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The Effects of Uncertainty about Oil Prices in G-7

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Abstract

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Keywords: Oil, Volatility, Vector autoregression, Multivariate GARCH-in-Mean VAR.
JEL Classification: E32, C32.

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Abstract

The failure of decreases in oil prices to produce expansions that mirror the contractions associated with higher oil prices has been a topic of considerable interest. We investigate for the G-7 one explanation for this feature - the role of uncertainty about oil prices. In particular, we examine the link between oil price uncertainty and industrial production utilizing a very general and flexible empirical methodology that is based on a structural VAR modified to accommodate multivariate GARCH in mean. Our primary result is that oil price uncertainty has had a negative and significant effect on industrial production in four of the G-7 countries - Canada, France, UK and US. Impulse-response analysis suggests that, in the short-run, both positive and negative oil shocks may be contractionary. Our result helps explain why the sudden collapse in oil prices in the mid-1980's failed to produce rapid expansion in the G-7, and why the steady increases in oil prices from 2003-2007 did not induce recessions.

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1 Introduction

There has long been interest in the effects of uncertainty about energy prices on economic activity. Early theoretical foundations were established by Henry (1974) and Bernanke (1983), who show that uncertainty about energy prices will induce optimizing firms to postpone irreversible investment decisions as long as the expected value of additional information exceeds the expected short run return to current investment. For similar reasons, oil price uncertainty may induce consumers to postpone the purchase of durable goods. This theory explains why uncertainty about oil prices may induce auto manufacturers to postpone decisions to invest in either the development of SUV's or compact cars, and why consumers may postpone decisions to purchase such vehicles. Uncertainty about energy prices may also induce consumers to increase precautionary savings, depressing broad measures of current consumption.

Such shifts in expenditures can have large effects on aggregate output and employment if there are substantive frictions in the sectoral reallocation of labor and capital, as illustrated by Hamilton (1988). An interesting feature of these models, in contrast to real business cycle models such as Kim and Loungani (1992), is that the effects of oil price increases and oil price decreases are not symmetric. That is, the mechanisms described by Bernanke (1983) and Hamilton (1988) may cause both oil price increases and oil price decreases to be contractionary in the short-run.

In this paper, we investigate the effects of oil prices and uncertainty about oil prices in G-7 countries by utilizing a simultaneous equations model that accommodates both effects. The model is based on a structural VAR that is modified to accommodate multivariate GARCH-in-Mean errors, as detailed in Elder (1995, 2004) and Engle and Kroner (1995). We measure uncertainty about the impending oil price as the conditional standard deviation of the forecast error for the change in the price of oil. This empirical model permits changes in oil prices which are accompanied by a high degree of uncertainty to have different effects than changes in oil prices that are more easily forecastable. Consistent with Bernanke

(1983), if changes in oil prices are accompanied by an increase in uncertainty, then both increases and decreases in oil prices may be contractionary in the short-run.

This investigation is interesting and relevant for a number of reasons. First, our results provide direct evidence on role of oil price uncertainty and whether the response of output to oil shocks is asymmetric, which are issues of considerable recent interest (cf. Hamilton (2003) and Kilian (2008)). Second, applying this empirical model to the G-7 provides a test of robustness of Elder and Serletis (2008), who find that oil price uncertainty adversely affects consumption, investment and output in the US. Third, the cross section of G-7 countries offers a diverse pattern of oil consumption, oil exports and economic conditions. For example, oil expenditures as a share of GDP for the US were 4.8% in 2003 and as high as 8% in the early 1980's. Such expenditures for both the US and Canada are considerably larger than for the remaining G-7 countries. Our sample also includes two countries that were oil exporters over much of our sample: Canada (since the mid 1980's) and the UK (prior to about 2005). Finally, applying this empirical model to the G-7 also provides additional insight into whether the apparent asymmetry in the response of US output to oil prices is actually due to domestic factors, such as US tax legislation, as suggested by Kilian (2008).

Empirical evidence related to asymmetries in the response of output to oil shocks in the US was reported by Loungani (1986), Davis (1987), Mork (1989) and Davis and Haltiwanger (2001). Evidence has been reported more recently in an international context by Cuñado and Pérez de Gracia (2003), Huang, Hwang and Peng (2005) and Jimenez-Rodriguez and Sanchez (2005). Cunado and Perez de Gracia (2003) find that oil price shocks have significant effects on economic growth for a sample of European countries. Huang, Hwang and Peng (2005) find that oil price shocks have asymmetric effects on economic growth in the US, Canada and Japan. Jimenez-Rodriguez and Sanchez (2005) find that oil price shocks have negative effects on output growth for the UK and all oil importing countries in the G-7, with the exception of Japan. With regard to oil price uncertainty, Ferderer (1996) reports empirical evidence that oil price uncertainty adversely affects US output, while Hooker (1996) reports evidence that this relationship has deteriorated. Elder and Serletis (2008) find that oil price uncertainty has adversely affected US output, investment, consumption and production.[1]

Our primary result is that oil price uncertainty has had a negative and statistically significant effect on output growth in four of seven countries (Canada, France, UK and the US). The results for three countries (Canada, UK and the US) are remarkably robust to various assumptions related to stationarity, as well as simplifications to our baseline VAR with Multivariate GARCH. Impulse-response analysis indicates that both increases in oil prices and decreases in oil prices have tended to reduce output in these countries in the short-run, which is consistent with the mechanism described by Bernanke (1983).

Graphical plots reveal that oil price uncertainty spiked when OPEC collapsed during the mid 1980's, suggesting that oil price uncertainty may have contributed to the stagnant economic growth during this period. More recently, oil price uncertainty was somewhat elevated in 2005, but, through 2007, did not equal the peaks of previous crises. Finally, our results suggest that uncertainty about oil prices has adverse effects for both net oil importers (such as the US and France) and net oil exporters (such as Canada and the UK). The remainder of the paper is structured as follows: Section 2 describes the empirical model. Section 3 discusses the data and issues related to identification. In section 4, we test for stationarity and cointegration and present our empirical results. Section 5 concludes.

2 The Empirical Model

Our empirical model was developed in Elder (1995, 2004), and is based on the VAR of Sims (1980) and structural version of Bernanke (1986), modified to accommodate multivariate GARCH-in-Mean. We assume that the dynamics of the structural system can be summarized by a linear function of the relevant vector of macroeconomic variables, modified to

[1] More generally, Elder (2004), Bredin and Fountas (2005), and Fountas and Karanasos (2007) show that aggregate price uncertainty has tended reduce US industrial production.

permit the conditional volatility of oil to affect the conditional mean;

$$By_t = C + \Gamma_1 y_{t-1} + \Gamma_2 y_{t-2} + ... + \Gamma_p y_{t-p} + \Lambda H_{oil}(t)^{1/2} + \epsilon_t \tag{1}$$

where dim(B) = dim(Γ_i) = (N×N), ϵ_t—$\psi_{t-1} \sim$ iid N($\mathbf{0}$, $\mathbf{H}_t$), $\mathbf{H}_t$ is diagonal, $\mathrm{H}_{oil}(\mathrm{t})^{1/2}$ is the conditional standard deviation of oil and ψ_{t-1} denotes the information set at time *t-1*, which includes variables dated *t-1* and earlier. Λ is a vector of zeros with one free parameter, as described below. We specify the vector $\mathbf{y}_t$ to include a measure of the price level, an index of industrial production, the growth rate in oil prices and a short term interest rate for each country.

This model relaxes two major assumptions in conventional VARs. First, we relax the assumption that the structural disturbances are homoskedastic. Second, we relax the assumption in conventional VARs that excludes the volatility of, say, oil prices from the output equation, by including the parameter matrix Λ. The vector ε_t represents the orthogonalized structural innovations, which are related to the choice of N(N–1)/2 free parameters in the matrix $\mathbf{B}$, with the diagonal elements normalized to one and subject to the condition that $\mathbf{B}$ is of full rank. Our model allows contemporaneous oil price volatility, denoted $\mathrm{H}_{oil}(\mathrm{t})^{1/2}$, to affect output growth by the coefficient matrix Λ. That is, if oil price volatility tends to decrease industrial production, then we would expect the coefficient on the conditional standard deviation of oil in the production equation would be negative and statistically significant.

