our SVAR model in two specifications; i.e., the “level specification” and the “bandpass filter specification”. For the reader’s convenience, we offer comparisons between these two specifications throughout the paper. For the former, we first take the natural logarithm of all variables except for the (net) Federal Funds rate, and then difference labor productivity and the relative price of investment goods. For the latter, we additionally apply a one-sided bandpass filter (Christiano & Fitzgerald 2003)<sup>6</sup> in order to control for low-frequency movements in per-capita hours. We prefer this particular filter since agents know only the past (Lucas 1980). More precisely, we apply the one-sided bandpass filter not only to per-capita hours, but to *all* series considered.<sup>7</sup> Figure (2) illustrates that this procedure allows us to maintain spectral coherence (Granger 1969) between labor productivity growth and per-capita hours. When all series are filtered — as in the top panel — we are able to break the low-frequency comovement and minimize distortions at higher (particularly, at business cycle) frequencies. When only per-capita hours are filtered — as in the bottom panel — we are less successful in breaking the low-frequency comovement and distort the relationship at business cycle frequencies.

## 2.2 Identification

We estimate the following four structural shocks using standard identifying assumptions. Commodity price shocks (Rotemberg & Woodford 1996) and monetary policy shocks (Christiano et al. 1996) are identified using short-run restrictions. Neutral and investment-

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<sup>6</sup>To be precise, we first remove the drift of the series and then apply the one-sided bandpass filter with following options:  $p_l = 2$ ,  $p_u = 52$ ,  $root = 1$ ,  $drift = 0$ ,  $ifilt = 0$ ,  $nfix = -1$ ,  $thet = 1$ . Where available, we use data from 1948Q1 to 2007Q4 and then drop the filtered data points prior to 1955Q3. The Federal Funds rate is only available from 1954Q3.

<sup>7</sup>Our long-run identification strategy (see Section 2.2) requires that labor productivity and the relative price of investment goods enter the SVAR in first differences. Hence, we difference these variables first and then apply the bandpass filter.

specific technology shocks are identified using long-run restrictions (Galí 1999, Fisher 2006).<sup>8</sup> The remaining shocks in the nine-dimensional SVAR model are identified via a recursive ordering scheme.

Consequently, the reduced-form VAR is given by:

$$\begin{aligned} x_t &= a + B(L)x_{t-1} + e_t & (1) \\ x_t &= \begin{bmatrix} \Delta q_t & \Delta a_t & z_t & r_t & p_t \end{bmatrix}' \\ z_t &= \begin{bmatrix} \pi_t & h_t & c_t & i_t & n_t \end{bmatrix}' \end{aligned}$$

where  $B(L)$  is a lag polynomial of order  $M$ . By premultiplying with  $\beta_0$ , we obtain the structural VAR:

$$\beta_0 x_t = \alpha + \beta(L)x_{t-1} + \epsilon_t \quad (2)$$

where  $\epsilon_t$  denotes the vector of fundamental shocks. The orthogonality assumption implies that its covariance matrix  $V_\epsilon = E(\epsilon_t' \epsilon_t)$  is diagonal. Moreover, we normalize the diagonal of  $\beta_0$  to a 9x1 vector of ones.

Both technology shocks are identified using long-run restrictions (Shapiro & Watson 1988, Blanchard & Quah 1989). Following Fisher (2006), we assume that *only* investment-specific technology shocks affect the relative price of investment goods in the long run. The long-run level of aggregate productivity may be affected by both investment-specific *and* neutral technology shocks. No other shock has any long-run effect on the relative price of investment goods or the level of labor productivity (Galí 1999).

The identification strategies of the commodity price shock and the monetary policy shock are based on short-run restrictions. We impose the constraint that no other variable may respond contemporaneously when the Fed's monetary policy — given by the Federal Funds rate — deviates from its linear rule. This presumes that, when setting the nominal interest rate, the Fed's information set includes the contemporaneous values of all other variables included in the SVAR (Christiano et al. 1996). Moreover, we identify the commodity price shock by assuming that nominal commodity prices are predetermined with respect to U.S. macroeconomic aggregates.<sup>9</sup> This identification strategy was originally developed by Rotemberg & Woodford (1996) in the context of nominal oil price shocks.<sup>10</sup> The assumption of predeterminedness is based on the observation that energy prices do not respond contemporaneously to macroeconomic news (Kilian & Vega 2011). As ex-

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<sup>8</sup>The current section follows largely the identification procedure described in Ravn & Simonelli (2008), with commodity prices instead of government spending. We are unable to find significant movement to government spending shocks when we include government spending per capita instead of commodity prices in our SVAR model

<sup>9</sup>In contrast, innovations in *real* commodity prices may be due to unexpected changes in U.S. inflation and, therefore, should be treated as endogenous (Rotemberg & Woodford 1996, Hamilton 2008). See Kilian & Lewis (2011) for a recent work on this issue.

<sup>10</sup>The shape of the impulse responses remains unchanged when we use the West Texas Intermediate spot oil price instead, but the broad commodity price index turns out to be quantitatively more important.



plained by Kilian (2008, p. 4), this short-run restriction provides a “good approximation” when working with quarterly data.<sup>11</sup> Nevertheless, when applying this approach to commodity prices, we take into account the possibility that the broad commodity price index may behave less sluggishly than the nominal oil price (Alquist et al. 2011). Moreover, following Edelstein & Kilian (2007, 2009), we focus our analysis on the average effects of commodity price shocks, irrespective of whether these movements are driven by changes in supply or demand (see Kilian 2009, for a detailed discussion of this topic). We subject these identifying assumptions to the following robustness checks. Therefore, Section (4.1) relaxes the contemporaneous exogeneity assumption to allow for immediate responses in the commodity price index to innovations in U.S. aggregate activity; i.e., labor productivity growth and per-capita hours. Section (4.2) includes a global demand indicator — which is ordered first before the commodity price index (as in Kilian & Lewis 2011) — in order to control for contemporaneous movements in the global demand for commodities.<sup>12</sup> Besides, note that our linear VAR model presumes that commodity price increases and decreases have symmetric effects (Kilian & Vigfusson 2011).

Consequently, the process for the Federal Funds rate depends on the current and past values of all other variables, but no other process depends on its current realizations. This implies that the second-last column of the contemporaneous coefficient matrix  $\beta_0$  consists of zeros, apart from the second-last element which is normalized to unity. The process for the commodity price, on the other hand, depends on the lagged values of commodity prices and all other variables, but not on the current realizations of any other variable. Hence, the last row of  $\beta_0$  consists of zeros, apart from the last element which is normalized to unity. Furthermore, the order of the variables included in the vector  $z_t$  imposes a number of additional short-run restrictions on  $\beta_0$ .

## 2.3 Estimation

The first equation of the structural VAR (equation 2):

$$p_t = \alpha^p + \sum_{j=1}^M \beta_{x,j}^p x_{t-j} + \epsilon_t^p \quad (3)$$

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<sup>11</sup>Applications of this identification strategy to quarterly data include Rotemberg & Woodford (1996), Edelstein & Kilian (2007), and Blanchard & Galí (2010).

<sup>12</sup>Note that, given that commodity prices are denominated in U.S. dollars, we are not able to identify endogenous (real) exchange rate driven commodity price movements.

identifies the commodity price shock  $\epsilon_t^p$ . We estimate equation (3) using ordinary least squares. The second equation of the SVAR:

$$\begin{aligned} \Delta q_t = & \alpha^q + \sum_{j=1}^M \beta_{q,j}^q \Delta q_{t-j} + \sum_{j=0}^{M-1} \beta_{a,j}^q \Delta^2 a_{t-j} \\ & + \sum_{j=0}^{M-1} \beta_{z,j}^q \Delta z_{t-j} + \sum_{j=1}^{M-1} \beta_{r,j}^q \Delta r_{t-j} + \sum_{j=0}^{M-1} \beta_{p,j}^q \Delta p_{t-j} + \epsilon_t^q \end{aligned} \quad (4)$$

identifies the investment-specific technology shock  $\epsilon_t^q$ . The long-run restriction is imposed by differencing all the regressors in  $x_t$  apart from the relative investment goods price itself (note that  $\Delta^2$  is the second difference operator). Moreover, we exclude the contemporaneous value of the Federal Funds rate from this regression. This implements the short-run assumption on the Fed's information set. Since  $\epsilon_t^q$  may be correlated with  $\Delta a_t$  (via equation 5) and  $z_t$  (via equation 7), we estimate equation (4) with 2SLS. The instruments are a constant, the vector  $[\Delta q_{t-j}, \Delta a_{t-j}, z_{t-j}, r_{t-j}, p_{t-j}]_{j=1}^M$  and  $\hat{\epsilon}_t^p$  (the estimate of  $\epsilon_t^p$ ). The third equation of the SVAR:

$$\begin{aligned} \Delta a_t = & \alpha^a + \sum_{j=0}^M \beta_{q,j}^a \Delta q_{t-j} + \sum_{j=1}^M \beta_{a,j}^a \Delta a_{t-j} \\ & + \sum_{j=0}^{M-1} \beta_{z,j}^a \Delta z_{t-j} + \sum_{j=1}^{M-1} \beta_{r,j}^a \Delta r_{t-j} + \sum_{j=0}^{M-1} \beta_{p,j}^a \Delta p_{t-j} + \epsilon_t^a \end{aligned} \quad (5)$$

identifies the neutral technology shock  $\epsilon_t^a$ . Note that we difference all regressors — except for  $\Delta q_t$  and  $\Delta a_t$  — and exclude the contemporaneous value of the Federal Funds rate. We estimate equation (5) using 2SLS, given that  $\epsilon_t^a$  may depend on  $z_t$  (via equation 7) and  $q_t$  (via equation 4). The instruments employed above are extended to include the estimate of  $\epsilon_t^a$ ; i.e.,  $\hat{\epsilon}_t^a$ . The fourth equation of the SVAR:

$$r_t = \alpha^r - \beta_{q,0}^r \Delta q_t - \beta_{a,0}^r \Delta a_t - \beta_{z,0}^r z_t - \beta_{p,0}^r p_t + \sum_{j=1}^M \beta_{x,j}^r x_{t-j} + \epsilon_t^r \quad (6)$$

identifies the monetary policy shock  $\epsilon_t^r$ . This equation is estimated with ordinary least squares.

Following Altig et al. (2011), we estimate the remaining parameters for the vector  $z_t$ . The components of  $z_t$  are denoted by  $z_t^i$ ,  $i = 1, \dots, 5$ . The parameters of the first equation are obtained by estimating:

$$\begin{aligned} z_t^1 = & \alpha^1 + \sum_{j=0}^M \beta_{q,j}^1 \Delta q_{t-j} + \sum_{j=0}^M \beta_{a,j}^1 \Delta a_{t-j} \\ & + \sum_{j=1}^M \beta_{z,j}^1 z_{t-j} + \sum_{j=1}^M \beta_{r,j}^1 r_{t-j} + \sum_{j=0}^M \beta_{p,j}^1 p_{t-j} + \epsilon_t^1 \end{aligned} \quad (7)$$

employing the above-used instruments including the vector of estimated shocks  $[\hat{\epsilon}_t^p, \hat{\epsilon}_t^q, \hat{\epsilon}_t^a]'$ . The second equation extends the set of regressors with  $z_t^1$  and the list of instruments with  $\hat{\epsilon}_t^1$ . We continue this procedure recursively for all the variables included in  $z_t$ .

In order to determine the optimal VAR order ( $M$ ), we apply the standard sequential likelihood ratio test (see e.g. Lütkepohl 2005, Chapter 4.2.2), which rejects  $M = 4$  at the 1% significance level. The choice to use a lag length beyond one year is also supported by the existing literature on energy price shocks (Hamilton & Herrera 2004). In addition, we test the null hypothesis of zero serial correlation using bootstrapped multivariate Portmanteau (Q) statistics (Altig et al. 2011). On the basis of this test, we do not reject the null hypothesis for  $M = 5$ .

## 3 Results

### 3.1 Dynamic Responses to Structural Shocks

We modify the code by Altig et al. (2011) to estimate the coefficients and to compute the impulse responses to the four identified structural shocks.<sup>13</sup> We examine the impulse response functions at horizons up to 32 quarters. The graphs depict the responses based on bootstrap sampling over 3,000 replications, where the first 1,000 draws are used to correct for small sample bias and departures from non-normality (Kilian 1998a,b).<sup>14</sup> The solid line is the median estimate. The gray shaded areas represent the associated 60%, 70%, 80% and 90% non-centered confidence intervals. For the reader's convenience, Figures (3)-(5), (7) in the Appendix contrast the impulse responses of the "bandpass filter specification" (a panels) with the impulse responses of the "level specification" (b panels).

#### 3.1.1 Commodity Price Shocks

Figure (3) depicts the impulse responses to the identified commodity price shock. We find that this shock triggers a temporary rise in the commodity price index, peaking shortly after the initial increase before slowly returning to its steady state level. Moreover, we observe a spike in the inflation rate, indicating that aggregate consumer prices are very flexible in response to commodity price shocks. In the following periods, the inflation rate declines sharply. The unexpected surge in the inflation rate prompts the Fed to elevate the nominal interest rate for a protracted period (about 6-8 quarters).<sup>15</sup> Consequently, the inflation rate falls below normal about two years after the shock. We also note that the relative price of investment goods decreases slightly, but the effect disappears

<sup>13</sup>We thank Lawrence Christiano for making the code available on his website.

<sup>14</sup>The Jarque-Bera test statistics reject the null hypothesis that the commodity price shocks and the monetary policy shocks are normally distributed at the 1% significance level.

<sup>15</sup>The peak response of the Federal Funds rate corresponds to 30 basis points. We infer this value from the *level* specification (see Figure 3, panel b).

relatively quickly. The adjustment paths of output, per-capita hours, employment, hours per worker, consumption, and investment display significant U-shaped responses. The estimated impulse responses of output and employment are qualitatively consistent with the results of Blanchard & Galí (2010) — output and employment decline persistently after a lag of 3-5 quarters and reach a trough after about ten quarters.

