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by Alessio Anzuini, Patrizio Pagano and Massimiliano Pisani

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INFORMATION FROM FINANCIAL MARKETS

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Abstract

This paper analyzes the macroeconomic effects on the U.S. economy of news about oil supply by estimating a VAR. Information contained in daily quotations of oil futures contracts is exploited to estimate the dynamic path of oil prices following a shock. Hence, differently from the VAR literature on oil shocks we do not need to rely on recursive identification. Impulse response functions suggest that oil supply disruptions have stagflationary effects on the U.S. economy. Historical decomposition shows that oil shocks contributed significantly to the US recessions of the last thirty years, but not all exogenous increases in oil prices have induced a recession. Finally, the contribution of oil shocks to inflation fluctuations seems to have declined over time.

JEL Classification: C2, E3, O41.
Keywords: vector autoregression, oil shock, futures, news.

Contents

* Bank of Italy, Economic Research Department.

1 Introduction[1]

The study of the effects of oil price changes on macroeconomic variables using Vector Autoregressive (VAR) systems is now well established in the economics literature, dating back at least to Hamilton (1983). The common approach to identification is the recursive scheme borrowed from the literature on monetary policy analysis.[2] The typical assumption – used for instance by Burbidge and Harrison (1984) and by Bernanke *et al.* (1997) – is that oil shocks do not have simultaneous effect on all the VAR variables, while, only some of the variables, if any, simultaneously affect oil prices. Yet there is a wide debate in the VAR literature about the validity of the recursive identifying assumptions. Some authors suggest that they will not hold exactly in the data and that conclusions are not robust to minor deviations from the assumptions.[3]

In this paper we focus on the effects that news about oil supply has on US macroeconomic conditions following the identification strategy proposed by Faust *et al.* (2004) in their analysis on monetary policy shocks.[4] In particular, we use information from oil futures and spot prices corresponding to daily events classifiable as oil supply disruptions. This permits us to avoid imposing an arbitrary recursive identification. The news we select can be interpreted as oil supply shocks in a broad sense, since they include events of both actual and potential oil supply disruptions. For instance, geopolitical tensions may drive up oil prices even in the absence of actual supply restraints, but just because precautionary demand increases for fear of future supply cutbacks.

We estimate the response of oil prices to unexpected shocks by regressing separately futures changes at various horizons on spot changes. From these regressions, we obtain the corresponding dynamic path of the oil price. We then impose that the VAR response of the oil price to its own shock matches the response estimated using futures data. The VAR that we estimate on

[1] We thank Jean Boivin, Fabio Canova, Andrea Finicelli, Oscar Jordà, Alessandro Secchi and participants to the Workshop "Monetary and financial implications of globalization" held at Banco de Mexico for comments. We also thank Stephen G. Donald for sharing his routines. Giovanna Poggi provided valuable research assistance. We are solely responsible for any errors. The opinions expressed in this paper do not necessarily reflect those of the Bank of Italy. Address: via Nazionale 91, 00184 Rome - Italy. E-mail: alessio.anzuini@bancaditalia.it; patrizio.pagano@bancaditalia.it; massimiliano.pisani@bancaditalia.it

[2] See for example Christiano *et al.* (2005).

[3] See, for example, Leeper *et al.* (1996), Uhlig (1997) and Faust (1998).

[4] Faust *et al.* (2004) use changes in Fed funds futures around policy decisions to identify monetary policy shocks. High-frequency data are also used to identify such shocks in Bagliano and Favero (1999) and Cochrane and Piazzesi (2002).

US monthly data from January 1965 to December 2005 is rather standard. The variables are those used in similar studies: oil spot prices (WTI quality), the effective Fed fund nominal interest rate, the consumer price index (CPI), industrial production, nominal wages and the money aggregate M2.[5]

The main results are as follows. First, impulse response analysis shows that after the oil price shock industrial production decreases, reaching a through after roughly one year, while the CPI level persistently increases within a few quarters. Hence, exogenous oil supply disruptions cause stagflation (defined as negative co-movement between the level of industrial production and the level of consumer prices). Second, historical decomposition shows that oil shocks have contributed to each of the US recessions of the last thirty years. Third, over time, the effect of oil shocks on CPI inflation has declined. For robustness we perform the analysis using a recursive identification of the shock (as in Burbidge and Harrison, 1984) and results do not change significantly.