To capture the clustered volatility typical of financial and macroeconomic time series, we permit the conditional variance matrix $\mathbf{H}_t$ to follow a multivariate GARCH process. Versions of such processes are presented in Bollerslev *et al.* (1988) and Engle and Kroner (1995), although they are too general for most applications, with a very large number of parameters and no assurance that $\mathbf{H}_t$ is positive definite. We address these issues by following Elder (1995, 2004) and taking advantage of the common identifying assumption in structural VARs, that the structural errors are orthogonalized. This implies that the conditional variance matrix $\mathbf{H}_t$ is diagonal, which vastly simplifies the structural variance function. If we also permit each conditional variance to depend on one lag of its own past squared errors and one lag of its own past conditional variances, then the diagonal elements of $\mathbf{H}_t$ can be represented as;

$$\begin{pmatrix} H_{CPI}(t) \\ H_{IP}(t) \\ H_{oil}(t) \\ H_{rate}(t) \end{pmatrix} = \begin{pmatrix} C_1 + F_1\varepsilon_{CPI}(t-1)^2 + G_1 H_{CPI}(t-1) \\ C_2 + F_2\varepsilon_{IP}(t-1)^2 + G_2 H_{IP}(t-1) \\ C_3 + F_3\varepsilon_{oil}(t-1)^2 + G_3 H_{oil}(t-1) \\ C_4 + F_4\varepsilon_{rate}(t-1)^2 + G_4 H_{rate}(t-1) \end{pmatrix} \tag{2}$$

$\mathbf{z}_t \sim$ iid N($\mathbf{0}$, $\mathbf{I}$);

$\varepsilon_t = \mathbf{H}_t^{1/2}\mathbf{z}_t$.

The standard homoskedastic VAR is typically estimated in two-stages, in which the reduced form parameters are estimated by OLS in a first stage, and the structural parameters are recovered in a second stage from the reduced form covariance matrix $\mathbf{B}\varepsilon_t\varepsilon_t\prime\mathbf{B}\prime$ – either by a Cholesky decomposition or, if $\mathbf{B}$ is not triangular, a maximum likelihood procedure over the N(N–1)/2 free parameters. In our model, the information matrix is not block diagonal, so that the parameters cannot be estimated consistently by a comparable estimation procedure. In our model, the conditional mean and conditional variance must be estimated simultaneously in order to obtain consistent estimates of the parameters of interest. In particular, we use the estimation procedure described in Elder (2004), in which the multivariate GARCH-in-Mean VAR can be estimated by full information maximum likelihood by numerically maximize the log likelihood with respect to the structural parameters. We set the pre-sample values of the conditional variance matrix $\mathbf{H}_0$ to their unconditional expectation and condition on the pre-sample values of $\mathbf{y}_t$. To ensure that $\mathbf{H}_t$ is positive definite, we enforce $C_i > 0, F_i \geq 0$ and $G_i \geq 0$. (c.f., Engle and Kroner (1995)). Provided that the standard regularity conditions are satisfied, full information maximum likelihood estimates are asymptotically normal and efficient, with the asymptotic covariance matrix given by the inverse of Fisher's information matrix. The algorithms for estimation and analysis are coded by the authors in Gauss, utilizing the OPTMUM optimization routine.

3 Data and Identification

There exists an extensive VAR literature that relates oil prices to the real economy, including for example, Hamilton (1996), Hooker (1996), Mork (1989), Lee *et al.* (1995), Bernanke *et al.* (1997), Hamilton and Herrera (2004), Edelstein and Kilian (2007a, 2007b) and Elder and Serletis (2008). We use this literature to help guide our empirical specification. An empirical macroeconomic model for each country should include a measure of the aggregate price level, real output, oil prices and a short term interest rate. These variables include the core variables in the existing related literature. For example, Hamilton and Herrera (2004) and Bernanke et al. (1997) use monthly observations on these variables plus commodity prices and other interest rate measures. Hamilton (1996) and Hooker (1996) use these variables plus a measure of import prices. This four variable model appears to represent a reasonable compromise between completeness and parsimony given the complexity of our model.

We measure the price level in each country by the domestic consumer price index. Oil prices are measured by the free on board cost of the imported oil, expressed in local currency. Following Blanchard and Gali (2007), we use the nominal price of oil in local currency rather than the theoretically important real price of oil, in order to avoid dividing by an endogenous variable. This also allows us to isolate uncertainty associated with oil prices from uncertainty associated with the aggregate price level. Since we are interested in the effects of oil price uncertainty on energy intensive sectors such as manufacturing activity, we measure output in each country by the domestic index of industrial production.

Our data sample is monthly from 1974:01-2007:10, including the pre-sample observations. Figures 2 and 3 plot the industrial production growth rates against the oil price for each country, with shading representing recessions as indicated by the National Bureau of Economic Research for the US and by the Economic Cycle Research Institute for the remaining six countries. These figures indicate that oil prices, denominated in local currency, appear to move in a similar fashion in each of the G-7 countries. In each of these countries, oil prices rose dramatically in the late 1970s, dropped dramatically in 1985, and stabilized from the mid 1980's to about 1999. Since that time, oil prices have increased substantively. Our baseline model therefore consists of a four variable VAR on *CPI*, *IP*, the price of oil expressed in domestic currency and a domestic short-term interest rates for each country. With regard to identification, we allow $\mathbf{B}$ to be lower triangular with the following ordering: inflation, the growth rate in industrial production, the growth rate in oil prices and the interest rate. These identifying restrictions are broadly consistent with the identified VAR literature, including Hamilton and Herrera (2004) and Bernanke *et al.* (1997).

4 Empirical Evidence

Our multivariate GARCH-in-Mean VAR is estimated using monthly data for the 1974:1 to 2007:10 period, including pre-sample observations, for the G-7 countries. To determine the appropriate variable transformations, we first conduct tests for unit roots and cointegration. Table 1 reports the results of augmented Dickey-Fuller (ADF) tests for unit roots and cointegration, conducted in the manner described by Elder and Kennedy (2001a and 2001b). We initially take the log of each series to remove possible exponential growth. For the log series that appear to exhibit a trend, which include $\log(CPI)$ for some countries and $\log(IP)$, we estimate the following univariate equation by OLS, with the lag length chosen by minimizing the Schwartz information criteria (SIC);

$$y_t = \alpha + \xi_1 \Delta y_{t-1} + \ldots + \xi_p \Delta y_{t-p} + \rho y_{t-1} + \delta t + \varepsilon_t \tag{3}$$

Two common ADF test statistics based on this estimation equation are the OLS t-test with the null of $\rho = 1$, denoted τ_τ, and the OLS F-statistic based on the joint null hypothesis of $\rho = 1$ and $\delta = 0$, denoted Φ_3, both of which have non-standard distributions. The latter test, Φ_3, has greater power, and is motivated by the observation that if the null of the unit root is accepted, then the trend should be zero, to rule out explosive growth.

As reported in table 1, for log(*IP*) the joint null hypothesis of a unit root and no trend is not rejected for each country. We therefore use the first difference of the log(*IP*) in our empirical model. For log(*CPI*), however, the joint null is rejected for Canada, Italy, Japan and United Kingdom. To further investigate the nature of this rejection, we plot the raw *CPI*, the log(*CPI*) and year-over-year growth rate of the *CPI* for these four countries. As an illustration, these plots for the UK are reported in Figure 1. Examination of these plots confirms that during our sample, each of these countries underwent considerable disinflation. As a consequence, the raw *CPI* series does not display the usual exponential growth, so that the log transformation introduces a noticeable convexity. Such a process is not described well by either a time trend or a unit root. To further investigate whether this convexity affects our unit root tests, we conduct the Φ_3 tests on the untransformed *CPI* series, which does not display the convexity, and do not reject the joint null of a unit root with no trend for each of these four countries. This suggests that we should model the untransformed *CPI*, rather than the log(*CPI*) series, as difference stationary. The differenced log(*CPI*) series, however, has the more intuitive interpretation as the continuously compounded inflation rate. We therefore estimate our model with two transformations of the *CPI* to ensure robustness and consistency across countries, the log(*CPI*) series in first differences and the raw *CPI* series in first differences.