In this context, Bernanke et al. (1997) argue that a substantial part of the recessionary effects of commodity price shocks is not due to the direct impact of higher producer prices, but rather due to the systematic contractionary response of the Federal Reserve.<sup>16</sup> Their conclusion stems from a counterfactual exercise (suggested by Sims & Zha 2006) which presumes that the Federal Funds rate is kept constant when the U.S. economy is hit by an unexpected increase in commodity prices. In the following, we perform the same counterfactual exercise using our estimated SVAR model. Importantly, our approach differs in many aspects from theirs as we adopt several suggestions — e.g. a linear SVAR model (Kilian & Vigfusson 2011) with a lag length beyond one year (Hamilton & Herrera 2004) — made in the subsequent debate. We observe that the imputed counterfactual movements in the Federal Funds rate deviate only moderately from the original series (see Figure 8). For this reason, we believe that our results are less prone to changing parameters due to the Lucas (1976) critique.

Figure (9) contrasts the impulse response functions of output, per-capita hours and the inflation rate in the bandpass filter specification (top panel) with the impulse responses under the counterfactual assumption (bottom panel). Indeed, we are unable to observe a significant downturn in output and per-capita hours if the Fed stayed passive. There is only an insignificant decline in output that occurs with a lag of about two years. The *initial* spike in the inflation rate, on the other hand, seems identical to the one estimated in the bandpass filter specification. *At medium horizons* (10-20 quarters), however, the counterfactual response cannot replicate the significant disinflationary rebound in CPI inflation. Thus, we conclude that the contractionary monetary policy feedback rule helped the Federal Reserve to achieve price stability in the *long run*, yet at the cost of a significant economic downturn in output and per-capita hours.<sup>17</sup>

In addition, Figure (10) illustrates the impulse responses of two CPI sub-indices; i.e. the so-called core inflation rate (all items less food and energy) and its counterpart (food and energy only). We observe that the spike in the headline inflation rate is mainly due to a sharp rise in food and energy prices. The core inflation rate, on the other hand, shows

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<sup>16</sup>This view did not remain unchallenged. Herrera & Pesavento (2009, p. 107), on the other hand, find that “the systematic monetary policy response dampened fluctuations in economic activity during the 1970s”. Furthermore, Barsky & Kilian (2002) argue that shifts in the monetary policy regime following the breakdown of the Bretton Woods system triggered the Great Stagflation of the 1970s.

<sup>17</sup>Kilian & Lewis (2011), on the other hand, perform an alternative counterfactual exercise which assumes that the Fed does not respond directly to oil price shocks, but to the movements in other macroeconomic aggregates triggered by these shocks. They conclude that monetary policy responses did not cause large fluctuations in U.S. output.

a lower — but still significant — and more persistent increase. This indicates that a little price rigidity at the level of intermediate goods may translate into persistent inflation movements in other sectors of the economy (Basu 1995) — so-called second-round effects. We also find a marginally significant disinflationary rebound in both CPI sub-indices at medium horizons. Moreover, by repeating the above-described counterfactual exercise, we notice that the initial increase in both sub-indices remained virtually unchanged if the Federal Reserve stayed passive. The disinflationary rebound, however, disappears in both impulse responses. Therefore, we conclude that the Fed’s contractionary monetary policy feedback rule is unable to avoid second-round effects in the *short run*. Yet, it exhibits medium-run disinflationary effects which help the Federal Reserve to achieve price stability at *longer forecast horizons*.

### 3.1.2 Monetary Policy Shocks

Figure (4) shows the responses to an expansionary shock in monetary policy. This shock represents a drop in the Federal Funds rate, due to an unexpected deviation from the Fed’s linear policy rule. Our identifying assumptions imply that the shock has only a temporary effect. Nevertheless, the Federal Funds rate remains below its steady state level for more than seven quarters. In response to this, we observe that output, per-capita hours, employment, consumption, and investment rise gradually. Peak effects take place about 5-6 quarters after the monetary stimulus. At longer forecast horizons, the adjustment paths show a slight rebound. The response of the relative price of investment goods, on the other hand, is not significant. Overall, the shapes and elasticities of the responses are in line with the estimates by Ravn & Simonelli (2008). Only labor input indicators behave slightly different. Employment seems somewhat less elastic. Hours per worker even display a very mild downturn. Consistent with Sims (1992), the impulse response of the inflation rate drops on impact, followed by a slow and persistent increase. According to our estimates, the inclusion of the commodity price index reduces the size of the drop slightly, but leaves the shape of the inflation response unchanged. This indicates that the “price puzzle” is a robust feature of the data (Hanson 2004).

Furthermore, we observe that an unexpected cut in the Federal Funds rate induces a slow, but persistent increase in commodity prices. The maximum impact does not occur until four to five years after the shock. In comparison to Anzuini et al. (2012), our estimated impulse response is much more gradual and resembles (qualitatively as well as quantitatively) the impulse response of the consumer prices index (i.e., the cumulative response of the inflation rate). In other words, the commodity price index shows no significant response in *real* terms.

### 3.1.3 Neutral Technology Shocks

Figure (5) illustrates the impulse response functions to the identified neutral technology shock. We observe that a permanent improvement in labor productivity induces a long-lasting rise in output and consumption. On impact, both variables jump up and then remain well above their original value for the entire time horizon. Moreover, the shock produces a large and protracted hump-shaped response in investment. The inflation rate falls on impact and then asymptotes to its steady-state level within four years. There is also a modest increase in the relative price of investment goods, but the effect disappears relatively quickly. The impulse response of per-capita hours is positive and marginally significant at the 10% level. A very similar response can be observed for the employment rate. Hours per worker, on the other hand, rise on impact and then slowly return to their steady state. Quantitatively, however, the impact of the intensive margin is small. The estimated impulse responses differ only in one important respect from those obtained by Altig et al. (2011). We find that the increase in consumption is not gradual, but rather abrupt.

The left panel of Figure (6) documents that the response of hours worked to neutral technology shocks is extremely sensitive to the treatment of the data<sup>18</sup> when the information set is reduced to three variables  $\{q_t, a_t, h_t\}$  following the set-up in Canova et al. (2010). If we remove the low-frequency movements — either by applying the one-sided bandpass filter, by taking first differences, by including a time trend and two structural breaks in level and trend, or by including the corresponding Francis & Ramey (2009) hours time series in the level specification — the estimated hours response is significantly *negative*. This result is consistent with the findings of Galí (1999) and Canova et al. (2010). Instead, if per-capita hours enter the VAR in levels, the response is significantly *positive* (Christiano et al. 2003, 2004).

The right panel of Figure (6) shows the response of hours worked when the information set is large. We observe that the dynamic response is positive and marginally significant at the 10% level across all specifications. Only the dummy specification predicts a negative response during the first few quarters, but the confidence intervals are wide. Interestingly, we are also able to replicate the counterfactual exercise conducted by Fernald (2007). When all high and medium frequencies in per-capita hours are reversed, the small-scale SVAR model predicts a significantly positive response. This outcome is driven by the distortionary low-frequency movements. The counterfactual large-scale SVAR model, on the other hand, predicts that the hours response flips horizontally. This result is consistent with both the bandpass filter specification and the level specification. Thus, we conclude

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<sup>18</sup>*Difference* specification: Like the level specification, but also per-capita hours enter in first differences. *Dummy* specification: We extend the level specification to include a time trend and two structural breaks in level and trend at the dates 1973Q2 and 1997Q2 (see Fernald 2007). The Francis & Ramey (2009) hours time series is taken from Valerie A. Ramey’s website, which is gratefully acknowledged.

that the low-frequency bias (present in the level specification) and the misspecification error (induced by overdifferencing) become much less important when the information set is sufficiently large.<sup>19</sup>

### 3.1.4 Investment-Specific Technology Shocks

Figure (7) displays the effects of an investment-specific technology shock. This shock leads to an unexpected and permanent drop in the relative price of investment goods. We observe that all variables (except labor productivity) move together in response to this type of disturbance. Consistent with [Altig et al. \(2011\)](#), we find that the dynamic adjustment paths show a marked hump-shaped pattern, with peak effects occurring after 3-4 quarters. The impulse response of labor productivity remains insignificant for more than four years before eventually rising. This result illustrates that, on impact, the elasticity of per-capita hours is of the same magnitude as aggregate output (also here, most variation in labor input is due to adjustments along the extensive margin). Thus, investment-specific technology shocks seem far more important for the cyclical behavior of the labor market than neutral technology shocks.

## 3.2 Importance of the Structural Shocks

We now examine the relative importance of the four identified structural shocks for the variance of all variables included in our SVAR. First, we present the share of variation explained by each identified shock at different forecast horizons. However, as explained by [Ravn & Simonelli \(2008\)](#), these figures do not allow us to draw direct conclusions about the importance of these shocks at business cycle frequencies. Therefore, we also compute the variance decomposition at business cycle frequencies (8-32 quarters) following the method proposed by [Altig et al. \(2011\)](#). In addition, we present the historical decomposition of the four identified structural shocks for aggregate output in the postwar period.

### 3.2.1 Forecast Error Variance Decomposition

Figure (11) displays the forecast error variance decomposition of three key macroeconomic variables (output, per-capita hours, and the inflation rate) at different horizons. We observe that neutral technology shocks explain a large share of the variation in output, particularly at long forecast horizons. Investment-specific technology shocks are the main determinant of fluctuations in per-capita hours, and the second most important determinant of fluctuations in output. Commodity price shocks (together with neutral

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<sup>19</sup>This conclusion is in line with the results of [Forni & Gambetti \(2011\)](#). They demonstrate that trivariate SVAR models do not capture sufficient information in order to find an unbiased estimate of the hours response to neutral technology shocks. Our choice to use a large information set is also supported by the outcome of cross-correlation tests (see Table 3).

technology shocks) appear to be the primary driving force for movements in the inflation rate. Monetary policy shocks, on the other hand, explain only small shares of macroeconomic fluctuations. The joint explanatory power of all four shocks lies between 35% (inflation rate) and 48% (output) in the short run, and between 48% (inflation rate) and 68% (output) in the long run.

### 3.2.2 Variance Decomposition at Business Cycle Frequencies

We now investigate the variance decomposition at business cycle frequencies (see Table 4). The results show that commodity price shocks are a principal driving force for macroeconomic fluctuations. In particular, we find that commodity price shocks explain a large share of cyclical movements in inflation.<sup>20</sup> The commodity price shock also turns out to be a very important determinant of cyclical fluctuations in many other macroeconomic variables (e.g., Federal Funds rate, investment, or consumption), second only to investment-specific technology shocks. Our results are in line with [Edelstein & Kilian \(2009\)](#), who find that energy price shocks are a quantitatively important (but not dominant) determinant of changes in aggregate consumption.<sup>21</sup> Furthermore, the neutral technology shock explains only a considerable share of the variation in labor productivity — the endogenous variable in the equation that identifies the neutral technology shock. The monetary policy shock seems even less relevant. Even though the identifying assumptions of our VAR model are inherently not testable ([Kilian & Vega 2011](#)), two observations provide indirect support for our strategy. On the one hand, only 11% of the changes in the nominal interest rate at business cycle frequencies are due to the unexpected shock. This implies that the Fed’s monetary policy has followed a rule-based approach over our sample period (this conclusion is consistent with the results of [Sims & Zha 2006](#)). On the other hand, more than 50% of the cyclical variability in the commodity price index is explained by the commodity price shock itself (note that, e.g., [Cochrane 1994](#), finds even higher values). This result indicates that the contemporaneous exogeneity assumption is a reasonable identifying restriction.

The importance of investment-specific technology shocks is in line with the results of several recent SVAR studies by [Fisher \(2006\)](#), [Ravn & Simonelli \(2008\)](#), [Canova et al. \(2010\)](#), and [Altig et al. \(2011\)](#).<sup>22</sup> Altogether, the four identified shocks account for 49%-73% of business cycle volatility in the data. At first glance, however, it seems surprising

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<sup>20</sup>This result is consistent with the finding of [Hanson \(2004\)](#), who shows that the commodity price index by the [BLS \(2012\)](#) is a good predictor of the inflation rate.

<sup>21</sup>Note that the same authors ([Edelstein & Kilian 2007](#)) are unable to find a significant response in nonresidential fixed investment. Their data set, however, differs from ours in that we include consumer durables in investment.

<sup>22</sup>[Smets & Wouters \(2007\)](#) and [Mumtaz & Zanetti \(2012\)](#) draw the conclusion that neutral technology shocks are more important. Both sets of authors use a data set that includes consumer durables in consumption (and not in investment). [Schmitt-Grohé & Uribe \(2011\)](#) argue that a common stochastic trend in neutral and investment-specific technology is the main driving force for the business cycle.



that neutral technology shocks do not explain larger shares at business cycle frequencies. Therefore, we analyze also the explanatory power of neutral technology shocks across the whole spectrum (Figure 12). Indeed, we find that neutral technology shocks play a very important role in explaining macroeconomic fluctuations (particularly, output, labor productivity, and consumption), but at low frequencies.

### 3.2.3 Historical Decomposition of Shocks

As suggested by Edelstein & Kilian (2009), the following section studies the cumulative effects of the identified structural shocks on output and inflation. The historical decomposition (see Figure 13) is based on the code provided by Altig et al. (2011). For stationarity considerations, we only report the results of the bandpass filter specification here (Kilian & Lewis 2011). When all four identified structural shocks are considered, we observe that our SVAR model is able to replicate the cyclical behavior of output remarkably well. There are only two episodes in U.S. postwar history that exhibit a noticeable tracking error. The model explains neither the short recession in the late 1960s, nor the depth of the recession after the burst of the so-called dot-com bubble.

We also investigate the time series elicited by the four individual shocks. The graphs illustrate that their contribution varies considerably across different episodes in the U.S. postwar period. *Commodity price* shocks contribute most to the high degree of macroeconomic volatility — both in output and inflation — in the 1970s, particularly during and after the first OPEC oil crisis.<sup>23</sup> In addition, commodity price shocks are also an important determinant of the double-dip in the early 1980s, the economic boom in the early 1990s, and the short early 2000s recession. We also note that commodity prices explain only a moderate share of inflation volatility in the later 1970s/early 1980s, but a large share of the decline in inflation during the last two recessions in our sample. Figure (8) evaluates the impact of the Fed’s response to commodity price shocks. Therefore, we examine the cyclical movements of aggregate output in the absence of the contractionary monetary policy feedback rule.<sup>24</sup> Interestingly, we observe that the monetary policy feedback rule — in particular, the contractionary response during the first OPEC oil crisis and the subsequent monetary easing — has amplified the output fluctuations caused by unexpected changes in commodity prices.