A building block of our identification scheme relies on the use of selected exogenous events. Other works also follow this strategy. For instance, Hamilton (2003) uses oil supply disruptions to detect oil price shocks, building on Hamilton (1985), who singles out exogenous oil supply shocks by using dummy variables associated with some events — probably exogenous to developments in the US economy — characterized by dramatic increases in the nominal price of oil. Kilian (2006a) derives a measure of oil supply shortfall for several oil-producing economies by comparing the level of observable oil production with the counterfactual level extrapolated by the supply of similar countries not affected by the exogenous event. Differently from them, and similarly to Cavallo and Wu (2006) and Kilian (2006b), we also pinpoint oil shocks due to the fear of future supply disruptions. Overall, we find that oil market event-day surprises are not trivial (the median daily change in spot prices is 1 per cent) and that during the most recent build-up of oil prices (since 2003) they have become more frequent and relatively smaller than in the past. For robustness, we further estimate the dynamics of oil prices by considering separately: (1) events unrelated to (possibly endogenous) OPEC decisions, (2) events in the period 2003-2005 and (3) big events (those with oil price changes larger than 5 per cent). Results are unchanged.

In addition to the literature on the macroeconomic effects of oil shocks, this paper also contributes to the reviving body of work that proposes news about future agents' expectations or changes in them as important sources of business fluctuations (e.g. Beaudry and Portier, 2006a, 2006b, Christiano *et al.* 2007, Jaimovich and Rebelo, 2006). This literature shows that when agents receive news that future fundamentals will be different from what was previously expected, their change of behaviour may influence current

[5] In particular this list is almost identical to that in Burbidge and Harrison (1984).

macroeconomic aggregates, generating volatility, co-movement, and persistence that are empirically plausible. Consistently with this approach, our findings suggest that news about potential disruptions in oil supply has a crucial role in explaining output and inflation dynamics.

Overall our results are similar to those obtained by other authors. Hamilton (1983) estimates a VAR and finds evidence, later updated and confirmed by Mork (1989), that oil price increases Granger-cause real output reductions. Burbidge and Harrison (1984) estimate VARs for five advanced economies and find that oil price innovations had a large role in the stagflation of 1973-74, but a minimal (except for Japan) in that of 1979-80. Using the same approach, Hooker (1996) concentrates on the US and discusses several possible explanations for the smaller impact of oil price shocks on macroeconomic variables after 1973. Bernanke *et al.* (1997) use a modified extension (by Hoover and Perez, 1994) of Hamilton (1985) and find that the effects of oil price shocks may be affected by the endogenous response of monetary policy. Finally, our results are in line with those of Cavallo and Wu (2006) and of Kilian (2006b), who find stagflationary effects on the US economy of oil price shocks related to possible future supply disruptions.

The rest of the paper is organized as follows. In the following section we illustrate how we select news on oil supply disruptions and how we use financial data to estimate the dynamic path of oil prices. In Section 3 we illustrate the results of the structural VAR analysis. Section 4 provides sensitivity analysis. The final section contains some concluding remarks.

2 Oil price responses to oil supply news

The identification strategy that we pursue in this paper is described in detail in Appendix A. The core of this strategy is the estimation of the following equation at various horizons:

$$\Delta_{d_t} f_{t+h} = \alpha_h + \beta_h \Delta_{d_t} s_t \tag{1}$$

where $\Delta_{d_t} f_{t+h} \left(= f_{t+h}^{d_t} - f_{t+h}^{d_{t-1}} \right)$ is the change in oil futures prices with maturity $h = 1, ..., 5$ months between the day d_{t-1} before an event that we classify as oil supply news and d_t, the day of the event itself.

Equation (1) states that each day oil supply news hits the market, futures prices at different horizons change proportionally to spot prices s_t. To estimate the coefficients of this relationship we need the dates of the news and changes in spot and futures prices.

We define as oil supply news those events that lead to actual or possible future disruption of oil supply – therefore exogenous with respect to global demand conditions – such as hurricanes, wars, civil unrest and political tensions in oil-producing countries, but also OPEC decisions to change output or quotas, the discovery of new oil fields or the release of strategic reserves.