The *Rate* series does not exhibit a clear trend, so we estimate the following univariate equation by OLS, with the lag length again chosen by minimizing the SIC;

$$y_t = \alpha + \xi_1 \Delta y_{t-1} + \ldots + \xi_p \Delta y_{t-p} + \rho y_{t-1} + \varepsilon_t \tag{4}$$

The common ADF test statistics based on this estimation equation are the OLS t-test with the null of $\rho = 1$, denoted τ_μ, and the F-statistic based on the joint null hypothesis of $\rho = 1$ and $\alpha = 0$, denoted Φ_1. The Φ_1 test is motivated by the observation that, under the null of a unit root, the drift term α should be zero, since a trend has been ruled out, and so this test should be expected to have greater power for large values of α. Elder and Kennedy (2001a) showed, however, that the Φ_1 test actually has less power than τ_μ, due to the invariance of this test statistic with respect the value of α. We therefore report the τ_μ test for *Rate* in table 1. Note that, for each country expect Germany, the null hypothesis of a unit root is not rejected. Many authors, however, have strong priors on the stationarity of short-term interest rates, the arguments for which were advanced most forcefully by Cochrane (1991). Bernanke and Blinder (1992, p 906, footnote 12) note simply that differencing the interest rate in such models is "not very sensible." Given these issues, we estimate our model with *Rate* both in levels and in first differences.

We next test for cointegration in two-variable pairs. To discount the probability of imposing a spurious cointegrating relationship, we test only for cointegrating relationships that are likely to be justified by economic theory. Such relationships include a Fischer effect, which posits cointegration between interest rates and either *CPI* or log(*CPI*), whichever is most appropriate based on the ADF tests, with a known cointegrating vector of (1, –1). We test this hypothesis by applying the ADF τ_τ test to the variable

$$z_t = log(CPI_t) - Rate_t$$

for France, Germany and the US and

$$z_t = CPI_t - Rate_t$$

for Canada, Italy, Japan and the UK We use the ADF τ_τ so as not to impose additional structure regarding the trend. For each country, the null hypothesis of a unit root in z_t is not rejected. Hence, we conclude that there is no cointegrating relationship between prices and *Rate* in any of the countries for our sample.

Based on the above discussion, our baseline model includes the log(*CPI*), log(*IP*) and log(*Oil*) in first differences, so these variables are interpreted as continuously compounded growth rates. We include the interest rate in levels. These are comparable to the transformations applied by, for example, Lee, Ni and Ratti (1995). However, to ensure that our results are not driven by our assumptions regarding stationarity, we also estimate our model with interest rates in first differences for all countries and the *CPI* in first differences for Canada, Italy, Japan and the UK Such transformations have very little effect on the relationship between oil price uncertainty and industrial production, as we describe below. Since our model is designed to capture short-run effects of oil prices, we include six lags in all our VARs, which is appropriate on the basis of sequential likelihood ratio tests. To examine whether our model captures important features of the data, we calculate the SIC for our MGARCH-in-Mean VAR, and two nested models; a homoskedastic VAR and a Multivariate ARCH VAR. The results are reported in table 2, and they clearly show that for each country the MGARCH-in-Mean VAR is the preferred specification.

Tables 3A, 3B and 3C report the coefficients of the conditional variance equations for each of the four variables for the G-7 countries and reveal a number of interesting results. First, in the majority of cases, both the ARCH and GARCH coefficients are statistically significant, thus supporting the conclusions of table 2. Second, in most cases we find very high persistence in the volatility of inflation, output, oil price and interest rates, based on the sum of the ARCH and GARCH coefficients.[2] The primary coefficient of interest, the coefficient on oil price uncertainty in the industrial production equation, is reported in table 4. This coefficient is negative for all G-7 countries. Moreover, it is statistically significant at the 5% level in four of the G-7 countries, namely, Canada, France, UK and US. The US result is consistent with Elder and Serletis (2008) who find a negative effect of oil price uncertainty on several measures of US real economic activity. Our results indicate that the net oil exporter nature of both Canada and the UK does little to limit the exposure to the contractionary effects of oil price uncertainty. The lack of a significant effect for the case of Italy and Germany may be due to the offsetting effects played by real effective exchange rate depreciation. Finally, the implications of structural reform and the relatively less reliance on oil for the case of Japan may explain the lack of statistically significance

[2]Exceptions are the conditional variance of inflation for France, Germany and UK, and the conditional variance of output for Germany, Japan and the UK.

for the uncertainty term.[3]

In figures 4 and 5 we plot the estimated conditional standard deviation of oil against the growth rate of industrial production for each of the G-7 countries. Several observations can be made based on these plots. First, oil price uncertainty was quite high in the mid 1980s, during a period of rapidly declining oil prices, and the early 1990s, during the rapid oil price increases just prior to the Gulf War. Figures 4 and 5 illustrate the concomitant stagnant or falling output growth in a number of cases (most notably Canada, UK, US) which is consistent with our empirical results in table 4. Second, the persistent increases in oil price since 2003 have been accompanied by only modest increases in oil price uncertainty in early 2005. The failure of these oil price increases to generate sustained uncertainty may be one reason why a recession had not materialized by the end of 2007.

4.1 Robustness

We investigate the robustness of our results by estimating numerous alternative specifications, in particular we investigate alternative assumptions related to appropriate transformations required for the VAR to be stationary, as well as simplifications to our simultaneous equations model. As discussed previously, for most countries the null of a unit root in the interest rate cannot be rejected. Despite this, monetary VARs are typically specified with the interest rate in levels even if the other variables are differenced. To investigate whether our results regarding the relationship between oil price uncertainty and industrial production are affected by our inclusion of the interest rate in levels, we re-estimate the MGARCH-in-Mean VAR for each country with the interest rate in first differences. This transformation has very little effect on our results, as the coefficient on oil price uncertainty is again negative and significant for Canada, UK and the US. For France, the statistical significance of oil price uncertainty in the *IP* equation declines modestly, as the absolute asymptotic t-statistics drops from 2.14 to 1.78. This coefficient remains significant, however, at the

[3]Similar results in relation to oil price shocks have been reported by Mork (1994) and Jimenez-Rodriguez and Sanchez (2005).

10% level.[4] We also investigate whether our model is sensitive to the transformation applied to the *CPI*. In our baseline model we differenced the log of the *CPI* for all countries, even though the *CPI* for Canada, Italy, Japan and the UK did not exhibit exponential growth over our sample. We re-estimate the model without the log transformation, so that the raw *CPI* is in first differences, and again confirm our previous finding that the effect of oil uncertainty on industrial production is negative and significant.[5]

4.2 Additional Tests of Robustness

Our results thus far clearly illustrate that oil price uncertainty has had a negative and statistically significant effect on production in Canada, France, US and UK We next consider additional tests of robustness related to the our empirical model. That is, our empirical model has many desirable features, but the complexity of a simultaneous equations model with multivariate GARCH may overshadow the robustness of our empirical result. To investigate whether our measure of oil price volatility is significant in a simple linear regression of industrial production on lagged inflation, industrial production, oil prices and interests rates, we estimate the following regression, in which the variables have the same transformation as in our baseline model

$$IP_t = c + \sum_{j=1}^{6} \beta_{1,j} CPI_{t-j} + \sum_{j=1}^{6} \beta_{2,j} IP_{t-j} + \sum_{j=1}^{6} \beta_{3,j} Oil_{t-j} + \sum_{j=1}^{6} \beta_{4,j} Rate_{t-j} + \Lambda \hat{H}_{oil}(t) + \varepsilon_t \quad (5)$$

where $\hat{H}_{oil}(t)$ is the measure of oil price uncertainty extracted from our baseline structural VAR with multivariate GARCH. Note that in equation (5), $\hat{H}_{oil}(t)$ is a generated regressor, similar to that examined by Pagan (1984). If the data generating process is properly specified, then the coefficient Λ can be estimated consistently by OLS and the OLS standard error is also consistent, under the null hypothesis that the coefficient is zero. Under the alternative hypothesis that the coefficient is non-zero, the OLS standard

[4]The results are not reported, but are available from the authors upon request.