Consistent with the estimated impulse response functions, the counterfactual time series seems to lag the estimated path by about one year. Moreover, we note that the Fed’s policy rule was not able to avoid the spike in the inflation rate around the year 1974, but had disinflationary effects in the subsequent period (a similar pattern can also be observed during the early 1900s recession). In other words, our SVAR indicates that

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<sup>23</sup>This result is not as trivial as it seems. Kilian & Lewis (2011), on the other hand, argue that the cumulative effects of oil price shocks were not large (even prior to 1987).

<sup>24</sup>See Section (3.1.1) for details and motivation of this counterfactual exercise.

a contractionary monetary policy feedback rule may help the Federal Reserve to achieve price stability at longer forecast horizons, yet at the cost of output destabilization.

*Monetary policy* shocks, on the other hand, have played a role in the late 1960s recession, the double-dip in the early 1980s as well as in the subsequent recovery. Furthermore, in line with our previous results, we are unable to find a systematic relationship between *neutral productivity* shocks and fluctuations in aggregate output at business cycle frequencies. Neutral technology shocks seem rather important at low frequencies. For example, neutral productivity shocks suggest a deep recession between 1976 and 1983, reflecting the productivity slowdown in that period, and two long-lasting economic booms — the first in the mid 1980s and the second in the late 1990s. Consistent with Greenwood et al. (1997), we find that *investment-specific* technology growth was particularly strong during the productivity slowdown of other factors in the 1970s. Furthermore, investment-specific technology shocks appear to be a principal driving force for the 1960-61 recession and the following economic expansion, the 1973-75 recession (both output and inflation), the early 1980s recession as well as of the subsequent recovery.

## 4 Robustness Analysis

The following section presents a number of robustness checks. We investigate the sensitivity to the identifying assumptions (predeterminedness, global demand), the usage of the Thomson Reuters (2013) Continuous Commodity Index with time-invariant weights (see also Footnote 5), the data treatment, the choice of the lag length, and the selected sample period. We demonstrate that our results are robust across alternative model versions.

### 4.1 Non-Predetermined Commodity Prices

Section (2.2) assumes that no single shock (but the commodity price shock itself) has an impact on the contemporaneous value of the commodity price index. Rotemberg & Woodford (1996) proposed this assumption to identify nominal oil price shocks. Given that the broad commodity price index may behave differently from the nominal oil price (Alquist et al. 2011), the current section examines the robustness of our identification strategy. Therefore, we relax the contemporaneous exogeneity assumption to allow for immediate responses in the commodity price index to unexpected changes in two main indicators of the U.S. economy (labor productivity growth and per-capita hours). This procedure is different to the one used by Blanchard & Galí (2010), who have explored the consequences of an alternative recursive ordering of the variables. Importantly, as our SVAR is overidentified, we are able to eliminate these two identifying assumptions without imposing a new one.

Consequently, the identified commodity price shock  $\epsilon_t^p$  is now obtained by estimating:

$$p_t = \alpha^p - \beta_{a,0}^p \Delta a_t - \beta_{h,0}^p h_t + \sum_{j=1}^M \beta_{x,j}^p x_{t-j} + \epsilon_t^p \quad (8)$$

Since  $\epsilon_t^p$  may be correlated with  $\Delta a_t$  (via equation 5) and  $h_t$  (via equation 7), we estimate equation (8) with 2SLS. The set of instruments includes a constant and the following vector:<sup>25</sup>

$$[\Delta q_{t-j}, \Delta a_{t-j}, z_{t-j}, r_{t-j}, p_{t-j}]_{j=1}^{M+1}.$$

We find that the estimated results are remarkably robust (see Figure 17). In particular, the CPI inflation rate behaves almost identical to Section (3.1.1). Also the median response of the Federal Funds rate matches the previous estimates closely – even though the confidence bands are somewhat wider. Moreover, Table (5) shows that the business cycle variance decomposition statistics remain virtually unchanged.

## 4.2 External Demand

The present identification procedure of the commodity price shock is unable to distinguish between supply- and demand-driven innovations. However, the assumption that commodity price shocks are contemporaneously exogenous to U.S. macroeconomic aggregates seems more defensible in the case of supply shocks (e.g., political strife in the Middle East, see Kilian 2008) than in the case of demand shocks.<sup>26</sup> Therefore, we extend our SVAR by adding a variable that captures variations in global demand for commodity goods. In particular, we choose to include the natural log of the ratio of real exports to real imports of goods and services (see Table 2). Based on this series, we identify an external demand shock using short-run restrictions. Following Abbritti & Weber (2010) or Kilian & Lewis (2011), we assume that the process for the real export/import ratio is independent of the current realizations of all other variables but the commodity price index.<sup>27</sup>

The impulse responses generated by the four remaining shocks, particularly by the commodity price shock (Figure 16a), remain virtually unchanged when we control for unexpected movements in external demand. In addition, the variance decomposition statistics at business cycle frequencies are remarkably robust (Table 5). We note only a mild reduction (4 percentage points) in the explanatory power of the commodity price

<sup>25</sup>Note that the parameters are overidentified, given that the number of instruments exceeds the number of parameters. Using an overidentifying restrictions test (Sargan 1964) we are unable to reject this specification at the 10% significance level.

<sup>26</sup>Kilian & Murphy (2010) emphasize that political strife in the Middle East may not only disrupt oil supply, but also boost speculative demand.

<sup>27</sup>Note that the idea to disentangle commodity demand and supply shocks is due to Kilian (2009).

shock with respect to the cyclical movements in the inflation rate.<sup>28</sup> Figure (16b) illustrates the effects of the identified external demand shock. This shock represents a temporary but persistent rise in the real exports/imports ratio. We observe that the external demand shock causes a hump-shaped increase in the commodity price index, representing commodity price changes due to heightened global demand. Except for investment and consumption, all other variables show barely significant responses, which may be attributed to the fact that the U.S. is a relatively closed economy. Besides, the external demand shock is unable to explain significant shares in the business cycle variance of any variable but the real export/import ratio.

### 4.3 Thomson Reuters Continuous Commodity Index

Figure (18) shows the impulse responses when the Thomson Reuters (2013) Continuous Commodity Index with time-invariant weights (see Footnote 5) is used for the estimation of our SVAR model. In contrast to Section (3.1.1), the alternative commodity price index is more persistent with peak effects occurring about one year after the initial increase. Nevertheless, the *relative* impulse responses remain almost unchanged. Only the increase in the inflation rate seems to be more long-lived. After the initial increase, we observe that the inflation rate remains elevated for more than one year before eventually falling back to normal. We also note that the confidence bands are tighter. Consequently, the commodity price shock now becomes significantly more important in terms of the cyclical variance decomposition statistics (see Table 5), particularly with respect to inflation. The investment-specific technology shock, on the other hand, loses some of its explanatory power.

### 4.4 Data Treatment

**Bandpass Filter vs. Level Specification** The (b) panels of Figures (3)-(5), (7) and Table (4), respectively, display the impulse responses and the business cycle variance decomposition when we estimate the level specification of our SVAR model. We observe that all major conclusions survive this type of test. Even the response of per-capita hours to neutral technology shocks remains virtually unchanged (see also Section 3.1.3). The only notable difference between these two specifications is that the cyclical variance decomposition statistics of the commodity price shock are higher in the level specification, but the cyclical variance decomposition statistics of the investment-specific technology shock are higher in the bandpass filter specification. Altogether, these results indicate that the low-frequency bias becomes less important when the information set is sufficiently

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<sup>28</sup>Alternatively, we have used the “rest of the world” GDP index (1972Q1-2006Q4) by Enders et al. (2011), the “global economic activity” index (1968Q1-2007Q4) by Kilian (2009) and the “world industrial production” index (1948Q1-2007Q4) by Baumeister & Peersman (2012). Our (subsample) tests indicate that all indices yield similar results.

large. Furthermore, the remarkable resemblance of the impulse responses suggests that concerns about leaving (Fernald 2007, Francis & Ramey 2009, Canova et al. 2010) or removing (Gospodinov et al. 2011) low-frequency movements in the data are quantitatively not very important as long as the size of the information set is sufficiently large.

**Treatment of the Hours Series** In addition, Table (5) provides the cyclical variance decomposition statistics of output, per-capita hours, and the inflation rate under different model specifications. The figures confirm that our findings are robust to different filtering methods (differences, dummies, including the corresponding Francis & Ramey (2009) hours time series in the level specification).

**Treatment of the Commodity Price Index** Figure (14) shows the impulse responses when the commodity price index is differenced prior to applying the one-sided bandpass filter. This implies that we now identify a *permanent* shock to the level of the commodity price index. We observe that the shapes of the impulse responses are almost identical to Section (10). The most interesting difference is that the response of labor productivity is no longer significant, indicating that the elasticities of output and per-capita hours are of the same magnitude. Also the business cycle variance decomposition statistics (see Table 5) are very similar.

## 4.5 Lag Length

The present section investigates whether the chosen lag length has any impact on our results. For this purpose, we reduce the number of lags to  $M = 4$ . Table (5) shows that, in this case, the investment-specific technology shock becomes less important, but remains the principal driving force for output and per-capita hours over the business cycle. This indicates that our SVAR may suffer from “truncation bias” (Erceg et al. 2005) when the VAR order is insufficiently short. Qualitatively, however, we are unable to note any significant differences in the results.

## 4.6 Subsample Stability

Furthermore, we examine the subsample stability of the bandpass filter specification. Figure (15) illustrates the impulse responses to the identified commodity price shock before and after the appointment of Paul Volcker as chairman of the Board of Governors in August 1979. For this exercise — due to the smaller number of observations — we reduce the VAR order to  $M = 3$ . Note that, when plotting these graphs, we normalize the standard deviation of the commodity price shock in both sub-periods to the one measured in the full sample. We observe that output and the (core) inflation rate respond less elastic in the late sub-sample, but remain statistically significant at the 10% level. This result

is consistent with previous estimates by [Blanchard & Galí \(2010\)](#),<sup>29</sup> who attribute the milder response in the late sub-sample to (a) the smaller share of oil in production, (b) the decline in real wage rigidity, and (c) improvements in monetary policy.<sup>30</sup>

In contrast to their study, our SVAR explicitly controls for the decreasing share of oil in production (by using a broad commodity price index with time-varying weights) and identifies neutral and investment-specific technology shocks. Our result of smaller second-round effects is consistent with their view of a decrease in real wage rigidity. Moreover, we find evidence in favor of increased credibility of monetary policy. In the pre-Volcker period, we notice that the Federal Funds rate stays above its steady state level for about five quarters. Following the initial rise, the Federal Reserve reduces the nominal interest rate and keeps it below its long-run mean for the next ten quarters. This pattern is known as stop-and-go monetary policy.<sup>31</sup> In line with the conventional wisdom, we find no evidence for stop-and-go monetary policy in the post-Volcker period. The contractionary response in the post-Volcker period seems to be driven by the statistically significant hump-shape in the core inflation rate. Furthermore, consistent with the muted impulse responses, [Table \(5\)](#) shows that the explanatory power of the commodity price shock is somewhat lower in the post-Volcker period.<sup>32</sup>

## 5 Summary and Conclusion

This paper evaluates the importance of commodity price shocks in the U.S. business cycle. Therefore, we extend the standard set of identified shocks to include unexpected changes in commodity prices. The resulting SVAR shows that commodity price shocks are a very important driving force for macroeconomic fluctuations, second only to investment-specific technology shocks. In particular, commodity price shocks explain a large share of cyclical movements in inflation.

The impulse response analysis shows that commodity price shocks generate significant U-shaped responses in output, consumption, and per-capita hours. Most notably, the inflation rate displays a significant spike, followed by a rapid return to the initial level. The unexpected surge in the inflation rate prompts the Fed to elevate the nominal interest rate. Results of a counterfactual exercise (in the style of [Bernanke et al. 1997](#)) indicate

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<sup>29</sup>Our sub-sample periods are not exactly identical to the ones chosen by [Blanchard & Galí \(2010\)](#). Given our nine-dimensional SVAR, the number of degrees of freedom is not sufficient in order to estimate the model accordingly. We also note that [Edelstein & Kilian \(2009\)](#) or [Herrera & Pesavento \(2009\)](#) find milder responses in the post-Volcker period.

<sup>30</sup>We also confirm their conclusion that the size of the shock in the post-Volcker period is larger than in the pre-Volcker period. This implies that the “Great Moderation” is *not due* to smaller commodity price shocks.

<sup>31</sup>The results of [Evans & Fisher \(2011\)](#) suggest that the stop-and-go pattern in the Federal Funds rate is triggered by oil price shocks, while the significant hike is due to changes in prices of other commodities.

<sup>32</sup>Also note that the investment-specific technology shock is somewhat less important when we exclude the late 1990s Internet boom from our sample.

that the systematic contractionary response helped the Federal Reserve to achieve price stability in the long run, yet at the cost of a significant economic downturn in output and per-capita hours.

Our SVAR model also addresses the hours response to neutral technology shocks. In particular, we find that the response of per-capita hours is positive and marginally significant. This result is surprising, given that we control for low-frequency movements in the data (Canova et al. 2010). Further investigations show that this result is very robust to the treatment of the data as long as the size of the information set is sufficiently large. This result, which is in line with the evidence found by Forni & Gambetti (2011), confirms our choice to estimate a large-scale SVAR.

The sub-sample properties of our model are consistent with Blanchard & Galí (2010). We find that the effects of a commodity price shock on output and the inflation rate are milder in the post-Volcker period, but remain statistically significant at the 10% level. Several further robustness checks confirm the findings of our model. In particular, we examine robustness to the choice of the lag length, the identifying assumptions, the specific commodity price index used, and the inclusion of an external demand shock (Abbritti & Weber 2010).

