Oil Volatility and the Option Value of Waiting:
An analysis of the G-7

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Abstract

There has recently been considerable interest in the potential adverse effects associated with excessive uncertainty in energy futures markets. Theoretical models of investment under uncertainty predict that increased uncertainty will tend to induce firms to delay investment. These models are widely utilized in capital budgeting decisions, particularly in the energy sector. There is relatively little empirical evidence, however, on whether such channels have industry-wide effects. Using a sample of G7 countries we examine whether uncertainty about a prominent commodity – oil – affects the time series variation in manufacturing activity. Our primary result is consistent with the predictions of real options theory – uncertainty about oil prices has had a negative and significant effect on manufacturing activity in Canada, France, UK and US.

Key Words: Oil, Volatility, Vector autoregression, Multivariate GARCH-in-Mean VAR.

JEL Classification: G1, G3, F3.

"If energy prices will trend higher, you invest one way; if energy prices will be lower, you invest a different way. But if you don't know what prices will do, often you do not invest at all."

Lawrence H. Summers, Director of the National Economic Council

In a speech at the Brookings Institution, March 13, 2009.

1. Introduction

As the above quote demonstrates, there has recently been considerable interest in the effects of uncertainty in energy markets, and a common view is that uncertainty about energy prices tends to delay investment. This notion is rooted in theoretical models of investment under uncertainty developed by, for example, Brennan and Schwartz (1985), Majd and Pindyck (1987), and Brennan (1990). These concepts are now widely taught in MBA programs, and widely applied in capital budgeting decisions, particularly in the energy sector.[1]

This mechanism is illustrated for financial options by Black and Scholes (1973), in that an increase in uncertainty about the return to the underlying asset tends to increase the time value of a call option – i.e., the value of waiting, which delays exercising. The analogy in capital budgeting is that an increase in uncertainty about the return to an irreversible investment may tend to increase the value of waiting, rather than committing the investment. With respect to energy prices, this mechanism was developed explicitly by Bernanke (1983). Bernanke (1983) shows that if, for example, oil prices are volatile, then firm's will tend to delay investment and production, until some of the uncertainty about the future path of oil prices is resolved. This will lead to lower output and production when uncertainty about oil prices is high, with the effect being strongest in sectors that are either energy intensive, or produce energy consuming goods.

More recently, as nearby oil futures since 2007 have skyrocketed from $60 to $130 per barrel, and then plunged to $40, the option value of waiting may induce auto manufactures, for

[1] See, for example, the Real Options Group at http://rogroup.com.

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example, to delay investment, as they decide whether to commit resources to the design and production of vehicles that are more or less fuel efficient. Similarly, we might expect firms to delay investment in energy exploration, energy conservation and alternative energy. The net effect would be lower production.[2]

Despite the wide acceptance of real options in capital budgeting decisions, there exists relatively little empirical evidence on aggregate effects of real options at the industry and multi-industry levels. One notable exception is Moel and Tufano (2002), who examine the effect of volatility in gold prices on mining. As they note, "because volatility does not affect the investment decision in simpler [discounted cash flow] models, its economic significance… informs us about the relevance of real option-like models."

In this paper, we examine the effects of the option value to delay investment and production induced by volatility, or uncertainty, about oil prices in G-7 countries. We do so by utilizing a simultaneous equations model that accommodates a role for oil price uncertainty on production. The empirical model is based on a structural VAR that is modified to accommodate multivariate GARCH-in-Mean errors, as detailed in Elder (2004). 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 allows us to isolate the effects of uncertainty about oil prices on measures of production in energy related industries. Consistent with Brennan and Schwartz (1985) and Bernanke (1983), if changes in oil prices are accompanied by an increase in uncertainty, then the option value of waiting may cause both increases and decreases in oil prices to dampen production in the short-run. In this sense, the effects of oil prices may be *asymmetric*.

The increased volatility of oil prices has drawn considerable interest in the finance literature, as evidenced by, for example, Cunado, Gil-Alana and De Gracia (forthcoming), Kogan, Livdan and

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[2] The very existence of futures markets suggests the ability to hedge some price risk over short horizons, but, in practice, firms typically do not, and cannot, completely determine their risk exposure (cf. Adam and Fernando, 2006).

Yaron (2009), Wang, Wu and Yang (2008) and Chang, Daouk, and Wang (2009), Switzer and El-Khoury (2007) and Bergin and Glick (2007).

Our investigation is therefore interesting and relevant for a number of reasons. First, our results provide evidence on whether fluctuations in observed industry output can be explained by real options models. Second, our results provide evidence on whether a mechanism such as the option value of waiting may be one reason to cause the response of production to oil shocks to be asymmetric, which is an issue of considerable recent interest (cf. Hamilton (2003)). Third, applying this empirical model to the G-7 provides a test of robustness of Elder and Serletis (2009), who find that oil price uncertainty adversely affects investment and production in the US. Fourth, the cross section of G-7 countries offers a diverse pattern of oil consumption, oil exports and economic conditions in which to analyze aggregate effects of the option value of waiting. 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 – considerably larger than for the remaining G-7 countries. Two countries in our sample were net oil exporters over at least part 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, or international, factors.