[5]The results are not reported, but are available from the authors upon request.

error is not consistent. Of course, generated regressors can be addressed by simultaneous estimation, but that is precisely the issue we are attempting to abstract from with this exercise.

We therefore estimate equation (5) by OLS, with the results reported in table 5. These results confirm our previous finding that the coefficient on oil price volatility is negative for all G-7 countries and is negative and statistically significant at the 5% level for Canada, France, the UK and the US. In addition, the coefficient estimates for these four countries are of roughly similar magnitude to those reported in table 4.

To further investigate the robustness of our results, we simplify our model by producing a new estimate of oil price uncertainty, from a simple univariate GARCH model. In particular, we estimate the univariate GARCH(1,1) model for each country

$$Oil_t = c + \sum_{j=1}^{6} \beta_{3,j} Oil_{t-j} + \varepsilon_t$$

where $\varepsilon_t \sim N\left(0, H_{oil}(t)^*\right)$ and $H_{oil}(t)^*$ follows a simple univariate GARCH(1,1) process. We then reestimate equation (5) for each country with $\hat{H}_{oil}(t)^*$ as our measure of oil price uncertainty. Our results are again surprisingly robust, the coefficient on $\hat{H}_{oil}(t)^*$ is negative and significant at the 5% level for Canada, the US and the UK For France, the coefficient is again negative, but the p-value falls to 0.15.[6]

4.3 Impulse-Response Analysis

The coefficient on oil price uncertainty in the output equation indicates that oil price uncertainty tends to be associated with lower production in four of the seven industrialized countries in our sample. In our empirical model, both positive oil shocks (higher oil prices) and negative oil shocks (lower oil prices) tend to increase oil price uncertainty, so this channel suggests that the effects of negative oil shocks will not mirror the effects the positive oil shocks. Standard economic theory, such as Kim and Loungani (1992), suggests that

[6]The results are not reported, but are available from the authors upon request.

negative oil shocks affect should be expansionary, even in the short-run. Whether the net short-run effect of negative oil shocks is contractionary or expansionary depends on whether the effects of oil price uncertainty outweigh the effects from other channels.

We can calculate this net short-run effect in our empirical model by simulating the response of output to an orthogonalized oil shock, accounting for the contemporaneous and lagged responses of interest rates and inflation, in a manner analogous to such simulations for conventional VARs. Elder (2003) describes how such impulse-response functions can be calculated for the structural VAR with MGARCH-in-Mean estimated in this paper, and Elder (2004) describes how to estimate one standard errors bands by Monte-Carlo methods. We therefore simulate the response of production to both positive and negative oil price shocks for each of the four countries in which the effect of oil price uncertainty is statistically significant. The absolute magnitude of the initial shock is one unconditional standard deviation of the growth rate in oil prices.

The response of production to a positive oil shock for Canada, France, UK and the US are reported in Figure 6. These impulse-responses indicate that higher oil prices tend to reduce production significantly after one or two months for Canada, the UK and the US - which is consistent with the effects predicted by standard economic theory. The effect of positive oil shocks in France is initially negative, although not significant. Note that by incorporating the effects of oil price uncertainty, the response of output to positive oil price shocks in the US in unambiguously negative. This apparently resolves the puzzle noted by Hamilton (1996), who finds that the VAR response of output to oil shocks is not significant in post-1973 samples, when the effects of oil price uncertainty are not explicitly accounted for.

We next examine the short-run response of production to a negative oil shock. Bernanke (1983) suggests that negative oil shocks (i.e., lower oil prices) may not be expansionary in the short-run if the oil shock creates uncertainty about future oil prices. The impulse-responses from our VAR with MGARCH, which is designed to capture such short-run effects, indicates that for Canada, UK and the US, a negative oil shock causes production to contract significantly for one to three months. The decline is similar in magnitude for each of three countries, before fading after two or three months. For France the effect of a negative oil shock is initially negative, but not statistically significant.

Our results therefore indicate that the effects of oil price uncertainty is significant, and that, in the short-run, both positive and negative oil shocks may be contractionary. In particular, the short-run effect of oil price uncertainty effect are sufficiently large that even falling oil prices may be contractionary.

5 Conclusion

The failure of decreases in oil prices to produce expansions that mirror the contractions associated with higher oil prices has long been of interest to researchers. In this paper, we investigate one prominent explanation for this feature, the role of uncertainty about oil prices. In particular, we examine the link between oil price uncertainty and industrial production in the G-7 countries utilizing a very general and flexible empirical methodology that is based on a structural VAR modified to accommodate multivariate GARCH in mean errors. Our primary result is that oil price uncertainty has had a negative and significant effect on industrial production in four of the G-7 countries, Canada, France, the UK and the US. Our result is robust to numerous assumptions regarding stationary and model specification, including substantive simplifications to our base-line model.

Given our measure of oil price uncertainty, our result helps explain why the steady but slow increases in oil prices from 2003-2006 failed to induce recessions in the G-7, and why dramatic decreases in oil prices may, in the short-run, be contractionary. Our result also indicates that the apparent asymmetry in the response of output to oil prices is a feature common to several industrialized economies. The adverse effects of oil price uncertainty irrespective of whether countries are net oil exporters or importers, is also evident. In particular, Canada and the UK, both net oil exporters are particularly sensitive to higher levels of oil price uncertainty. Finally, our result suggests that the dramatic increase in the variability of oil prices observed since early 2008 is likely to have adverse effects of world-wide economic growth.

6 Appendix: Data Description

Series	Transformations	Description
CPI	12*ln(CPI_t/CPI_{t-1})	Consumer Price Index
	$CPI_t - CPI_{t-1}$	Alternative transformation
Output	12*ln(IP_t/IP_{t-1})	Industrial Production, seasonally adjusted
Oil	12*ln(Oil_t/Oil_{t-1})	F.O.B. cost of imported crude oil, in local currency
Rate	None in baseline	Short term interest rate. Treasury bill rate rate is used for France and Canada. The Call Money rate is used for Germany and Japan. The money market rate is used for Italy and the overnight interbank rate for the UK. The Federal Funds rate is used for the US.
	$Rate_t - Rate_{t-1}$	Alternative transformation

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Table 1: **Tests for Stationarity**

Series	log(CPI)	CPI	log(IP)	log(Oil)	Rate	Cointergration
ADF test	Φ_3	Φ_3	Φ_3	τ_μ	τ_μ	τ_τ
5% Critical value	6.30	6.30	6.30	-2.87	-2.87	-3.42
Canada	93.84**	5.96	4.12	-2.34	-2.17	-1.68
France	5.91	2.61	3.91	-2.53	-1.65	-2.61
Germany	3.57	4.78	2.97	-1.26	-3.16**	-3.17
Italy	31.91**	2.79	2.98	-2.46	-1.63	-2.29
Japan	8.65**	5.09	1.86	-1.65	-2.40	-2.58
UK	8.73**	4.36	2.02	-2.73	-2.57	-3.05
US	2.89	2.95	4.56	-2.24	-2.24	-2.94

Note: This table reports ADF tests for a unit root. The Φ_3 test is based on equation (3) in the text with the joint $H_0 : \rho = 1$ and $\delta = 0$. The τ_μ test is based on equation (4) in text with H_0: $\rho = 1$. The τ_τ test is based on equation (3) in text with H_0: $\rho = 1$. The last column reports the cointegration results between Rate and log(CPI) or CPI.