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# A Tables

## A.1 Sources and Definitions of Data

Series	Definition	Source	Mnemonic
POP	civilian non-institutional population 16+	FRED	CNP16OV
FFR	effective (net) Federal Funds rate	FRED	FEDFUNDS
CPI	consumer price index (all urban consumers)	FRED	CPIAUCSL
PPI	producer price index (crude materials)	FRED	PPICRM
GOV	real government consumption expenditures & gross investment	FRED	GCEC96
EXP	real exports of goods & services	FRED	EXPGSC1
IMP	real imports of goods & services	FRED	IMPGSC1
HOU	hours in the business sector	BLS	PRS84006033
OUT	real output per hour in the business sector	BLS	PRS84006093
EMP	employment in the business sector	BLS	PRS84006013
RPI	quality-adjusted relative price of investment	DiCecio (2009)	p_i
CON	real personal consumption expenditures (nondurables & services)	DiCecio (2009)	cmdq + csq
INV	real quality adjusted gross private fixed investment + PCE durables, divided by 100	DiCecio (2009)	r_inv
CCI	continuous commodity index	Datastream	NYFECRB

**Table 1:** This table displays the definitions of the raw series used. The BLS (2012, p. 5) defines crude materials for further processing as “[...] unprocessed commodities not sold directly to consumers. Crude foodstuffs and feedstuffs include items such as grains and livestock. The crude energy goods category consists of crude petroleum, natural gas to pipelines, and coal. Examples of crude nonfood materials other than energy include raw cotton, construction sand and gravel, and iron and steel scrap”. Current and historical weights can be downloaded at: [ftp : //ftp.bls.gov/pub/special.requests/ppi/](ftp://ftp.bls.gov/pub/special.requests/ppi/); e.g. [sopnew08.txt](#) summarizes the weights in December 2007. We also thank Riccardo DiCecio for kindly sharing his data. The quality-adjustment follows Gordon (1990), Cummins & Violante (2002), and Fisher (2006). Consumer durables are included in investment, but the change in inventories is not. We aggregate all monthly series to quarterly data.

## A.2 Definition of Variables in the SVAR

Variable	Symbol	Definition
growth in labor p'tivity	$\Delta a_t$	first difference of log (OUT)
growth in RPI	$\Delta q_t$	first difference of log (RPI)
per-capita hours	$h_t$	log of (HOU/POP)
inflation rate	$\pi_t$	first difference of log (CPI)
nominal interest rate	$r_t$	FFR
employment rate	$n_t$	log of (EMP/POP)
commodity price index	$p_t$	log of (PPI)
consumption share	$c_t$	log of (CON/(CON+INV+GOV+EXP-IMP))
investment share	$i_t$	log of (INV/(CON+INV+GOV+EXP-IMP))
export/import ratio	$d_t$	log of (EXP/IMP)
continuous commodity index	$cci_t$	log of (CCI)

**Table 2:** This table displays the variables that enter the SVAR. The trivariate model (Canova et al. 2010) uses only the first three variables. The last two variables are only used for robustness checks.

### A.3 Cross Correlations with Technology Shocks

		0	1	2	3	4	5
$c_t$	lag	0.091	0.080	0.090	0.082	0.094	0.059
	lead	0.091	0.108	0.108	0.069	0.028	0.033
$n_t$	lag	-0.092	-0.033	-0.018	-0.019	-0.012	-0.001
	lead	-0.092	-0.105	-0.119	-0.085	-0.064	-0.077
$r_t$	lag	-0.104	-0.052	-0.016	0.021	0.020	0.068
	lead	-0.104	-0.152*	-0.190**	-0.210**	-0.197**	-0.145*
$\pi_t$	lag	-0.280**	-0.040	-0.047	-0.122	-0.074	-0.010
	lead	-0.280**	-0.171*	-0.136	-0.136	-0.107	-0.112
$i_t$	lag	0.001	-0.026	-0.012	-0.033	-0.045	-0.033
	lead	0.001	0.012	0.038	0.076	0.100	0.094
$p_t$	lag	-0.188**	-0.065	-0.050	-0.062	-0.025	0.003
	lead	-0.188**	-0.154*	-0.134	-0.129	-0.103	-0.100

(a) neutral technology

		0	1	2	3	4	5
$c_t$	lag	-0.038	-0.143*	-0.118	-0.148*	-0.083	-0.099
	lead	-0.038	-0.017	0.018	0.033	0.044	0.064
$n_t$	lag	-0.019	0.011	0.022	0.007	0.004	0.003
	lead	-0.019	-0.036	-0.081	-0.111	-0.115	-0.091
$r_t$	lag	0.053	0.096	0.095	0.021	0.010	-0.018
	lead	0.053	0.003	-0.056	-0.075	-0.048	-0.074
$\pi_t$	lag	-0.092	0.188**	0.175*	0.016	0.156*	0.123
	lead	-0.092	-0.012	-0.015	-0.137	-0.100	0.014
$i_t$	lag	-0.081	0.024	0.031	0.028	0.000	0.016
	lead	-0.081	-0.104	-0.171*	-0.177*	-0.133	-0.103
$p_t$	lag	-0.007	0.066	0.103	0.086	0.081	0.064
	lead	-0.007	0.000	-0.021	-0.051	-0.002	0.038

(b) investment-specific technology

**Table 3:** The table displays cross correlation coefficients between the the neutral and the investment specific technology shock series estimated using a trivariate SVAR following [Canova et al. \(2010\)](#), respectively, and the remaining six variables in our SVAR at leads and lags ( $\pm 5$  quarters). Stars (\*, \*\*) indicate significance at the 5% and 1% level, respectively.

## A.4 Variance Decomposition at Business Cycle Frequencies

	commodity prices	monetary policy	neutral tech	investment specific	all four shocks
$y_t$	12 (8)	8 (5)	11 (7)	21 (12)	52 (14)
$y_t/h_t$	8 (6)	6 (4)	17 (9)	13 (9)	44 (14)
$q_t$	9 (6)	3 (3)	11 (7)	29 (15)	51 (16)
$h_t$	11 (8)	6 (5)	6 (6)	29 (14)	52 (14)
$n_t$	11 (8)	7 (5)	6 (6)	27 (14)	51 (14)
$h_t/n_t$	13 (8)	5 (4)	9 (6)	29 (13)	55 (14)
$\pi_t$	17 (9)	6 (4)	12 (8)	14 (10)	49 (14)
$r_t$	14 (8)	11 (6)	7 (6)	26 (14)	58 (14)
$i_t$	14 (8)	8 (5)	9 (6)	20 (12)	51 (13)
$c_t$	14 (8)	8 (5)	13 (8)	20 (11)	55 (14)
$p_t$	51 (15)	4 (3)	7 (6)	11 (8)	73 (14)

(a) Benchmark Specification

	commodity prices	monetary policy	neutral tech	investment specific	all four shocks
$y_t$	17 (8)	7 (4)	14 (9)	17 (10)	55 (11)
$y_t/h_t$	11 (7)	5 (3)	16 (10)	19 (10)	51 (12)
$q_t$	13 (7)	3 (2)	10 (6)	18 (14)	43 (14)
$h_t$	14 (8)	6 (4)	9 (8)	20 (12)	50 (13)
$n_t$	14 (8)	6 (4)	9 (8)	23 (13)	52 (13)
$h_t/n_t$	17 (8)	5 (4)	9 (7)	14 (9)	45 (11)
$\pi_t$	25 (10)	5 (4)	9 (7)	17 (11)	57 (11)
$r_t$	17 (9)	13 (5)	8 (7)	20 (12)	58 (13)
$i_t$	17 (8)	8 (5)	11 (8)	16 (10)	53 (11)
$c_t$	20 (9)	8 (5)	14 (8)	15 (9)	57 (11)
$p_t$	60 (11)	6 (4)	5 (4)	7 (4)	77 (10)

(b) Level Specification

**Table 4:** The table displays the decomposition of variance at business cycle frequencies based on estimated spectral densities (following Altig et al. 2011). Numbers are means of point estimates across bootstrap simulations, numbers in parentheses are the corresponding standard deviations.

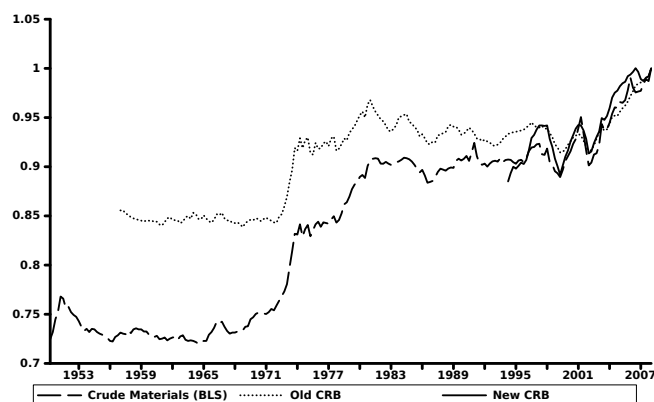
## A.5 Robustness

	commodity prices	monetary policy	neutral tech	investment specific	all four shocks
Benchmark Specification					
$y_t$	12 (8)	8 (5)	11 (7)	21 (12)	52 (14)
$h_t$	11 (8)	6 (5)	6 (6)	29 (14)	52 (14)
$\pi_t$	17 (9)	6 (4)	12 (8)	14 (10)	49 (14)
Level Specification					
$y_t$	17 (8)	7 (4)	14 (9)	17 (10)	55 (11)
$h_t$	14 (8)	6 (4)	9 (8)	20 (12)	50 (13)
$\pi_t$	25 (10)	5 (4)	9 (7)	17 (11)	57 (11)
Difference Specification					
$y_t$	15 (7)	9 (5)	15 (9)	15 (10)	53 (12)
$h_t$	12 (7)	7 (5)	11 (9)	21 (12)	51 (13)
$\pi_t$	24 (9)	6 (4)	9 (6)	15 (10)	54 (12)
Dummy Specification					
$y_t$	16 (8)	8 (5)	13 (8)	17 (10)	54 (11)
$h_t$	15 (8)	6 (4)	7 (6)	21 (13)	50 (13)
$\pi_t$	27 (10)	6 (4)	8 (6)	12 (9)	53 (11)
Francis & Ramey (2009) Hours					
$y_t$	11 (6)	6 (4)	15 (9)	18 (11)	50 (12)
$h_t$	9 (6)	6 (4)	9 (8)	22 (13)	46 (13)
$\pi_t$	19 (9)	4 (3)	8 (7)	19 (13)	50 (12)
Commodity Prices in First Differences					
$y_t$	10 (6)	8 (5)	14 (8)	18 (11)	50 (14)
$h_t$	9 (6)	6 (5)	6 (5)	28 (13)	50 (14)
$\pi_t$	17 (9)	7 (5)	8 (6)	12 (9)	44 (15)
Benchmark Specification with 4 Lags					
$y_t$	11 (7)	6 (4)	13 (8)	15 (9)	45 (13)
$h_t$	10 (7)	5 (4)	8 (6)	21 (12)	44 (13)
$\pi_t$	20 (10)	7 (4)	12 (7)	7 (5)	45 (13)
Pre-Volcker Period (1959Q1-1979Q2)					
$y_t$	16 (10)	12 (7)	13 (10)	15 (11)	57 (14)
$h_t$	17 (11)	13 (7)	11 (8)	14 (11)	55 (14)
$\pi_t$	17 (10)	12 (7)	10 (8)	15 (8)	55 (13)
Post-Volcker Period (1980Q1-2007Q4)					
$y_t$	9 (7)	8 (5)	8 (7)	23 (14)	48 (14)
$h_t$	9 (8)	7 (5)	8 (7)	23 (16)	47 (15)
$\pi_t$	13 (8)	6 (4)	7 (6)	18 (12)	44 (13)
External Demand Shock					
$y_t$	10 (7)	7 (5)	11 (7)	20 (12)	54† (14)
$h_t$	10 (7)	6 (5)	7 (6)	28 (15)	57† (14)
$\pi_t$	13 (8)	5 (4)	11 (9)	16 (10)	51† (14)
Non-Predetermined Commodity Prices					
$y_t$	12 (9)	7 (5)	12 (8)	18 (11)	49 (15)
$h_t$	12 (9)	6 (5)	7 (6)	27 (14)	52 (15)
$\pi_t$	17 (10)	5 (4)	11 (8)	15 (10)	48 (15)
Thomson Reuters Continuous Commodity Index					
$y_t$	22 (10)	5 (3)	9 (6)	14 (9)	50 (13)
$h_t$	20 (10)	4 (3)	7 (6)	14 (10)	45 (13)
$\pi_t$	31 (12)	3 (3)	8 (6)	11 (7)	53 (12)

**Table 5:** The table displays the decomposition of variance at business cycle frequencies based on estimated spectral densities (following Altig et al. 2011). Numbers are means of point estimates across bootstrap simulations, numbers in parentheses are the corresponding standard deviations. The “external demand shock” specification includes five shocks in total (denoted by a dag symbol †).

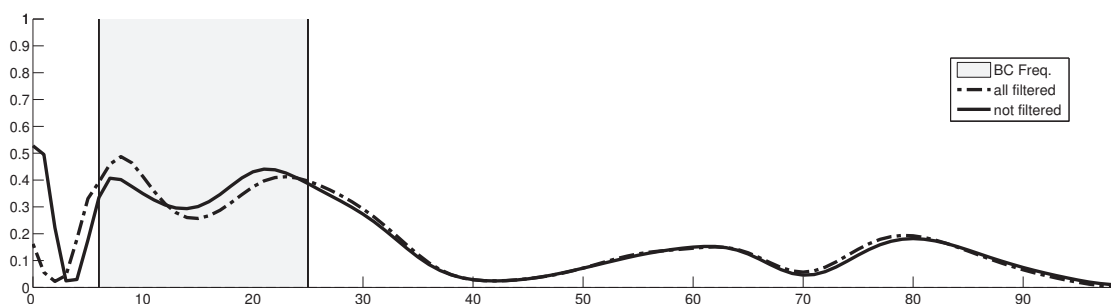
## B Figures

### B.1 Commodity Price Indices

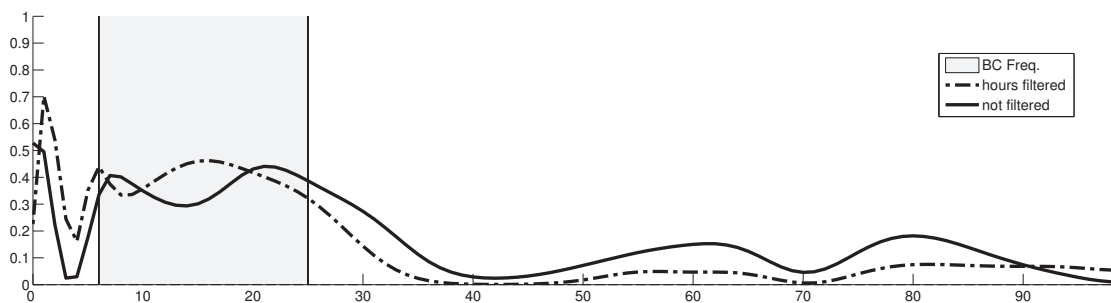


**Figure 1:** The figure illustrates the evolution of the “PPI: crude materials for further processing” index by the BLS (2012, p. 8, dashed line), the “old CRB” by Thomson Reuters (2013, calculated backwards until 1956Q4, dotted line), and the “new” Thomson Reuters/Jefferies (2013, solid line) CRB index, introduced in 2005, but calculated backwards until 1994Q1. All three indices are logged and normalized to unity in 2007Q4.