Our results suggest that the aggregate effect of oil price uncertainty on the option value of waiting, as captured by output in manufacturing related industries, is both measurable and significant in four of the countries in our sample (Canada, France, UK and the US). The results for three of these 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 real options mechanisms described by, for example, Bernanke (1983) and Brennan and Schwartz (1985).

Graphical plots reveal that oil price uncertainty spiked when OPEC collapsed during the mid 1980's, suggesting that uncertainty about oil prices may have contributed to the stagnant economic growth during this period. More recently, between 2005 and 2007 oil price uncertainty has been considerably more elevated it has not equaled 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 (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 permit the conditional volatility of oil to affect the conditional mean;

$$\mathbf{B}\mathbf{y}_t = \mathbf{C} + \mathbf{\Gamma}_1\mathbf{y}_{t-1} + \mathbf{\Gamma}_2\mathbf{y}_{t-2} + \cdots + \mathbf{\Gamma}_p\mathbf{y}_{t-p} + \mathbf{\Lambda}H_{oil}(t)^{1/2} + \mathbf{\varepsilon}_t, \tag{1}$$

where $\dim(\mathbf{B}) = \dim(\mathbf{\Gamma}_i) = (N{\times}N)$, $\mathbf{\varepsilon}_t \mid \psi_{t-1} \sim$ iid $N(\mathbf{0}, \mathbf{H}_t)$, $\mathbf{H}_t$ is diagonal, $H_{oil}(t)^{1/2}$ 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. $\mathbf{\Lambda}$ 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 $H_{oil}(t)^{1/2}$, 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 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 (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 $\mathbf{B}\varepsilon_t\varepsilon_t'\mathbf{B}'$ -- 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), Mork (1989), Lee *et al*. (1995), Bernanke *et al*. (1997), Hamilton and Herrera (2004), Kilian (2008) and Elder and Serletis (2009). 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) uses 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 in local currency. Following Blanchard and Gali (2007), we use the nominal price of oil in local currency rather than the real price of oil, for the reasons they elaborate. 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, 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. Terminating our sample prior to 2008 excludes a recent period of extreme volatility in oil prices that coincided with a collapse in the international financial system.   By excluding this period, we avoid inadvertently attributing adverse effects associated with the financial system to uncertainty about oil prices. Figures 1 and 2 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 the appropriate transformations, as discussed below. With regard to identification, we allow **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). In the next section, we discuss the transformations of these variables related to stationarity.

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, if necessary, cointegration. Table 1 reports the results of augmented Dickey-Fuller (ADF) tests for unit roots, 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.[3]

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). In the baseline model, the vector y_t is therefore

$$\mathbf{y_t} = \begin{pmatrix} \Delta\ln(CPI_t) \\ \Delta\ln(IP_t) \\ \Delta\ln(oil_t) \\ i_t \end{pmatrix}.$$

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, so that the vector

of data becomes

$$\mathbf{y_t} = \begin{pmatrix} \Delta\ln(CPI_t) \\ \Delta\ln(IP_t) \\ \Delta\ln(oil_t) \\ \Delta i_t \end{pmatrix}.$$

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[3] Note that we could, at this stage, examine cointegration in two-variable pairs. To discount the probability of imposing a spurious cointegrating relationship, we would consider only cointegrating relationships that are likely to be justified by economic theory. Such relationships include a Fischer effect, which posits cointegration between interest rates and inflation, with a known cointegrating vector of $(1, -1)$. However, since the interest rate is stationary by assumption, there is no cointegrating relationship to impose.

Finally, we also estimate the model with the *CPI* in first differences for Canada, Italy, Japan and the UK, so that the data vector for these countries becomes

$$\mathbf{y_t} = \begin{pmatrix} \Delta(CPI_t) \\ \Delta\ln(IP_t) \\ \Delta\ln(oil_t) \\ i_t \end{pmatrix}.$$

Such transformations have very little effect on the estimated relationship between the oil price uncertainty and 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.[4]

The primary coefficient of interest, the coefficient on oil price uncertainty in the 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 (2009) who find a negative effect of oil price uncertainty on several measures of US economic activity. Our results indicate that the net oil

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

exporter nature of both Canada and the UK does little to limit the exposure to the contractionary

effects of oil price uncertainty. Our results are consistent with the theory of the option value of

waiting, as increased oil price uncertainty may delay investment in both net oil producing and net oil

consuming industries and countries. Our results also indicate the lack of significance of oil price

uncertainty for the case of Germany, Italy and Japan. This is particularly surprising for the case of

Japan and Germany, the second and fourth largest oil importers in the world.[5] However, in all three

countries there has been particular emphasis on improving energy efficiency. This is particularly the

case in Japan where its oil dependency has fallen dramatically in recent years due to a drive for

energy diversity.[6] This is also the case in Germany, but particularly following unification. Policy has

also played a role in Italy, where a number of measures have been introduced including freezing

energy tariffs. Finally, the lack of a significant effect for the case of Italy and Germany may also be

due to the offsetting effects played by real effective exchange rate depreciation, in particular since

economic and monetary union (EMU).