Table 2: **Model Specification Tests**

	Schwarz Criterion Values		
Series	VAR	Multivariate ARCH VAR	Multivariate GARCH-M VAR
Canada	208	-30	-200
France	555	-360	-420
Germany	99	-251	-414
Italy	555	180	-175
Japan	110	-89	-397
UK	1221	877	636
US	-782	-1239	-1569

Note: These are the Schwarz criterion values for the estimated VAR, where 'VAR' refers to the homoskedastic VAR, the Multivariate ARCH-VAR given by equations (1) and (2) with $\varepsilon_t \sim \mathrm{N}(\mathbf{0}, \mathbf{H}_t)$ and $\mathrm{G}_1 = 0$, and the Multivariate GARCH-M VAR given by equations (1) and (2) with the diagonal elements of F_1 and G_1 unrestricted.

Table 3a

Coefficient Estimates for the Variance Function of the MGARCH-M VAR

Equation	Con. Var.	Constant	$\varepsilon_i(t-1)^2$	$H_{i,i}(t-1)$
Canada				
Infl	$H_{1,1}(t)$	0.000*	0.211**	0.635**
		(1.99)	(2.96)	(5.01)
Output	$H_{2,2}(t)$	0.000	0.121**	0.869**
		(1.15)	(3.91)	(24.02)
Oil	$H_{3,3}(t)$	0.005	0.253**	0.732**
		(1.79)	(5.14)	(14.49)
Rate	$H_{4,4}(t)$	0.002	0.167**	0.823**
		(1.05)	(5.88)	(24.91)
Germany				
Infl	$H_{1,1}(t)$	0.001**	0.524**	0.062
		(5.62)	(3.99)	(0.649)
Output	$H_{2,2}(t)$	0.022**	0.132**	0.142
		(4.11)	(1.94)	(0.77)
Oil	$H_{3,3}(t)$	0.014*	0.197**	0.778**
		(1.93)	(5.18)	(19.33)
Rate	$H_{4,4}(t)$	0.001**	0.237**	0.754**
		(3.75)	(7.54)	(25.33)
Japan				
Infl	$H_{1,1}(t)$	0.000**	0.050**	0.940**
		(2.14)	(3.02)	(100.19)
Output	$H_{2,2}(t)$	0.019**	0.137	0.000
		(5.63)	(1.62)	(0.00)
Oil	$H_{3,3}(t)$	0.017*	0.266**	0.713**
		(2.11)	(5.55)	(14.81)
Rate	$H_{4,4}(t)$	0.001**	0.145**	0.845**
		(3.75)	(7.52)	(46.95)

Note: These are the parameter estimates for the free elements in **F** and **G** from the model given by equations (1) and (2) with $\varepsilon_t \sim N(0, H_t)$. Each row in the table represents an equation from the associated bivariate GARCH-in-Mean VAR. Asymptotic t-statistics are in parentheses. A coefficient of 0.000 indicates that the nonnegativity constraint is binding.

** and * denotes significance at the 5% and 10% level respectively.

Table 3b

Coefficient Estimates for the Variance Function of the MGARCH-M VAR

Equation	Con. Var.	Constant	$\varepsilon_i(t-1)^2$	$H_{i,i}(t-1)$
France				
Infl	$H_{1,1}(t)$	0.001**	0.00	0.00
		(5.08)	(0.00)	(0.00)
Output	$H_{2,2}(t)$	0.001**	0.056**	0.934**
		(5.08)	(2.84)	(53.97)
Oil	$H_{3,3}(t)$	0.006	0.194**	0.794**
		(0.75)	(5.17)	(19.03)
Rate	$H_{4,4}(t)$	0.054**	0.984**	0.001
		(7.14)	(31.63)	(0.03)
Italy				
Infl	$H_{1,1}(t)$	0.001**	0.139**	0.851**
		(3.84)	(5.51)	(35.79)
Output	$H_{2,2}(t)$	0.001**	0.081**	0.909**
		(5.11)	(4.03)	(51.57)
Oil	$H_{3,3}(t)$	0.014*	0.238**	0.743**
		(1.93)	(4.60)	(14.00)
Rate	$H_{4,4}(t)$	0.005	0.466**	0.524**
		(0.84)	(15.82)	(5.48)
UK				
Infl	$H_{1,1}(t)$	0.002**	0.913**	0.000
		(10.15)	(6.64)	(0.00)
Output	$H_{2,2}(t)$	0.006**	0.952**	0.000
		(4.74)	(4.79)	(0.00)
Oil	$H_{3,3}(t)$	0.003	0.180**	0.808**
		(1.32)	(3.14)	(22.74)
Rate	$H_{4,4}(t)$	0.002	0.209**	0.781**
		(1.41)	(4.65)	(31.38)

Note: These are the parameter estimates for the free elements in **F** and **G** from the model given by equations (1) and (2) with $\varepsilon_t \sim N(0, H_t)$. Each row in the table represents an equation from the associated bivariate GARCH-in-Mean VAR. Asymptotic t-statistics are in parentheses. A coefficient of 0.000 indicates that the nonnegativity constraint is binding.

** and * denotes significance at the 5% and 10% level respectively.

Table 3c

Coefficient Estimates for the Variance Function of the MGARCH-M VAR

Equation	Con. Var.	Constant	$\varepsilon_i(t-1)^2$	$H_{i,i}(t-1)$
US				
Infl	$H_{1,1}(t)$	0.001	0.095**	0.893**
		(1.17)	(3.08)	(23.23)
Output	$H_{2,2}(t)$	0.002**	0.346**	0.351**
		(3.66)	(3.66)	(2.78)
Oil	$H_{3,3}(t)$	0.003**	0.221**	0.769**
		(2.20)	(8.15)	(29.17)
Rate	$H_{4,4}(t)$	0.003	0.355**	0.635**
		(1.63)	(8.15)	(13.02)

Note: These are the parameter estimates for the free elements in **F** and **G** from the model given by equations (1) and (2) with $\varepsilon_t \sim N(0, H_t)$. Each row in the table represents an equation from the associated bivariate GARCH-in-Mean VAR. Asymptotic t-statistics are in parentheses. A coefficient of 0.000 indicates that the nonnegativity constraint is binding.

** and * denotes significance at the 5% and 10% level respectively.

Table 4

Coefficient Estimates on Oil Volatility in the IP Equation

Model	Variables	Sample	Coefficient on $H_{oil}(t)^1/2$ Oil Volatility
Canada	*CPI,IP,Oil,Rate*	1974:01-2007:07	-0.065**
			(2.99)
France	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.048**
			(2.14)
Germany	*CPI,IP,Oil,Rate*	1974:01-2007:07	-0.015
			(0.61)
Italy	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.026
			(1.02)
Japan	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.013
			(0.59)
UK	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.077**
			(4.87)
US	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.038**
			(3.37)

Note: These are the parameter estimates for Λ from the structural VAR with multivariate GARCH model given by equations (1) and (2). $H_{oil}(t)^{1/2}$ denotes the conditional standard deviation of the oil price, measured in the domestic currency. Absolute asymptotic t-statistics are in parentheses. Variable transformations are described in Appendix A1.

** and * denotes significance at the 5% and 10% level respectively.

Table 5

Coefficient Estimates on Oil Volatility in OLS Regression

Model	Variables	Sample (including pre-sample)	Coefficient on $H_{oil}(t)^1/2$ Oil Volatility
Canada	*CPI,IP,Oil,Rate*	1974:01-2007:07	-1.619**
			(3.20)
France	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.058**
			(2.26)
Germany	*CPI,IP,Oil,Rate*	1974:01-2007:07	-0.031
			(0.93)
Italy	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.051
			(1.36)
Japan	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.013
			(0.50)
UK	*CPI,IP,Oil,Rate*	1974:01-2007:10	-1.50**
			(3.20)
US	*CPI,IP,Oil,Rate*	1974:01-2007:10	-0.038**
			(2.92)

Note: These are the parameter estimates on oil price volatility in a regression of the growth rate of industrial production on lagged CPI, lagged IP, lagged Oil, lagged Rate and oil price volatility as given by equation (5). $H_{oil}(t)^{1/2}$ denotes the conditional standard deviation of the oil price, measured in the domestic currency. Absolute asymptotic t-statistics are in parentheses. Variable transformations are described in Appendix A1.