### B.2 Coherence Analysis



(a) all series are filtered

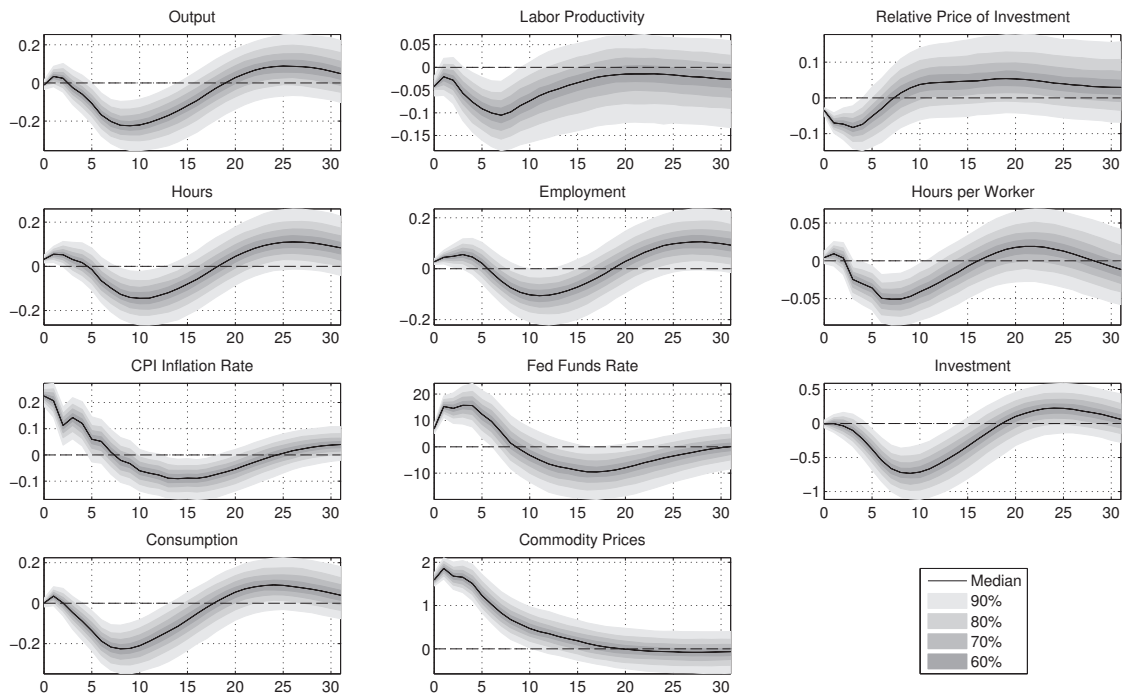


(b) only per-capita hours are filtered

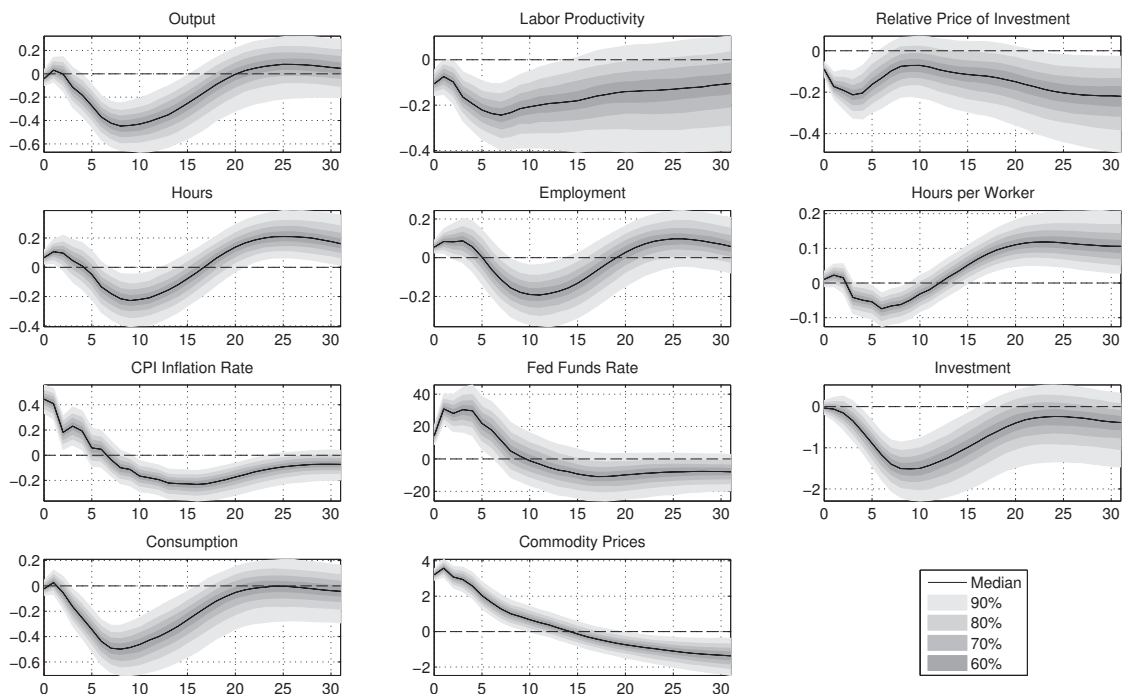
**Figure 2:** The figure illustrates the coherence between labor productivity growth and per-capita hours, estimated with five lags.



## B.3 Impulse Response Functions

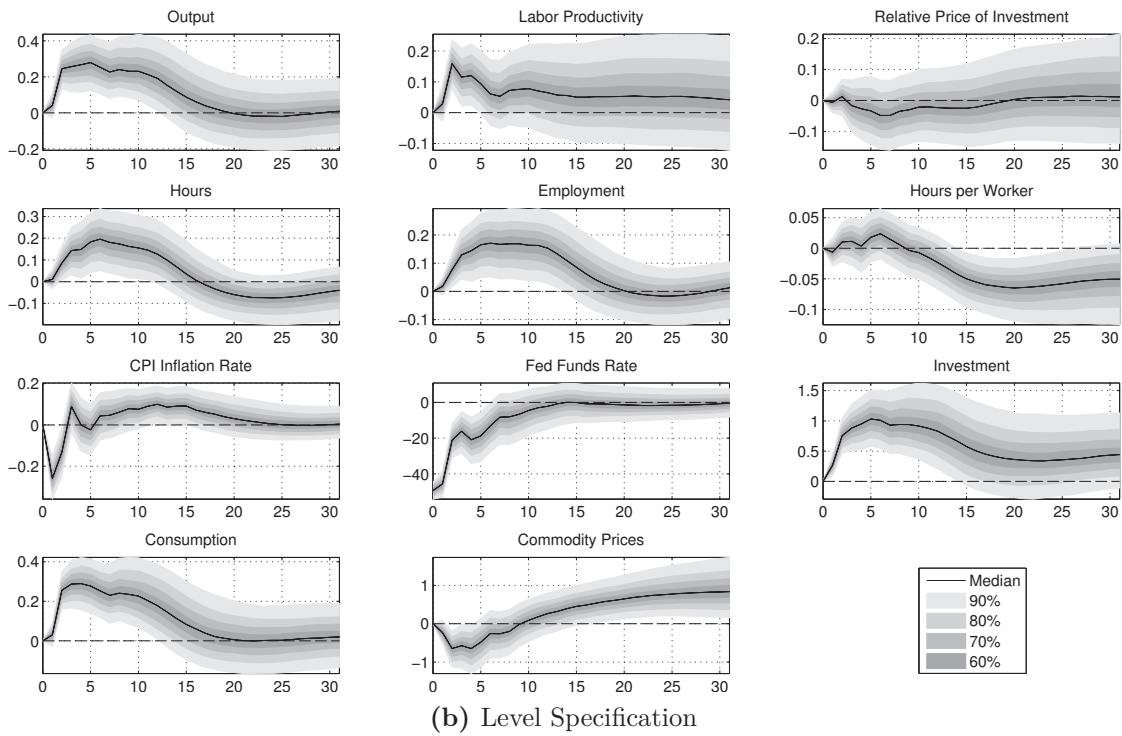
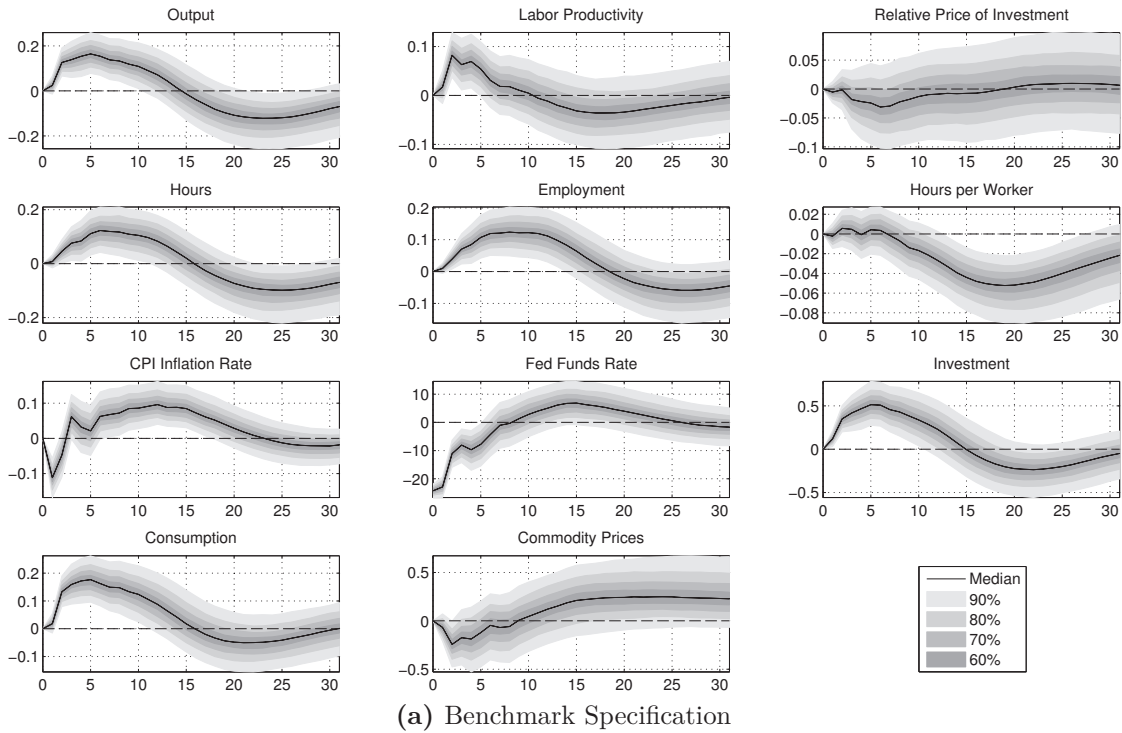


(a) Benchmark Specification

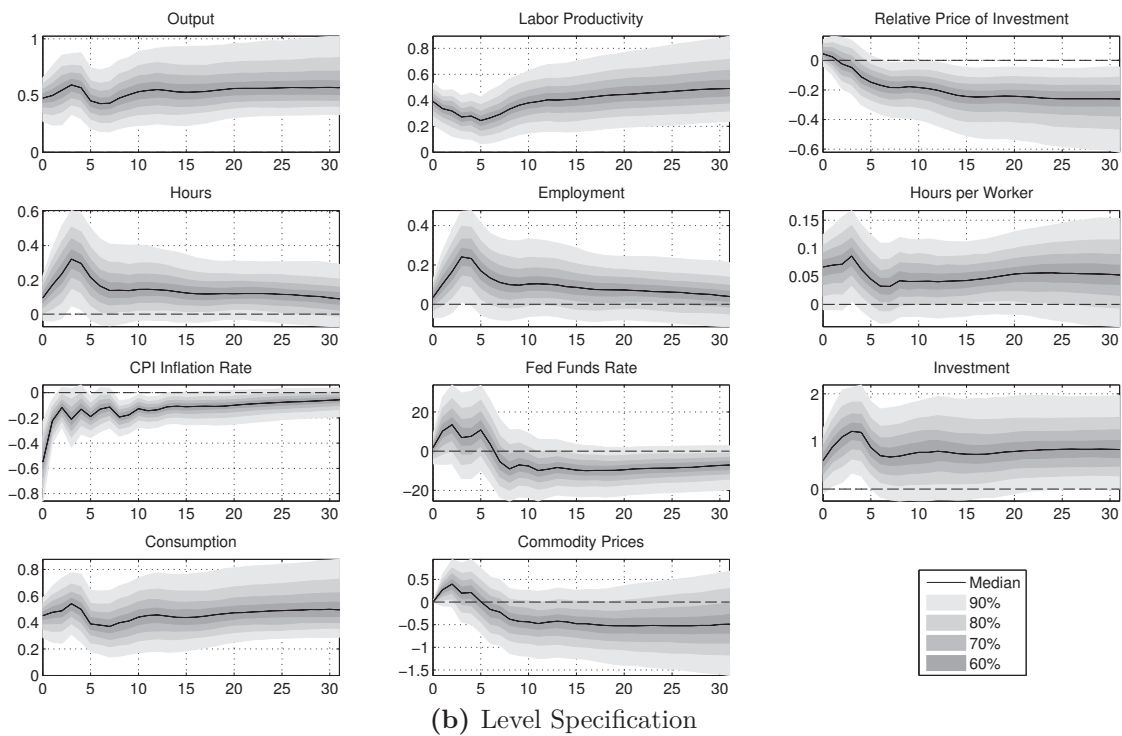
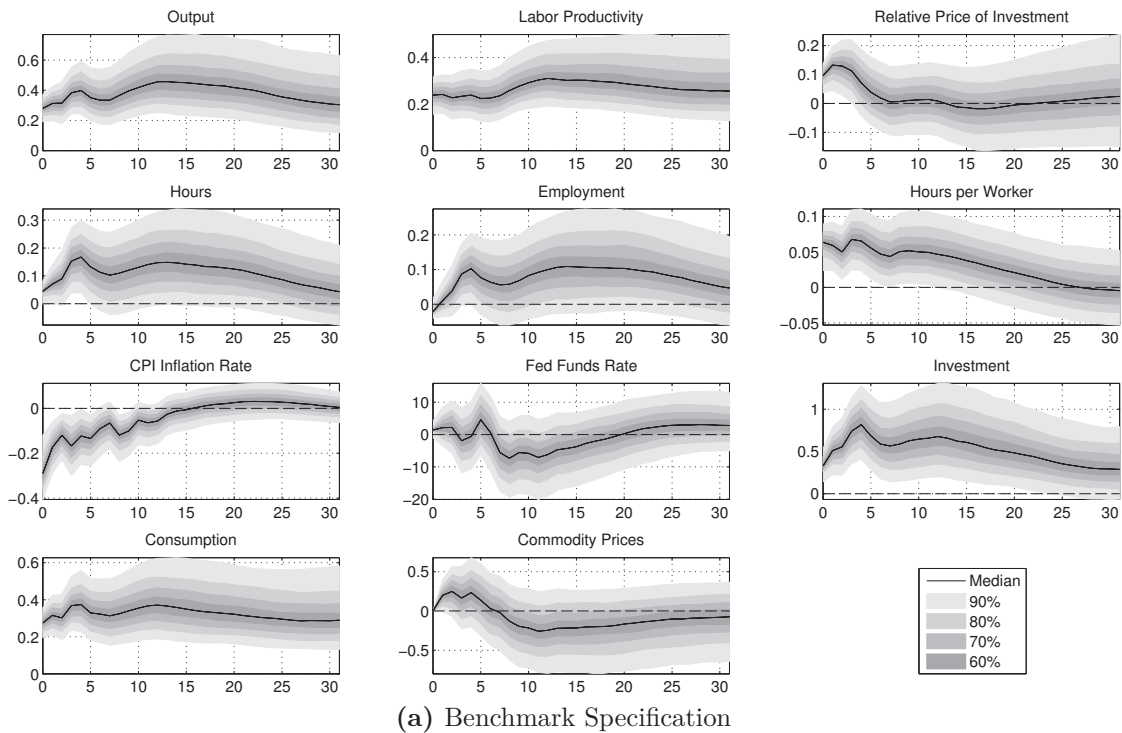