In figures 4 and 5 we plot the estimated conditional standard deviation of oil against the

growth rate of 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 prices from 2003 through the end of 2007 were accompanied

by only relatively 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 prior to the end of 2007.

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[5] See the Energy Information Administration (www.eia.doe.gov).
[6] A further complication in the case of Japan may the prolonged economic downturn. Further, Mork (1994) has reported that Japan was relatively less sensitive to oil price shocks of the early 1970's.

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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 10% level.[7] 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.[8]

Additional Tests of Robustness

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

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

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 for each country, in which the variables have the same transformation as in our baseline model

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

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 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

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

where $\varepsilon_t * \sim N(0, H_{oil}(t)*)$ 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, with the coefficient on $\hat{H}_{oil}(t)*$ 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.[9]

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. This effect is consistent with uncertainty about oil prices increasing the option value

of waiting, which would tend to delay and investment and production, as detailed by Bernanke

(1983).

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 negative oil shocks affect should be expansionary,

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

even in the short-run. Whether the net short-run effect of negative oil shocks is contractionary or expansionary depends on whether the effect due to increased uncertainty about oil prices outweighs the effects from other channels.

We can calculate this net short-run effect in our empirical model by simulating the response of production 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 may be one factor that helps resolve the observation 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. Real options theory 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, thereby inducing firms to delay investment and production. 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 are significant, and that,

in the short-run, both positive and negative oil shocks may be contractionary. In particular, our

results suggest that the real options effect may be sufficiently large that even falling oil prices,

accompanied with increased uncertainty, may induce firms to delay investment and production.

5. Conclusion

Theoretical models of investment under uncertainty predict that uncertainty about the return

to an investment tends to induce firms to wait, rather than exercise, an investment option. This

implies that volatility, or uncertainty, in commodity futures may tend to decrease production,

particularly in energy intensive or energy extensive industries. In addition, there has been recently

been considerable interest in the effects of excessive volatility in energy futures markets.

In this paper we examine this prediction, utilizing a commodity – oil – that is central to the

global economy (cf. Hamilton 2003). To do so, we model uncertainty about oil prices as the variance

of the one-step-ahead forecast error, and we utilize 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 consistent the predictions from the real options literature -- 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 not stimulate production.

Our result also indicates that the apparent asymmetry in the response of output to oil prices is a feature common to several industrialized economies. Finally, our result suggests that the dramatic increase in the variability of oil prices observed since early 2008 has likely been a contributing factor in the economic contraction among industrialized countries.

Series	Transformation in baseline and alternative models	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 model	Short term interest rate. Treasury bill 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*
ADF Test	Φ_3	Φ_3	Φ_3	τ_μ	τ_μ
5% Critical Value	6.30	6.30	6.30	–2.87	–2.87
Canada	93.84**	5.96	4.12	–2.34	–2.17
France	5.91	2.61	3.91	–2.53	–1.65
Germany	3.57	4.78	2.97	–1.26	–3.16**
Italy	31.91**	2.79	2.98	–2.46	–1.63
Japan	8.65**	5.09	1.86	–1.65	–2.40
United Kingdom	8.73**	4.36	2.02	–2.73	–2.57
United States	2.89	2.95	4.56	–2.24	–2.24

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 $H_0: \rho = 1$.

TABLE 2. MODEL SPECIFICATION TESTS

	Schwarz criterion values		
	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
United Kingdom	1221	877	636
United States	-782	-1239	-1569

Notes: 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 $\varepsilon_t \sim N(\mathbf{0}, \mathbf{H}_t)$ and $G_1 = 0$, and the Multivariate GARCH-M VAR given by equations (1) and (2) with the diagonal elements of F_1 and G_1 unrestricted.

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

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

** denotes significance at the 5% level.

 * denotes significance at the 10% level.

Equation	Conditional Variance	Constant	$\varepsilon_i(t-1)^2$	$H_{i,i}(t-1)$
France				
Infl	$H_{1,1}(t)$	0.001**	0.000	0.000
		(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)
United Kingdom				
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)

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

** denotes significance at the 5% level.

 * denotes significance at the 10% level.

Equation	Conditional Variance	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)

Notes: These are the parameter estimates for the free elements in **F** and **G** from the model given by equations (1) and (2) with $\varepsilon_t \sim N(\mathbf{0}, \mathbf{H}_t)$. Each row in the table represents an equation from the associated multivariate GARCH-in-Mean VAR. Asymptotic *t*-statistics are in parentheses. A coefficient of 0.000 indicates that the nonnegativity constraint is binding.
** denotes significance at the 5% level.
 * denotes significance at the 10% level.