** and * denotes significance at the 5% and 10% level respectively.

Fig. 1. Industrial Production and Oil Prices

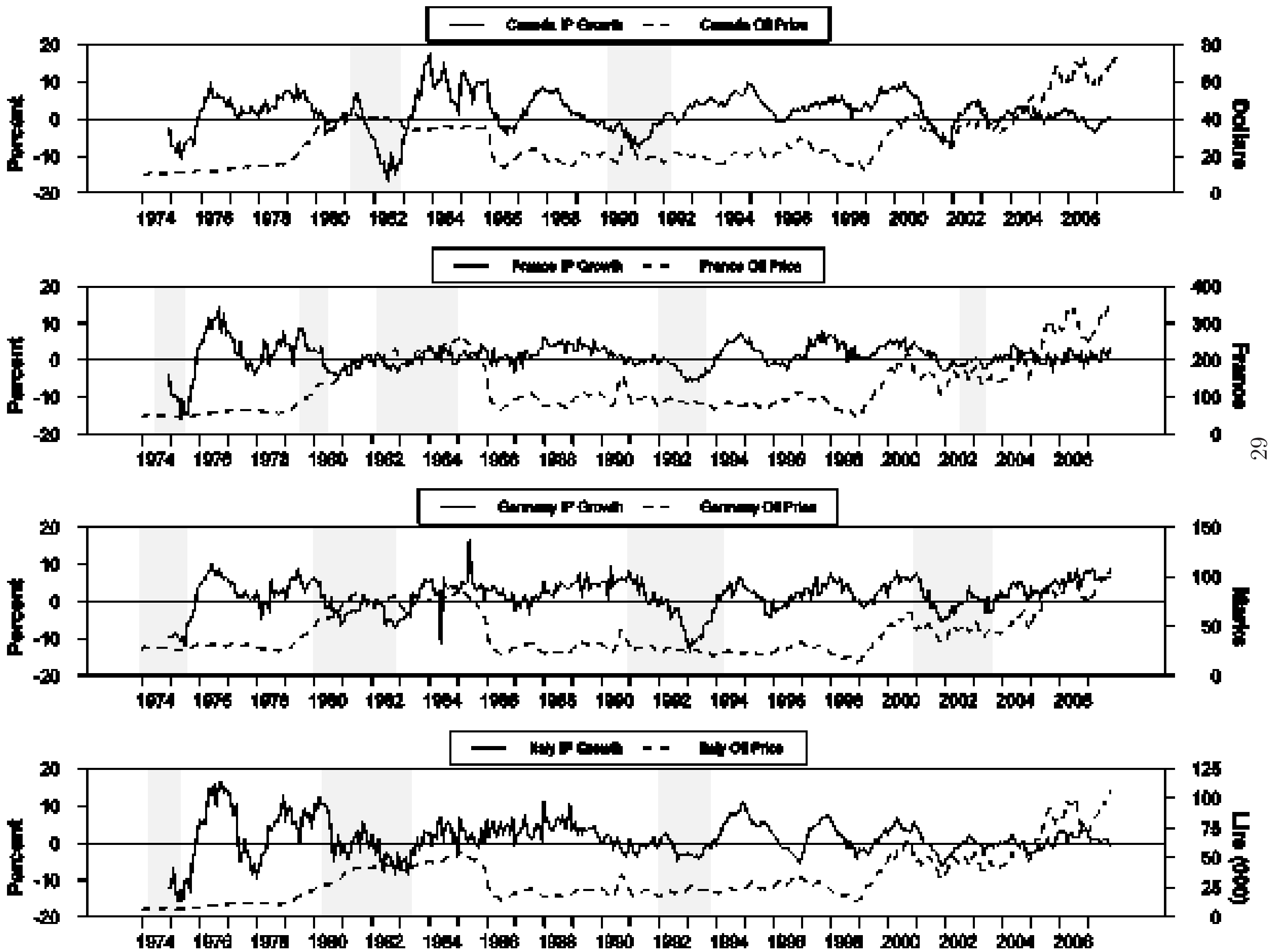

Fig. 2. Industrial Production and Oil Prices