(b) Level Specification

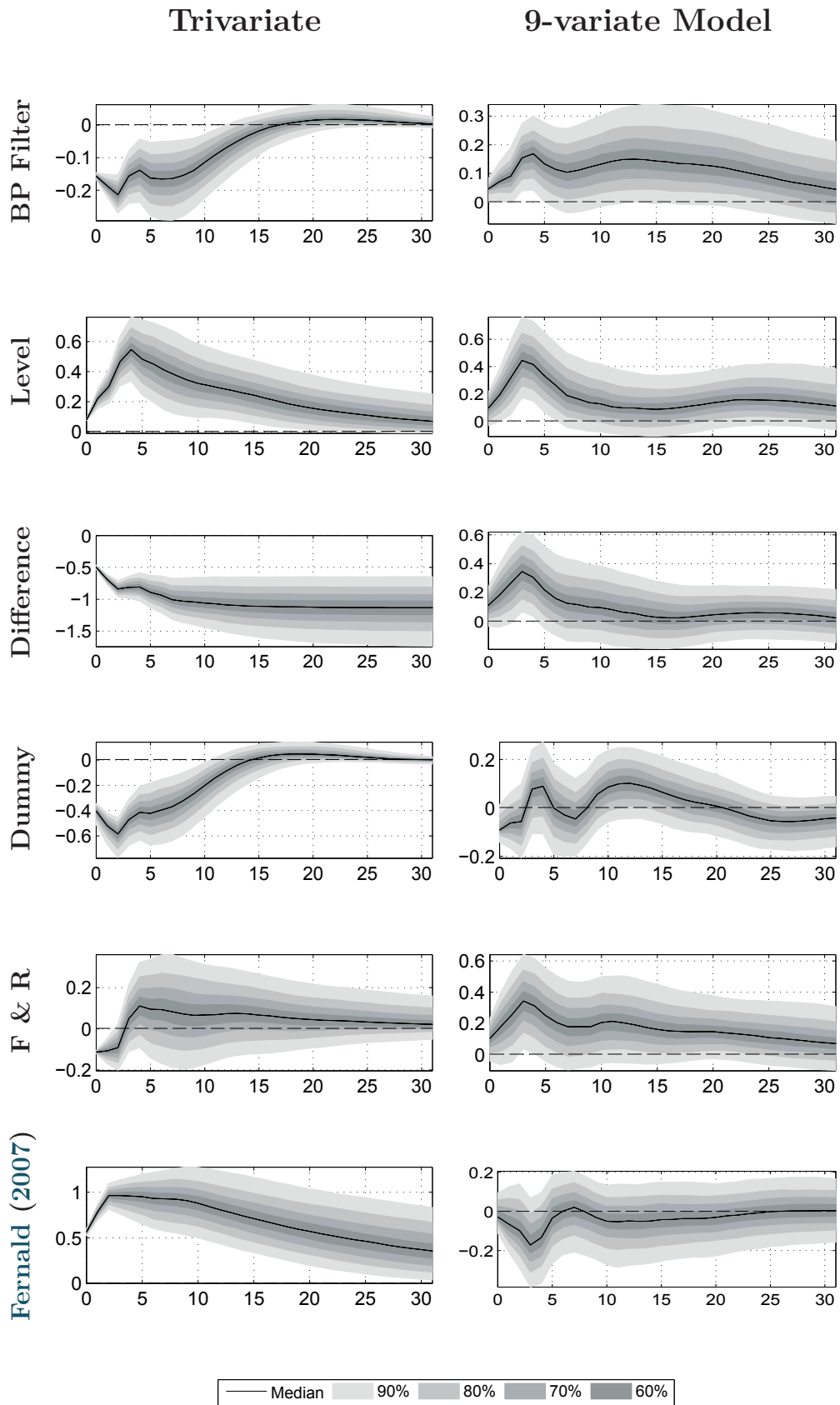
**Figure 3:** The figure illustrates the impulse responses to a commodity price shock.



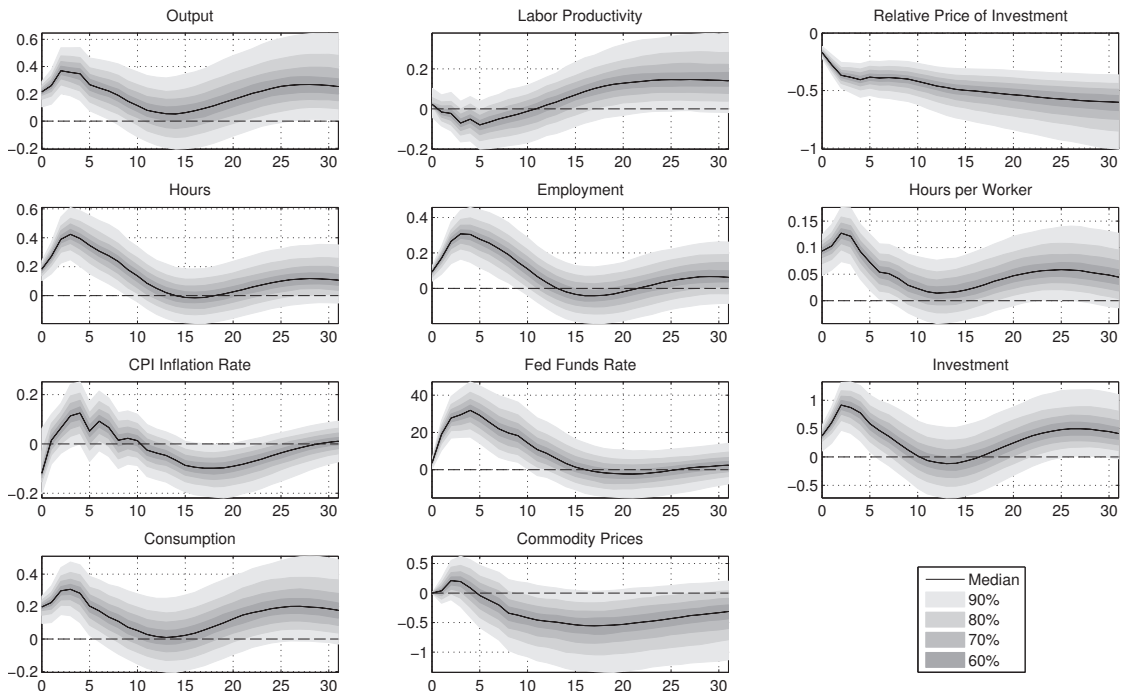
**Figure 4:** The figure illustrates the impulse responses to a monetary policy shock.



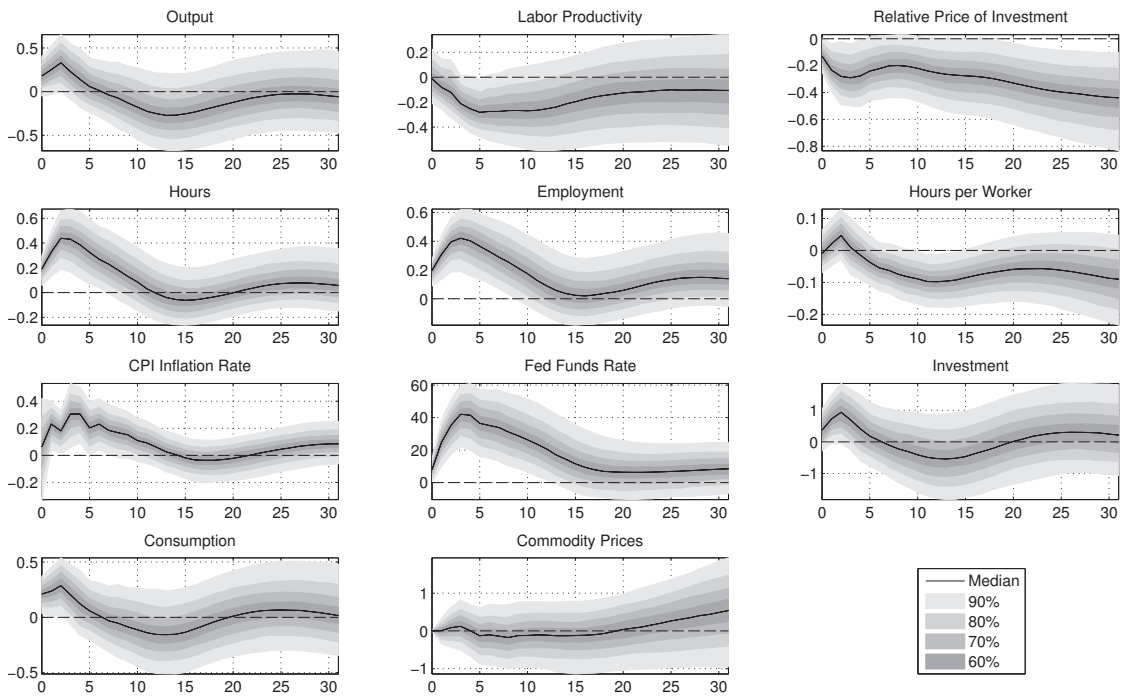
**Figure 5:** The figure illustrates the impulse responses to a neutral technology shock.



**Figure 6:** The figure illustrates the low-frequency bias by the means of the per-capita hours response to a neutral technology shock.



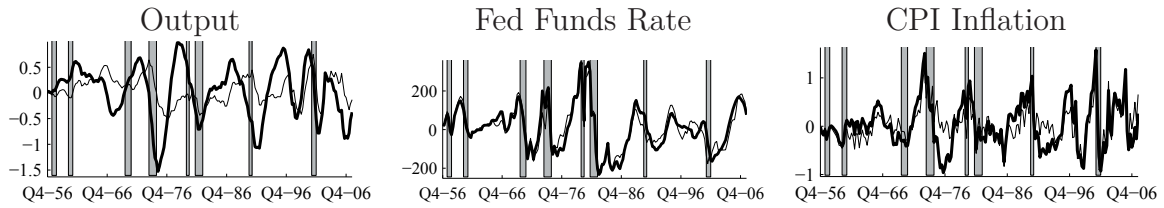
(a) Benchmark Specification



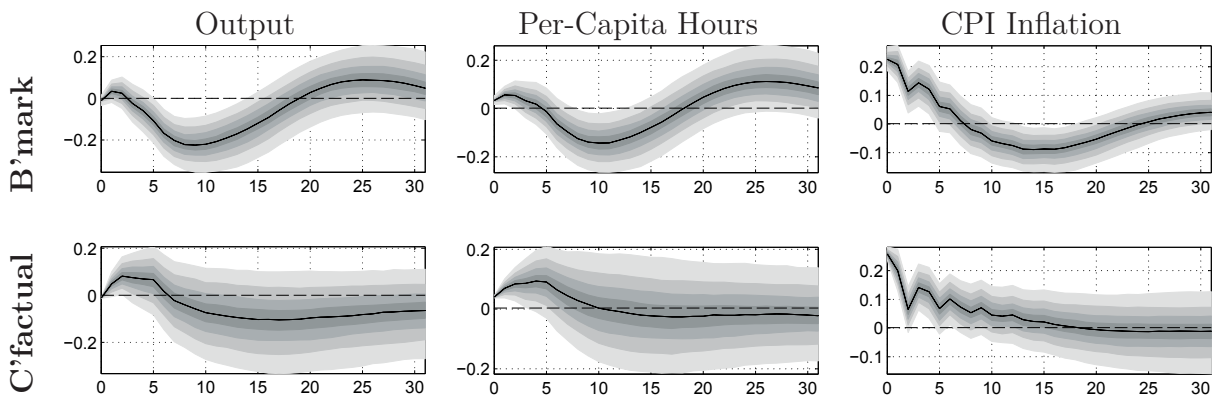
(b) Level Specification

**Figure 7:** The figure illustrates the impulse responses to an investment-specific technology shock.

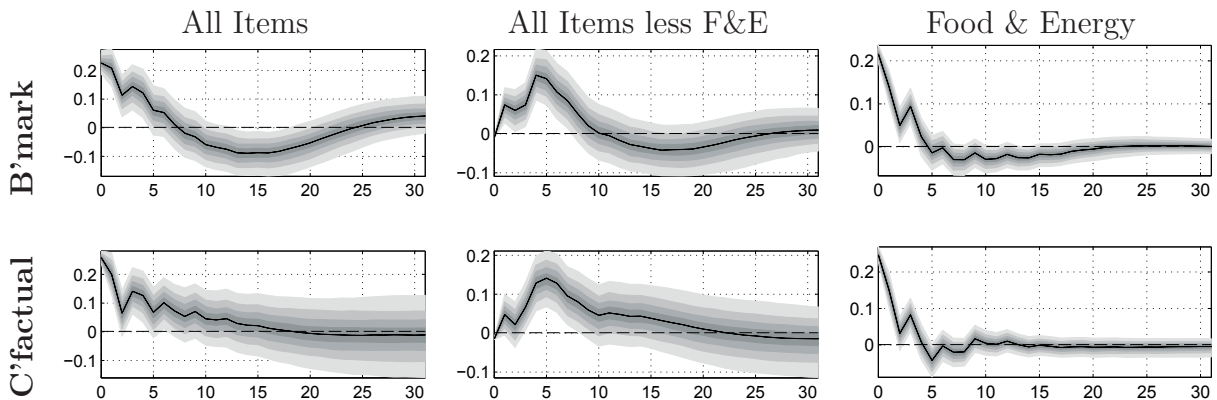
## B.4 Counterfactual Exercise



**Figure 8:** The graphs illustrate the time path of output, the Federal Funds rate, and inflation predicted by commodity price shock (bold line) with the same series predicted by the commodity price shock in the absence of a monetary policy feedback rule (thin line).