TABLE 4. COEFFICIENT ESTIMATES ON OIL VOLATILITY IN THE IP EQUATION

Model	Sample	Coefficient on $H_{oil}(t)^{\frac{1}{2}}$, Oil Volatility
Canada	1974:01-2007:07	−0.065** (2.99)
France	1974:01-2007:10	−0.048** (2.14)
Germany	1974:01-2007:07	−0.015 (0.61)
Italy	1974:01-2007:10	−0.026 (1.02)
Japan	1974:01-2007:10	−0.013 (0.59)
United Kingdom	1974:01-2007:10	−0.077** (4.87)
United States	1974:01-2007:10	−0.038** (3.37)

Notes: These are the parameter estimates for Λ from the structural VAR with multivariate GARCH model given by equations (1) and (2), with the following variables and transformations: $\Delta\ln(CPI)$, $\Delta\ln(IP)$, $\Delta\ln(Oil)$, *Rate*. $H_{oil}(t)^{\frac{1}{2}}$ denotes the conditional standard deviation of the oil price, measured in the domestic currency. Absolute asymptotic *t*-statistics are in parentheses. Variable transformations are detailed in Appendix A1.
** denotes significance at the 5% level.
* denotes significance at the 10% level.

TABLE 5. COEFFICIENT ESTIMATES ON OIL VOLATILITY IN OLS REGRESSION

Model	Sample (including pre-sample)	Coefficient on $H_{oil}(t)^{\frac{1}{2}}$ Oil Volatility
Canada	1974:01-2007:07	−1.619** (3.20)
France	1974:01-2007:10	−0.058** (2.26)
Germany	1974:01-2007:07	−0.031 (0.93)
Italy	1974:01-2007:10	−0.051 (1.36)
Japan	1974:01-2007:10	−0.013 (0.50)
United Kingdom	1974:01-2007:10	−1.50** (3.20)
United States	1974:01-2007:10	−0.038** (2.92)

Notes: These are the parameter estimates on oil price volatility in a regression of the growth rate of industrial production on lagged $\Delta \ln(CPI)$, $\Delta \ln(IP)$, $\Delta \ln(Oil)$, *Rate* and oil price volatility as given by equation (5). $H_{oil}(t)^{\frac{1}{2}}$ 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.
** denotes significance at the 5% level.
* denotes significance at the 10% level.