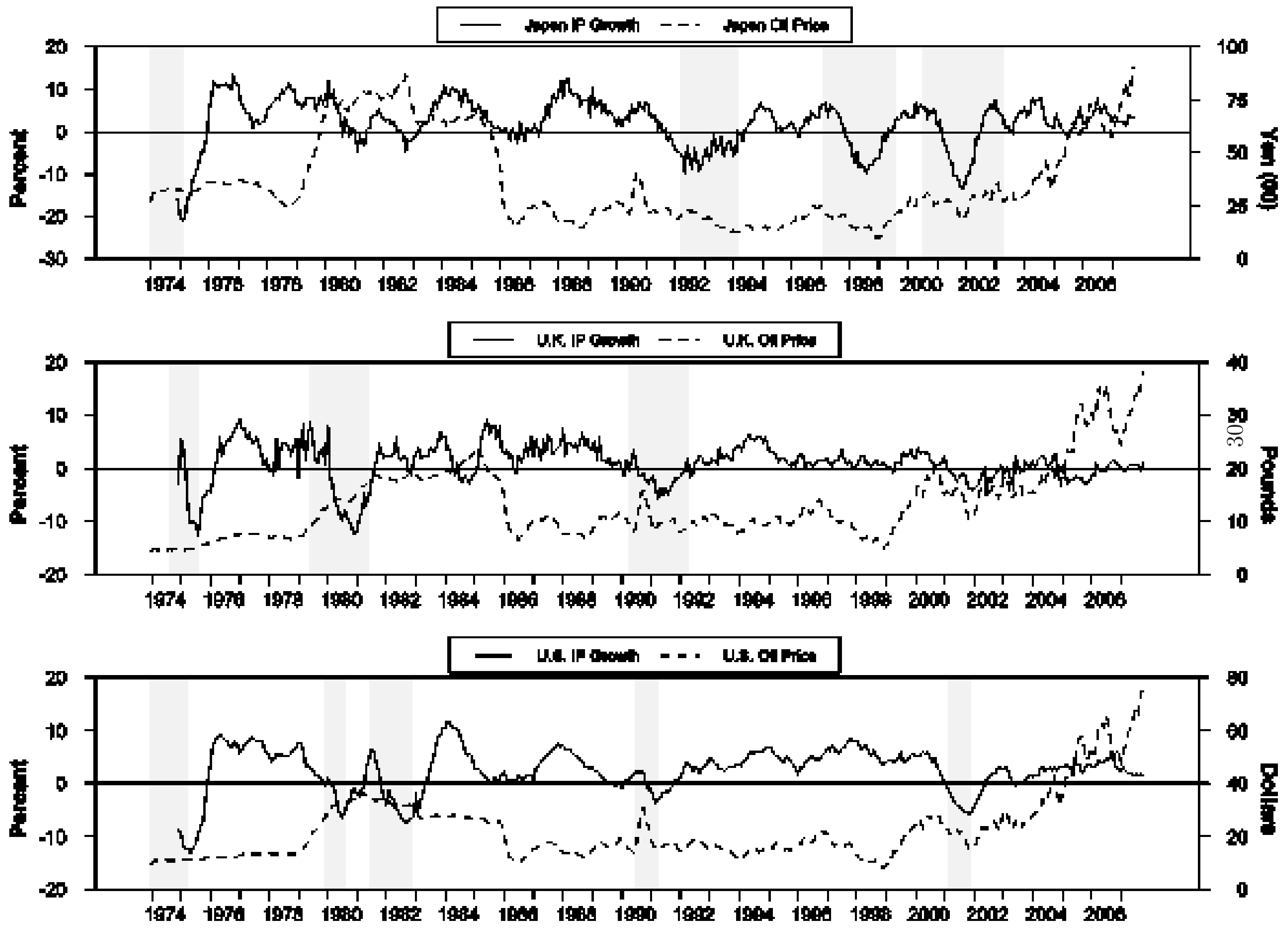

Fig. 3. UK CPI Transformations

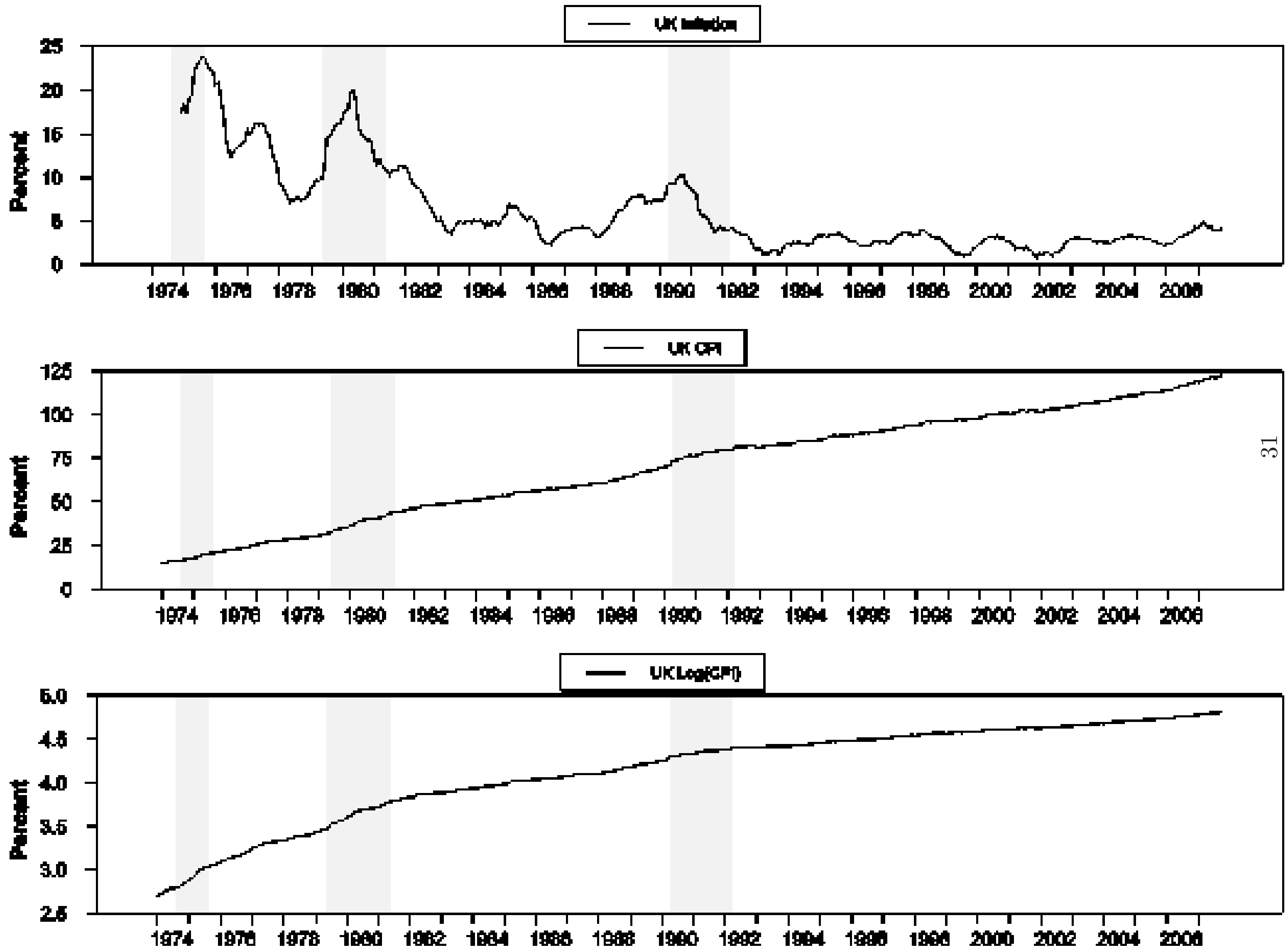

Fig. 4. Industrial Production and Oil Volatility

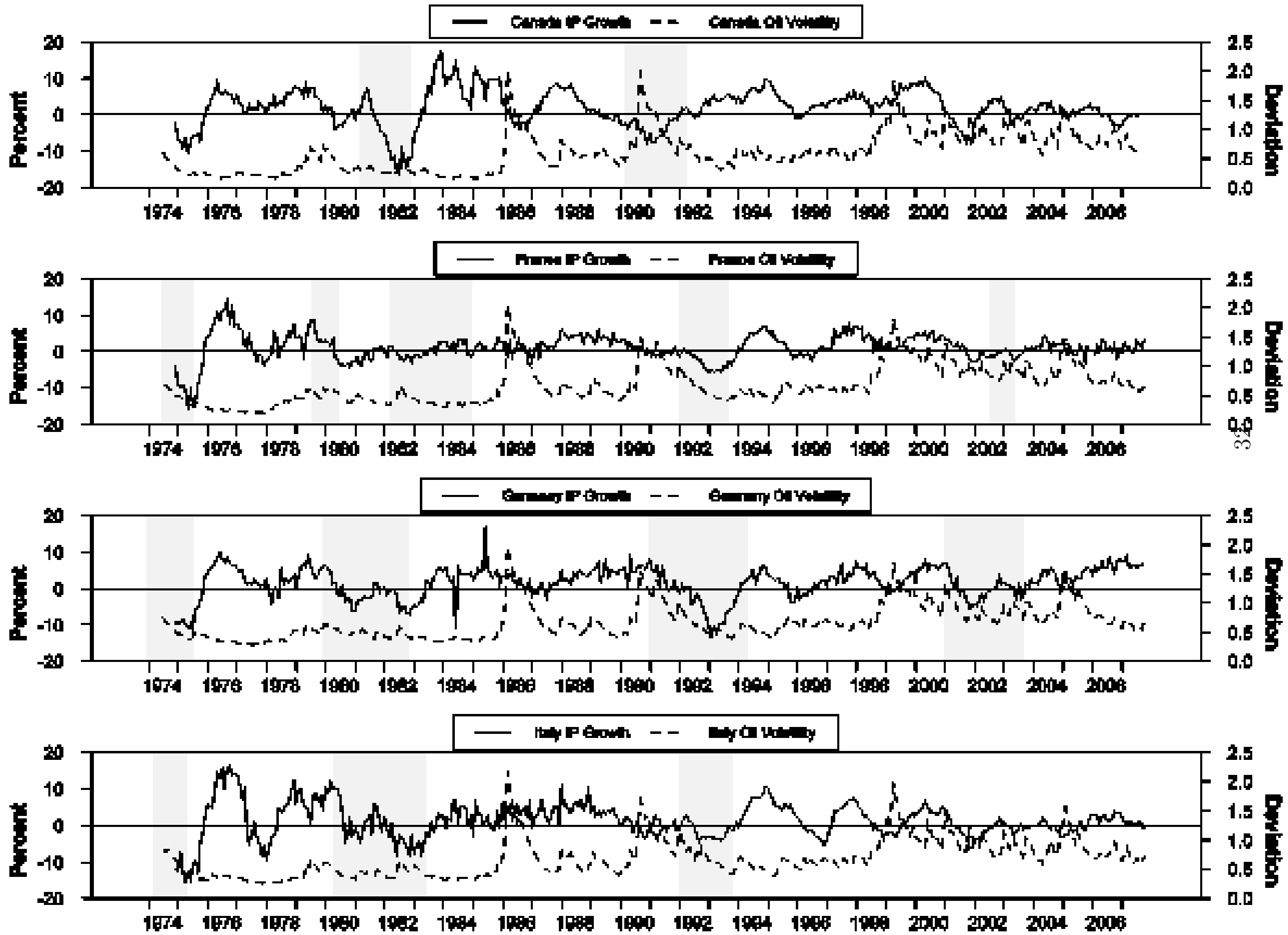

Fig. 5. Industrial Production and Oil Volatility