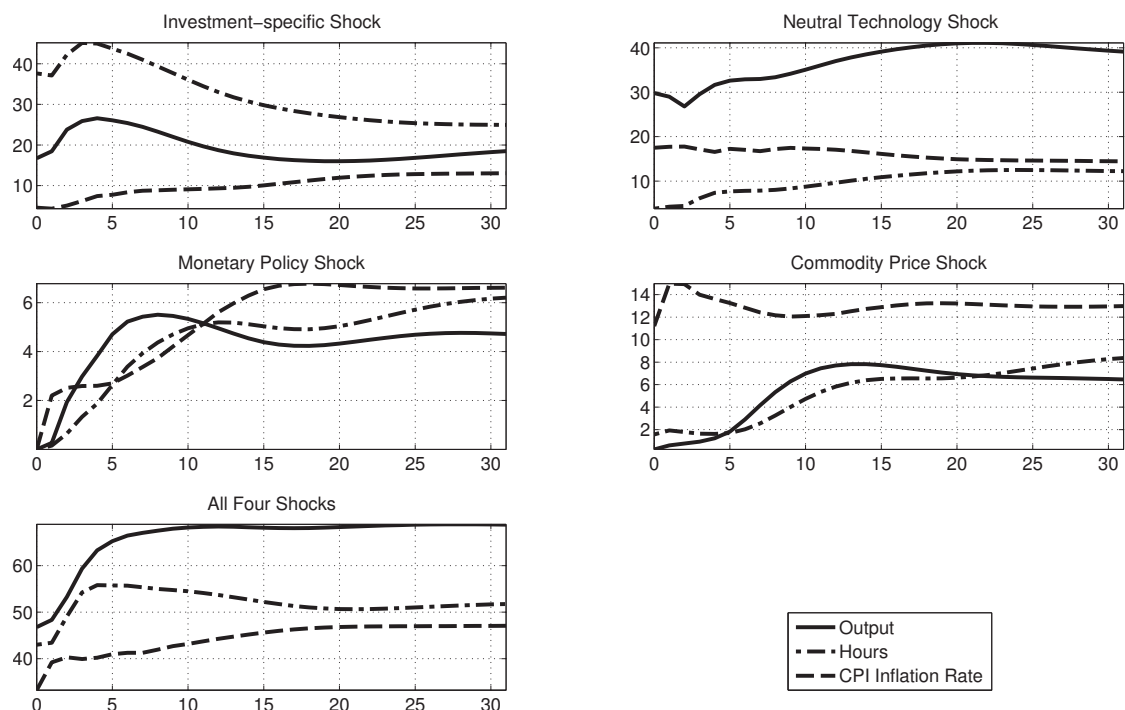


**Figure 9:** The top panel illustrates the responses of output, per-capita hours, and inflation to the estimated commodity price shock. The bottom panel illustrates the same responses when the Federal Funds rate — counterfactually — is assumed to be constant.



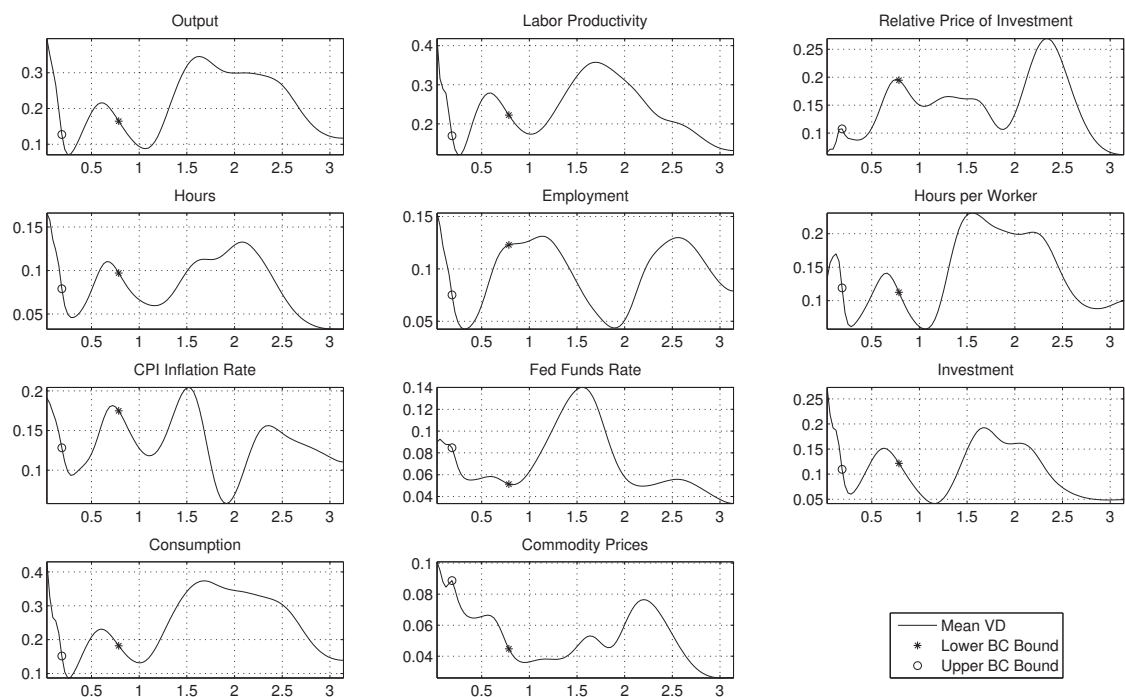
**Figure 10:** The figure illustrates the impulse responses of three CPI inflation measures to the identified commodity price shock; i.e., the “headline” inflation rate (all items), the “core” inflation rate (all items less food and energy), and the “food and energy” inflation rate. Due to limited data availability, the latter two responses are estimated using a slightly reduced sample period (1958Q2-2007Q4). The CPI “food and energy” is a weighted average of its components, using time varying weights (based on own calculations). All data are taken from FRED.

## B.5 Forecast Error Variance Decomposition



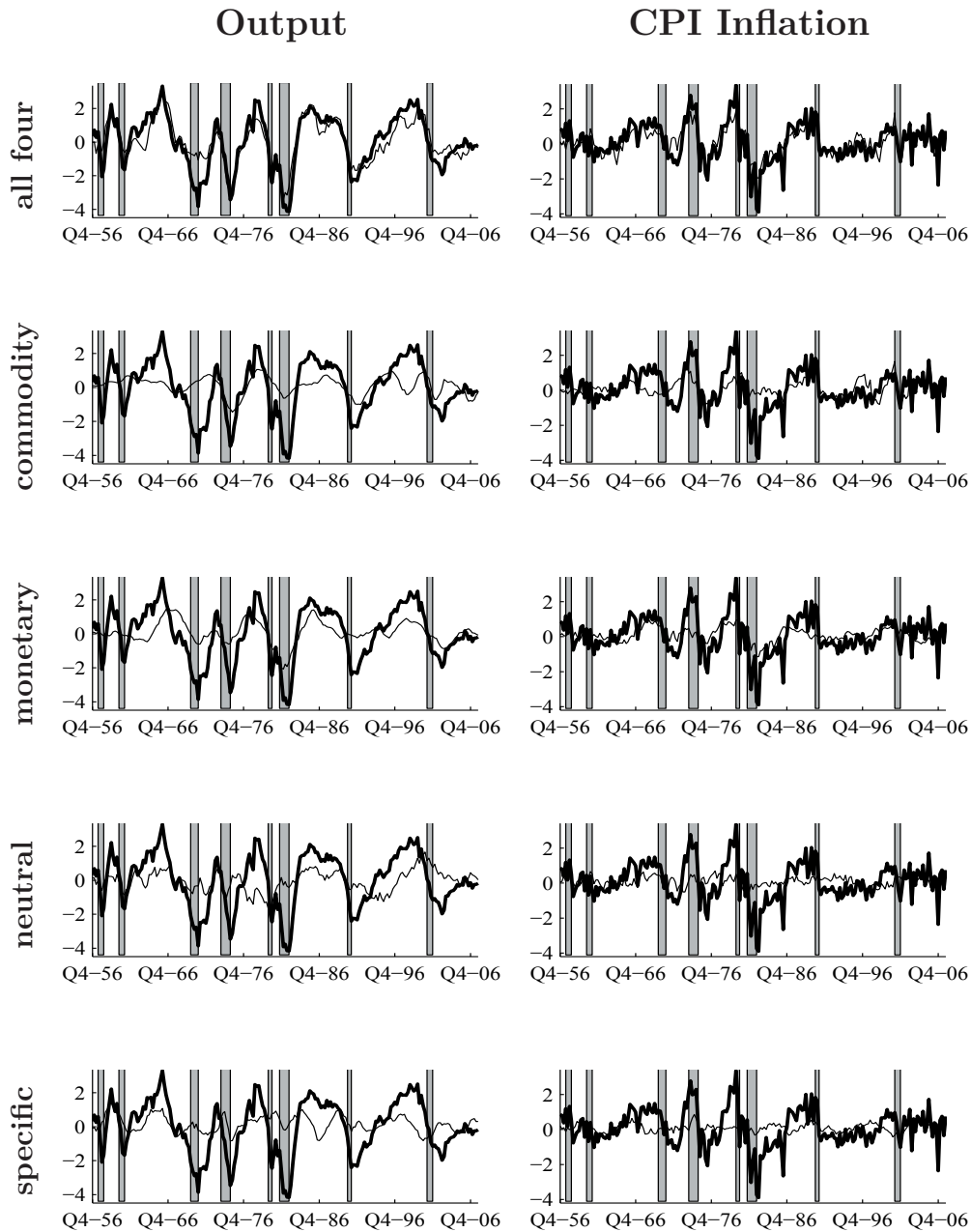
**Figure 11:** The figure illustrates the forecast error variance decomposition in our benchmark specification.

## B.6 Variance Decomposition at the Frequency Domain



**Figure 12:** The figure illustrates the explanatory power of neutral technology shocks across the whole spectrum in our benchmark specification.

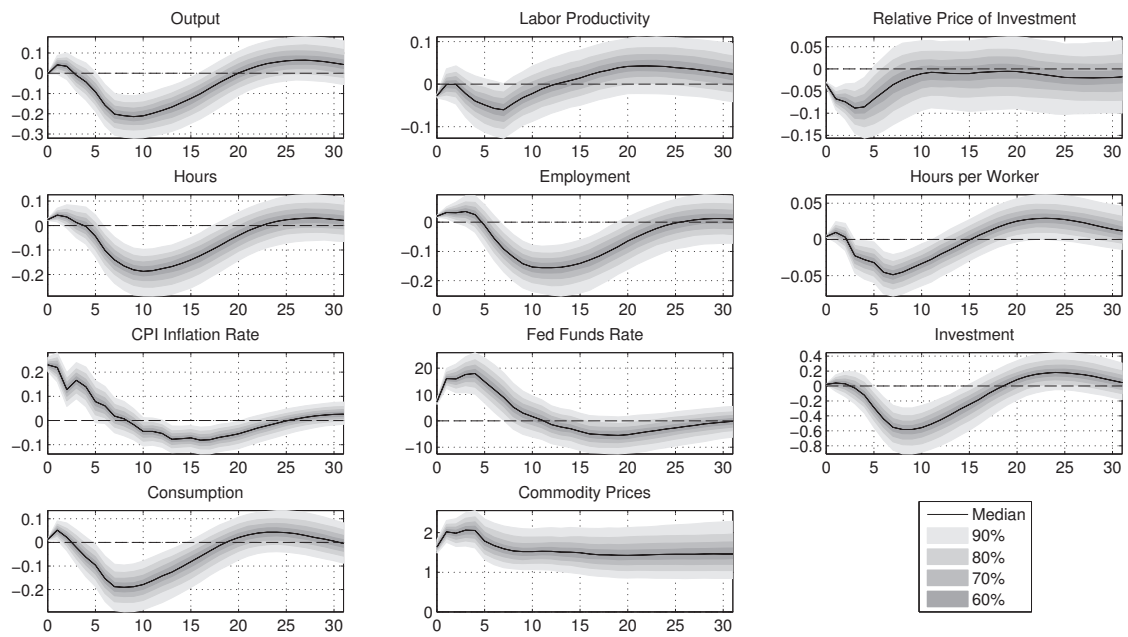
## B.7 Historical Decomposition of Shocks



**Figure 13:** The figure illustrates the historical decomposition of the four identified structural shocks for output and inflation. The bold line represents the bandpass filtered data, the thin line represents the time series predicted by the respective shock(s). In addition, the counterfactual exercise contrast the output/inflation series predicted by the commodity price shock (bold line) with the output/inflation series predicted by the commodity price shock in the absence of the monetary policy feedback rule (thin line).