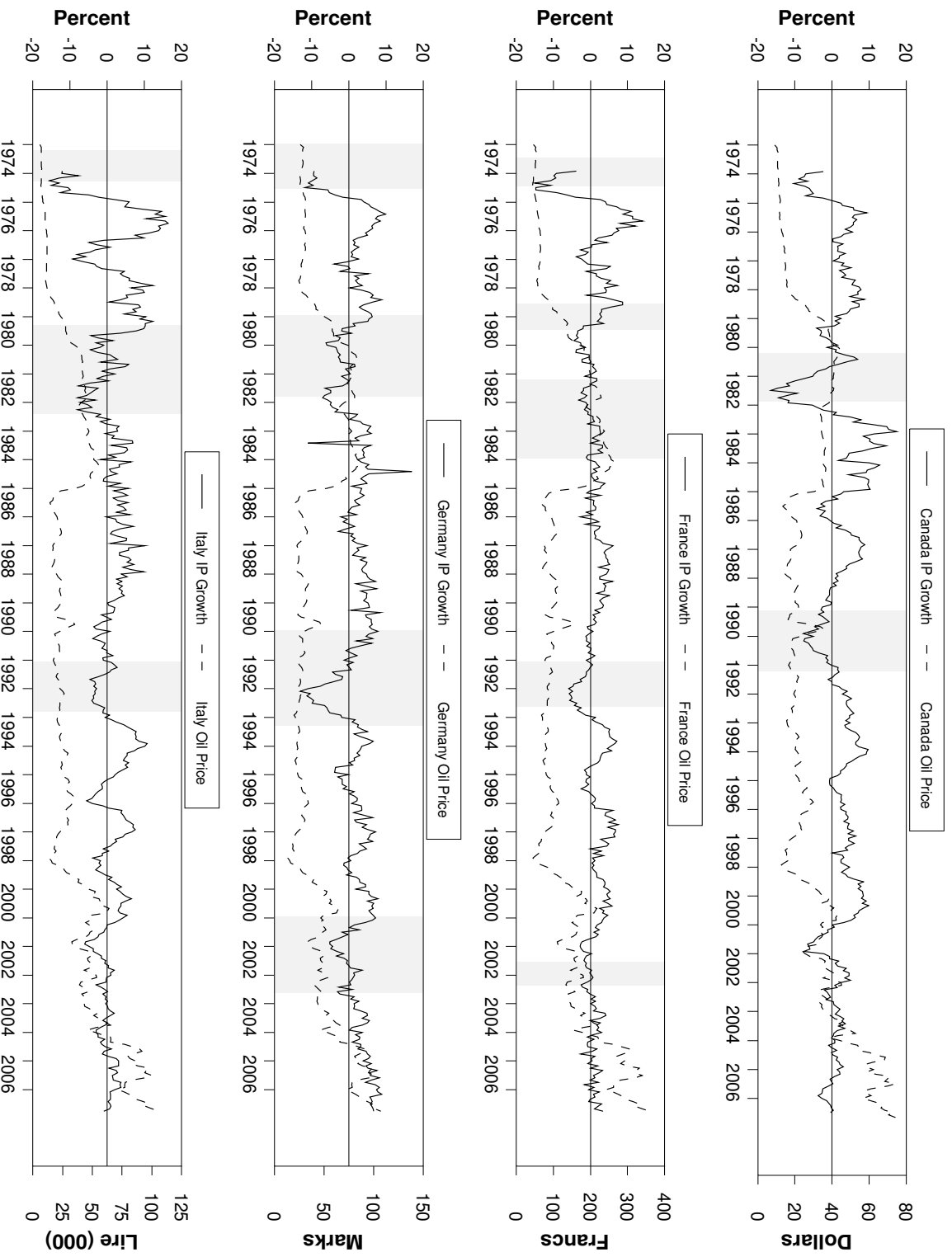**Fig. 1. Industrial Production and Oil Prices**
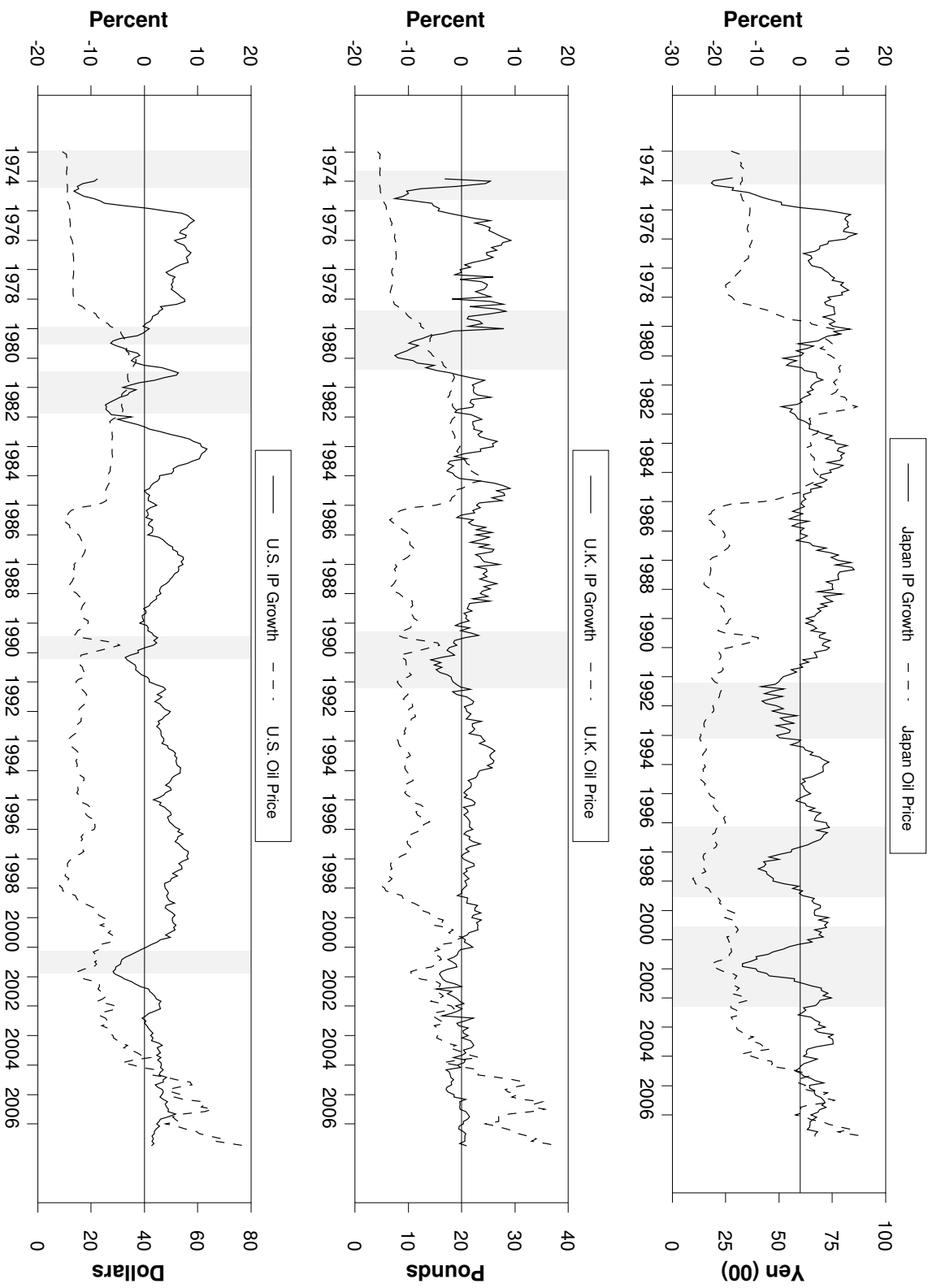
Fig. 2. Industrial Production and Oil Prices