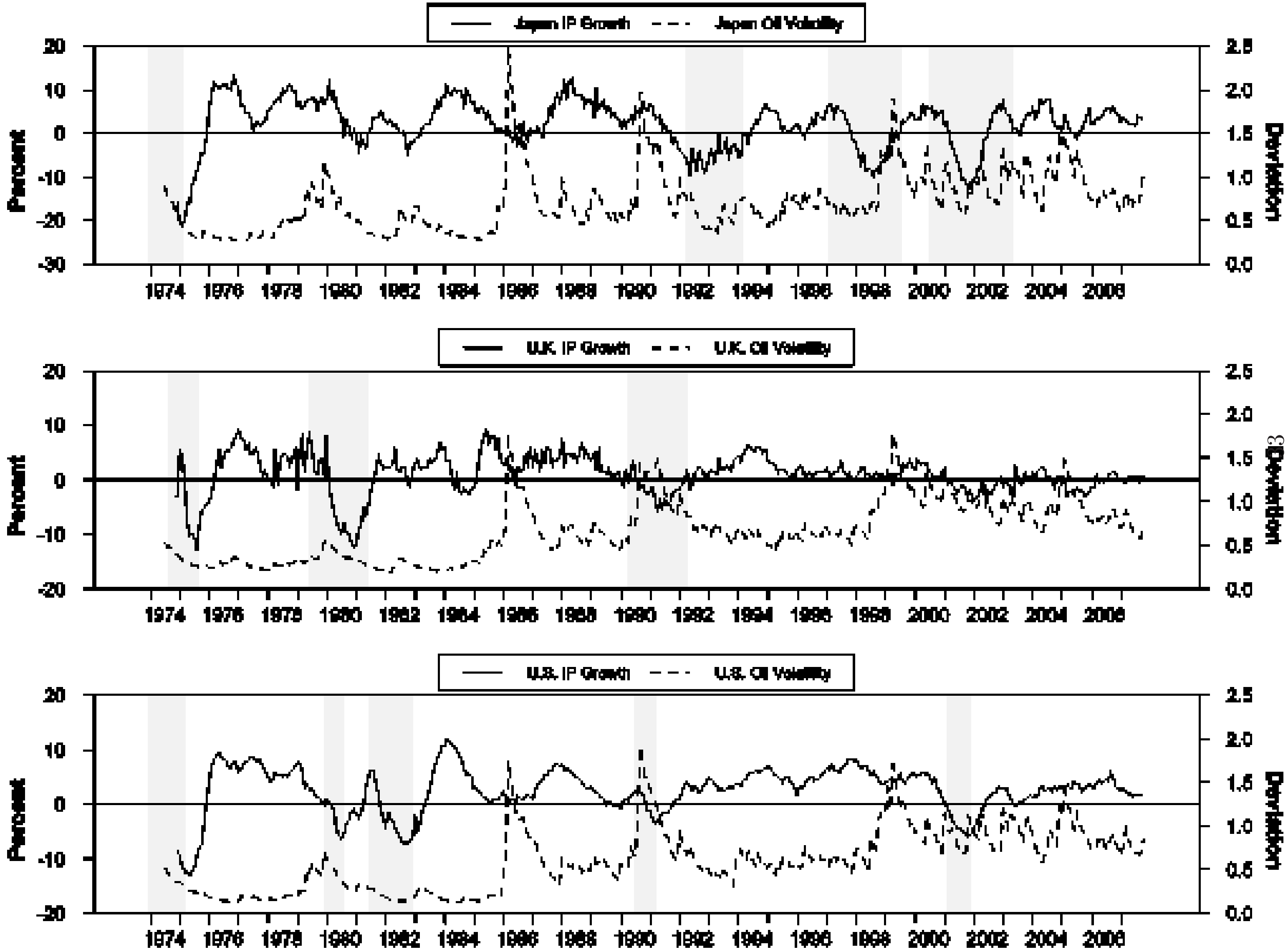

Figure 6. Response of Production to Positive Oil Shock

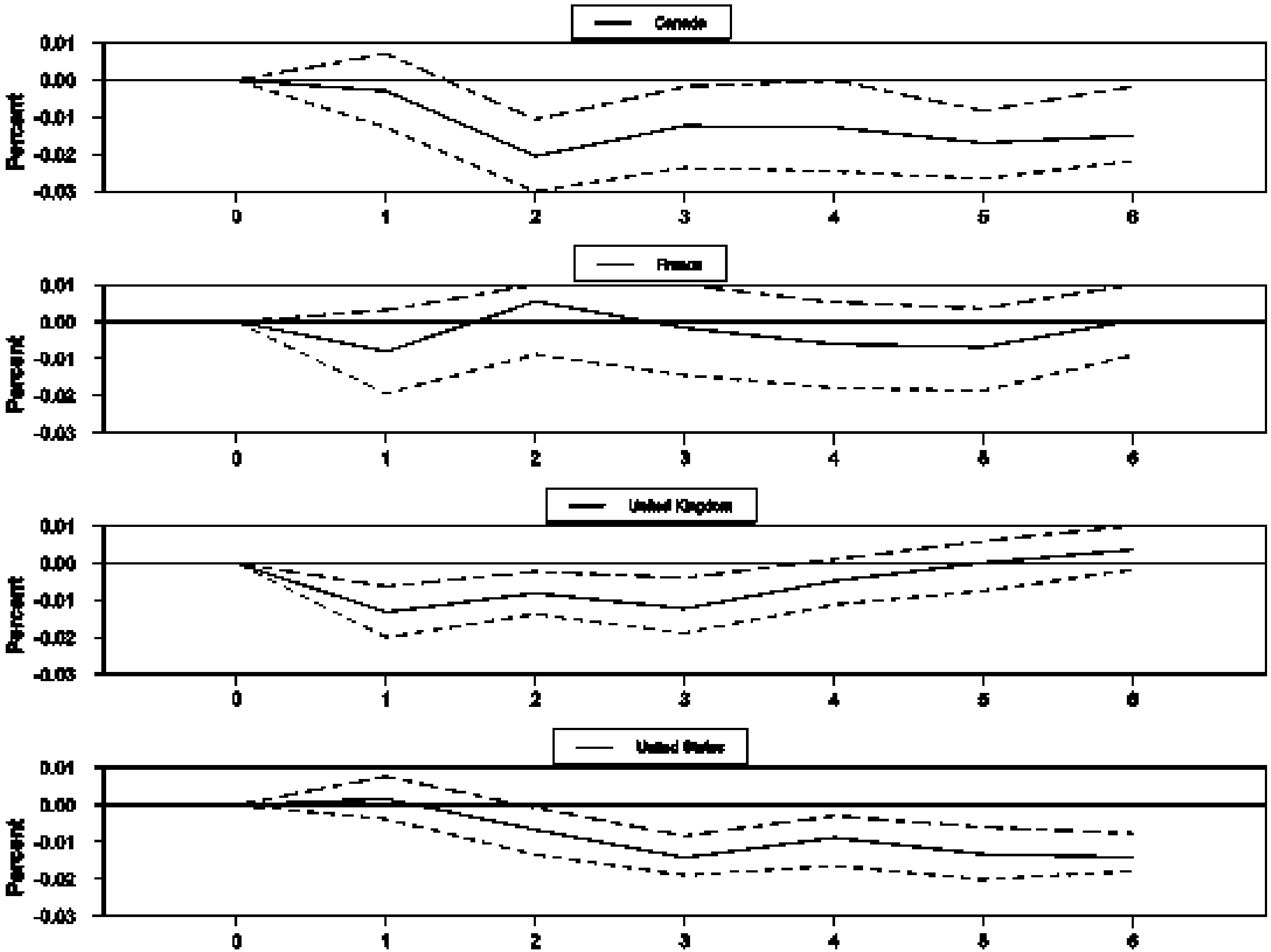

Figure 7. Response of Production to Negative Oil Shock