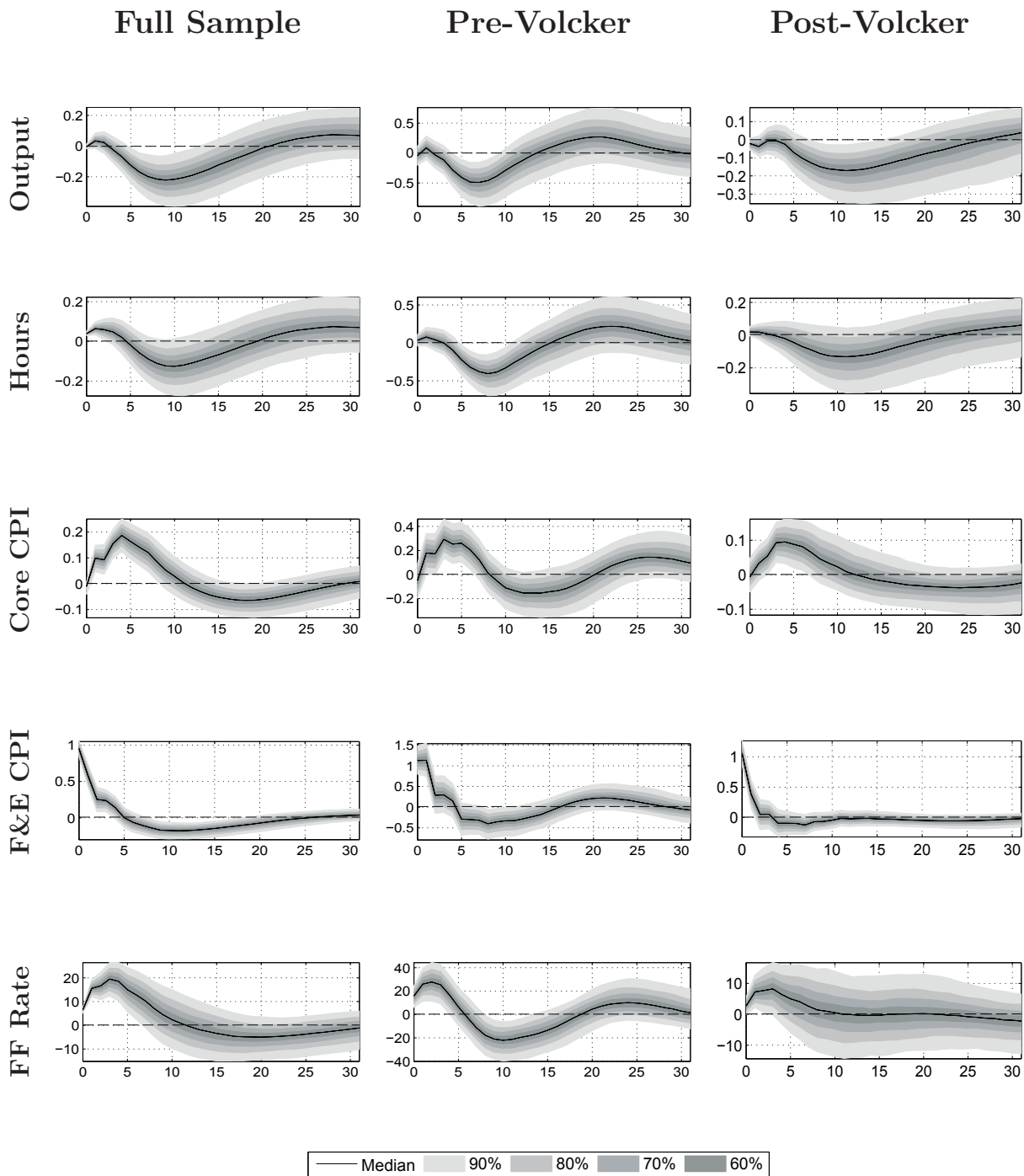


## B.8 Commodity Price Index in First Differences



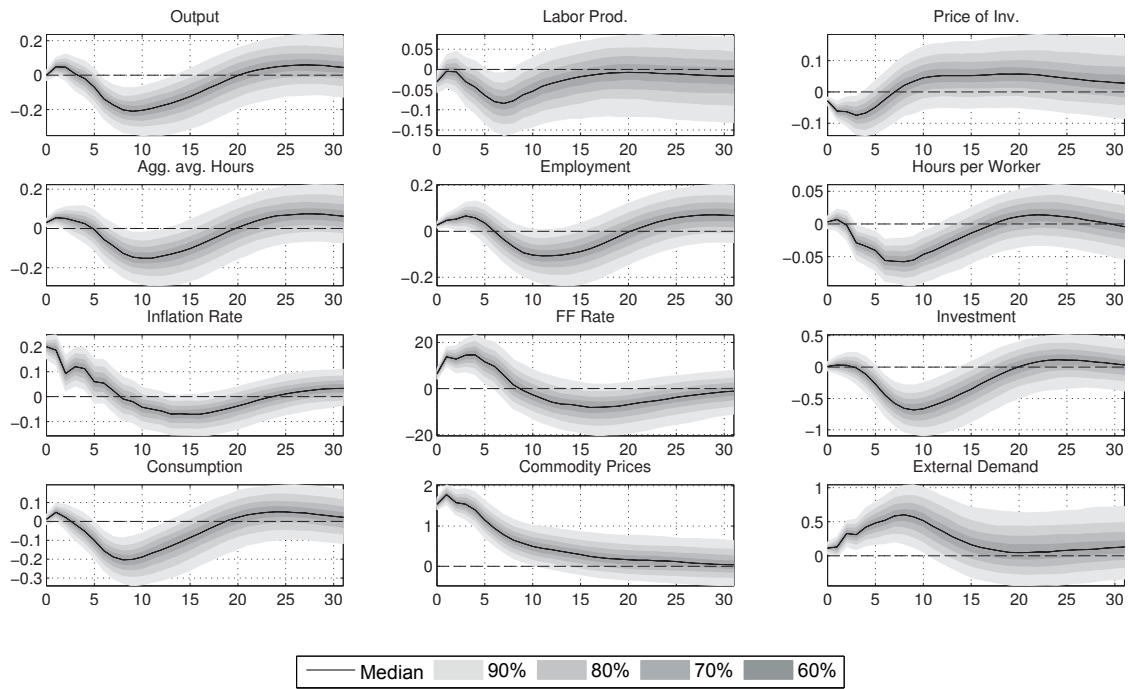
**Figure 14:** The figure illustrates the impulse responses to a commodity price shock when we first difference the natural logarithm of the commodity price index and then apply the one-sided bandpass filter.

## B.9 Subsample Stability

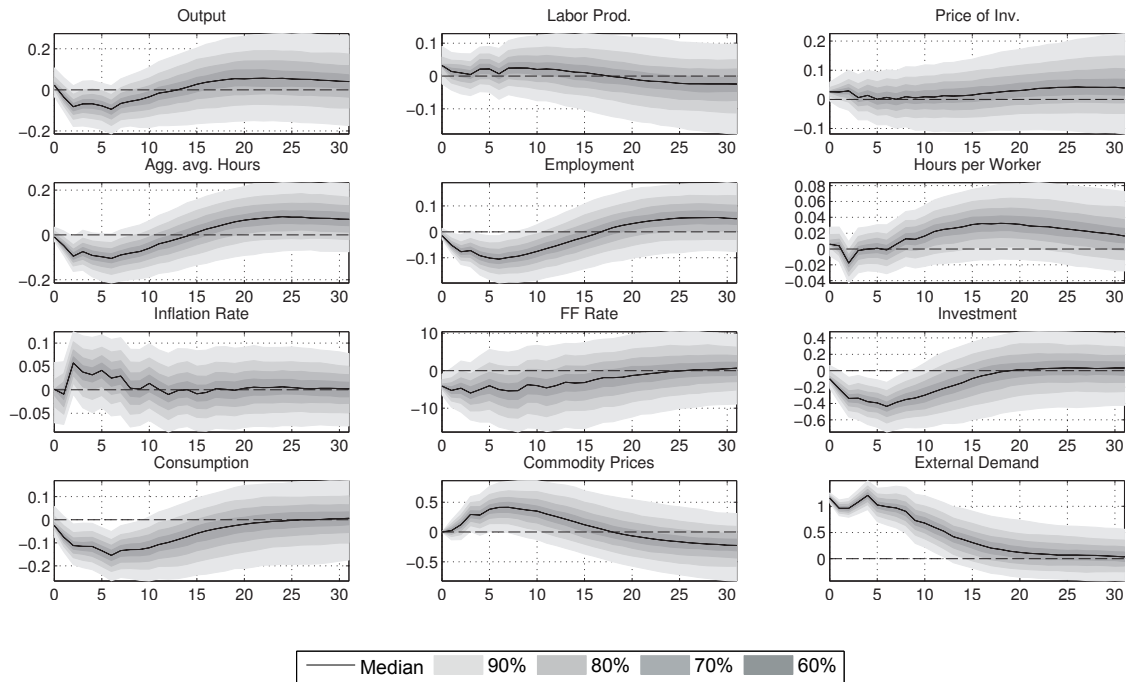


**Figure 15:** The figure contrasts the impulse responses to a commodity price shock in the “full” sample (1959Q1-2007Q4), the pre-Volcker period (1959Q1-1979Q2), and in the post-Volcker period (1980Q1-2007Q4). All three models are estimated with three lags ( $M = 3$ ). In order to facilitate comparability across sub-samples, the standard deviation of the commodity price shock is normalized to the standard deviation over the full sample.

## B.10 External Demand



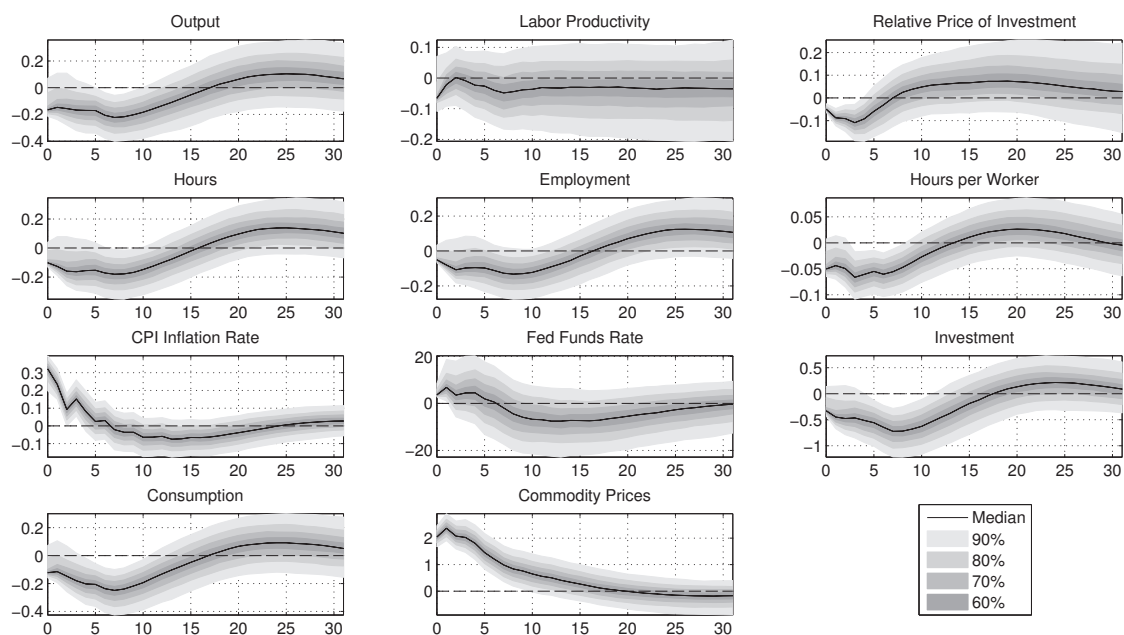
(a) Commodity Price Shock



(b) External Demand Shock

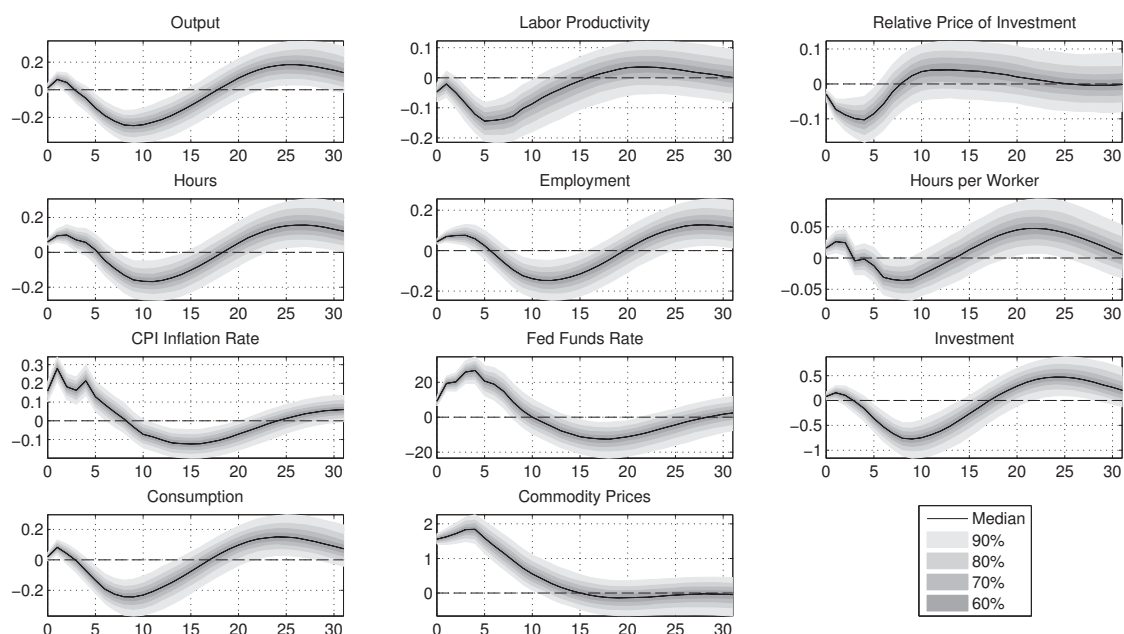
**Figure 16:** The figure illustrates the impulse responses when we include external demand shocks.

## B.11 Commodity Prices: Relaxed Exogeneity Assumption



**Figure 17:** The figure illustrates the impulse responses to a commodity price shock when commodity prices may depend on the current values of labor productivity and per-capita hours.

## B.12 Thomson Reuters Continuous Commodity Index



**Figure 18:** The figure illustrates the impulse responses to a commodity price shock when commodity prices are represented by the Thomson Reuters Continuous Commodity Index.

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