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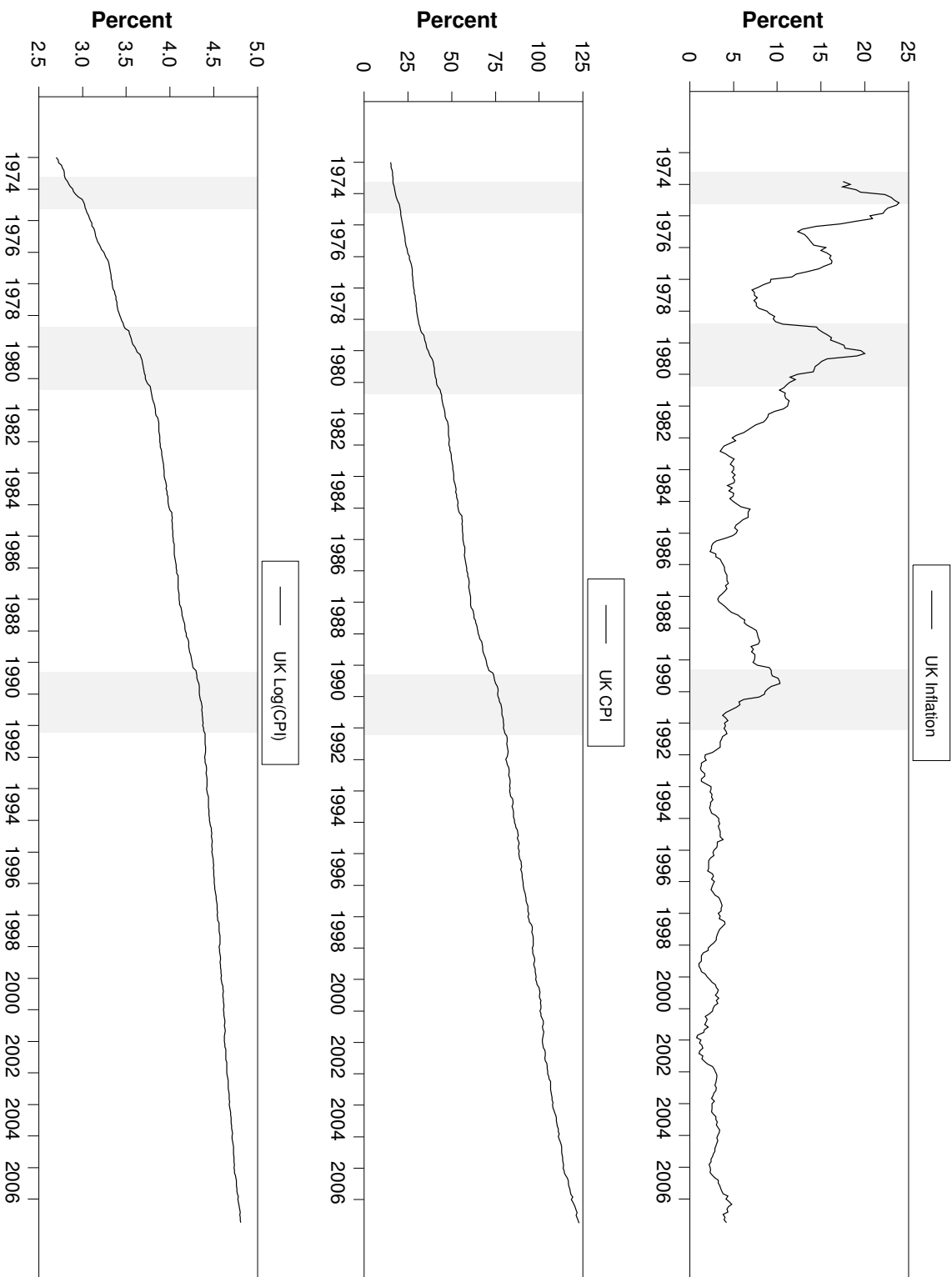
Fig. 3. UK CPI Transformations

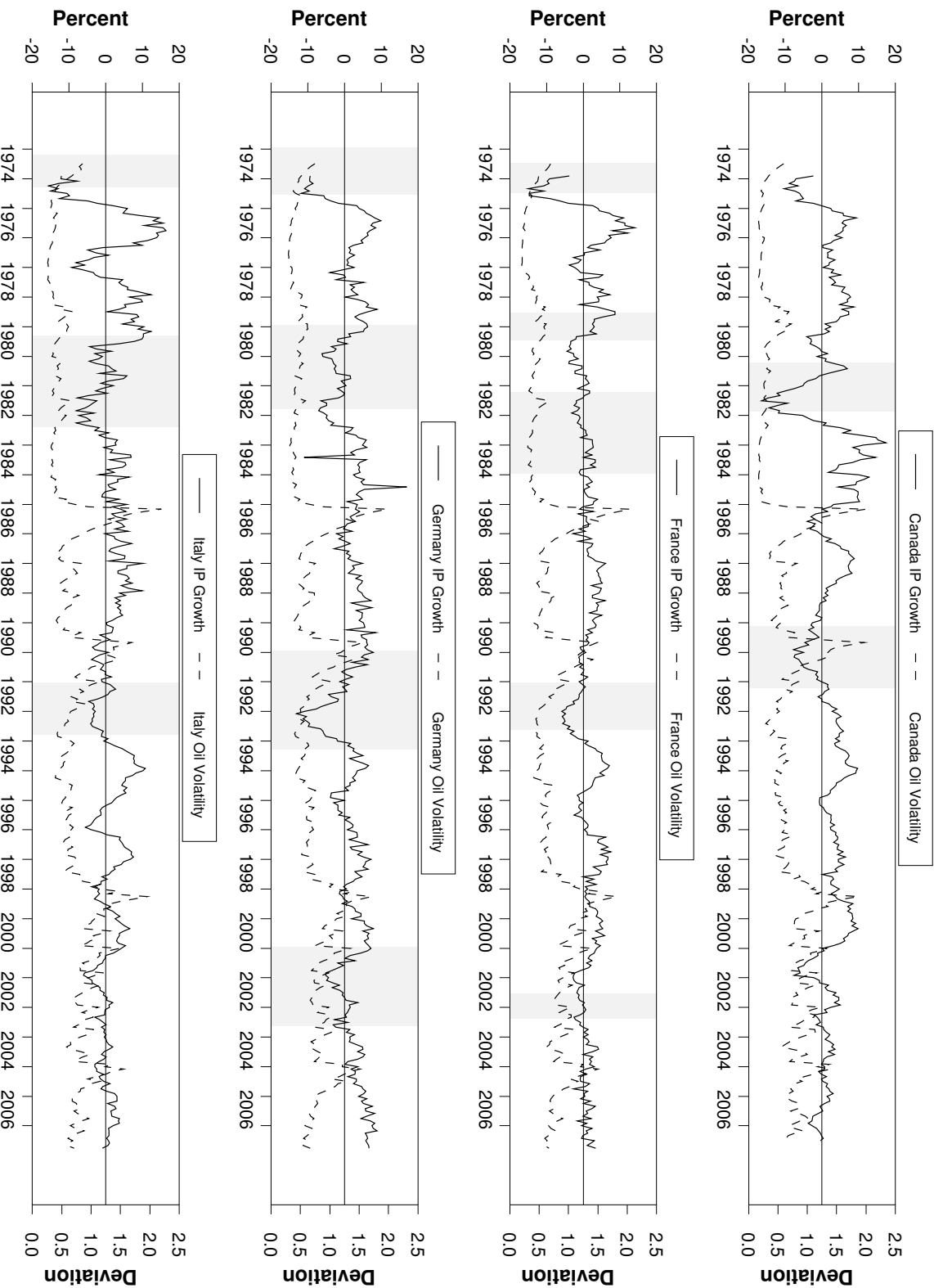# Fig. 4. Industrial Production and Oil Volatility

33

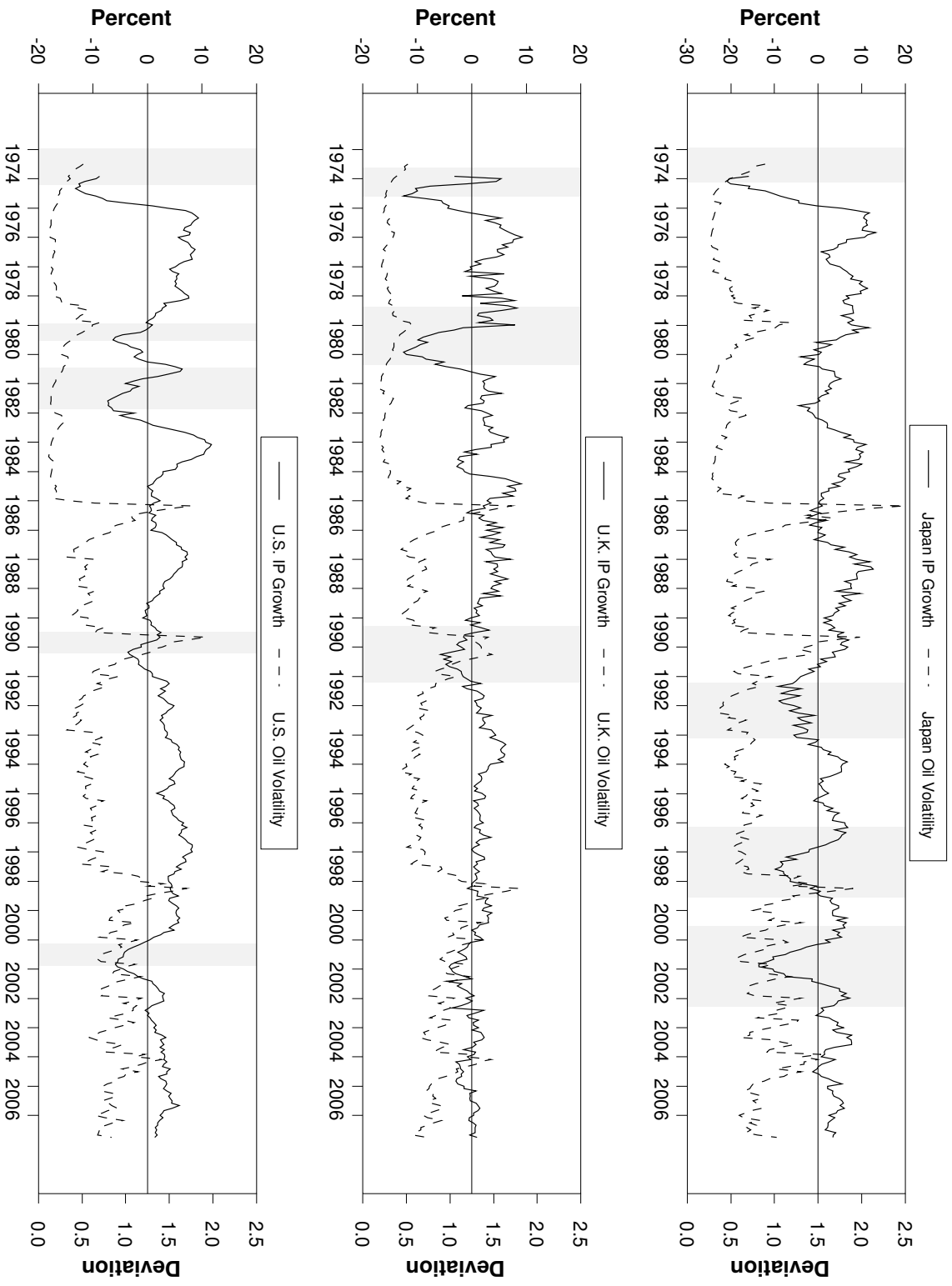**Fig. 5. Industrial Production and Oil Volatility**
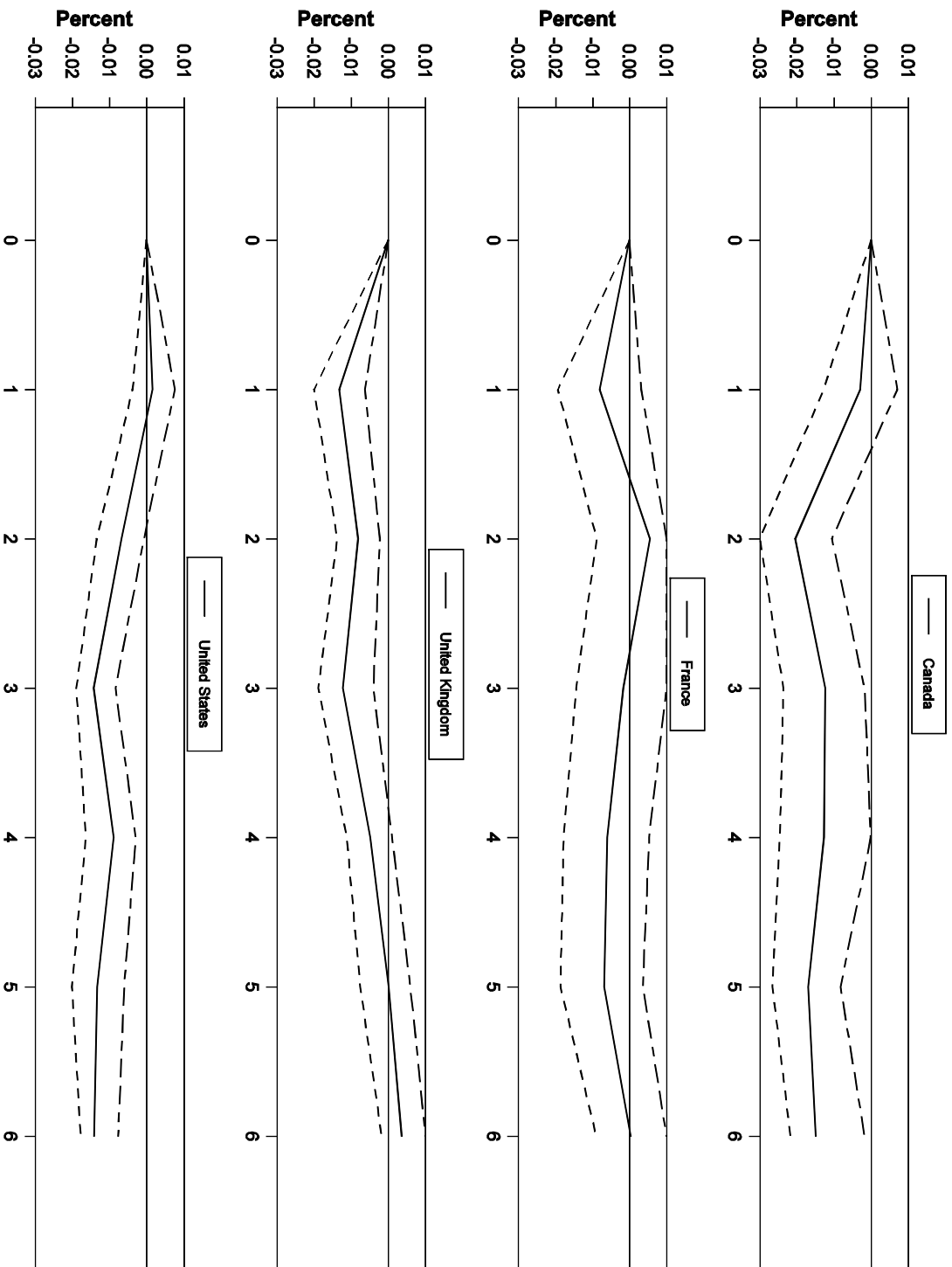
Figure 6. Response of Production to Positive Oil Shock

Figure 7. Response of Production to Negative Oil Shock