Table 1: Oil price spot and futures following oil supply news: summary statistics (percentage changes)

percentiles	spot	t+1	t+2	t+3	t+4	t+5
Jan.1986-Dec.2005 (125 events)						
10th	-3.68	-3.51	-2.87	-2.80	-2.45	-2.40
25th	-1.78	-1.35	-1.34	-1.13	-1.04	-0.97
50th	1.02	0.93	0.83	0.77	0.78	0.71
75th	2.87	2.87	2.37	2.37	2.22	2.13
90th	6.22	6.13	5.18	4.65	4.68	4.25
Apr.2003-Dec.2005 (47 events)						
25th	-1.73	-1.20	-0.94	-0.77	-0.71	-0.69
50th	0.48	0.87	0.73	0.71	0.70	0.68
75th	1.86	2.17	2.04	2.01	1.97	1.95

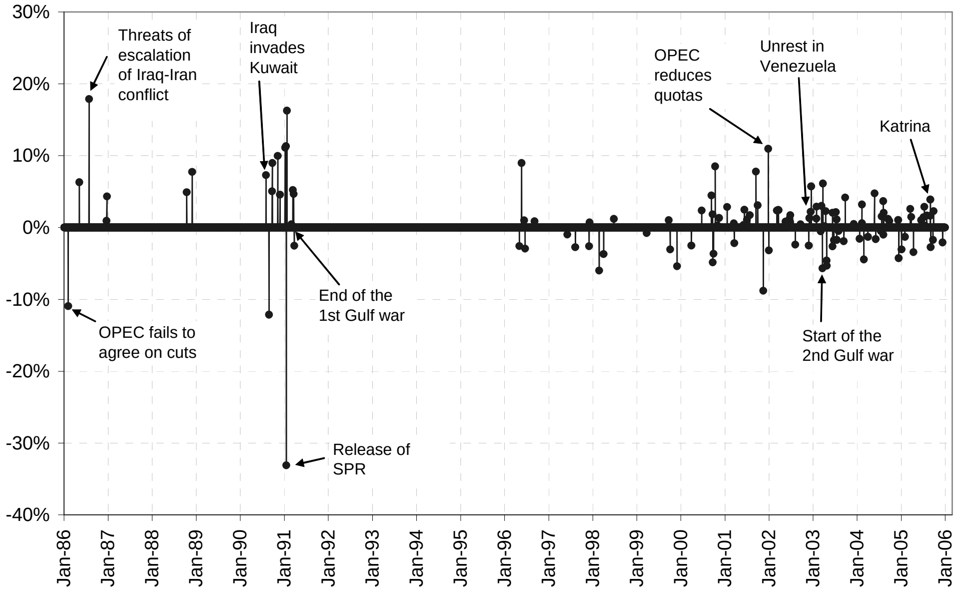
We use data on the closing prices of oil spot and futures contracts for 1-5 months, exchanged on NYMEX from January 1986 to December 2005. On the basis of the "World oil market and oil price chronology 1970-2004" and of the "Monthly energy chronology 2005" we select 125 events.[6] We then measure the unexpected change in the oil price as the difference between the closing spot price on the day of release of the news and the spot price on the day before. We do the same thing for the 1- to 5-month ahead futures prices. If the event happens in the last week of a month, when usually 1-month ahead futures are no longer quoted, we take 2- to 6-month ahead contracts.

The spot and futures changes are summarized in Table 1. The oil market event-day surprises are not trivial. The median change is 1 per cent and more than 50 per cent of price changes are below -1.8 per cent or above 2.9 per cent. The change in oil spot prices following oil supply news is shown in Figure 1. Those changes were relatively sharper – albeit less frequent – in the first part of the sample period. More recently, surprise changes have been more frequent, but relatively smaller. In fact, since April 2003, the beginning of the latest oil price build up, we record 47 events (or 38 per cent of the whole sample), with a median change (0.5 per cent) equal to half the median of the whole sample and with 50 per cent of the cases ranging between -1.7 and 1.9 per cent.

In the whole period the number of big events, defined as changes in the spot prices larger than 5 per cent in a single day, is 25. For instance, in Feb-

[6] We download them from the website of the Energy Information Administration (www.eia.doe.gov). An appendix with the detailed specification of the events is available from the authors upon request.

Figure 1: Oil price surprises



Notes: Daily percentage changes in oil spot prices (WTI) following oil supply news.

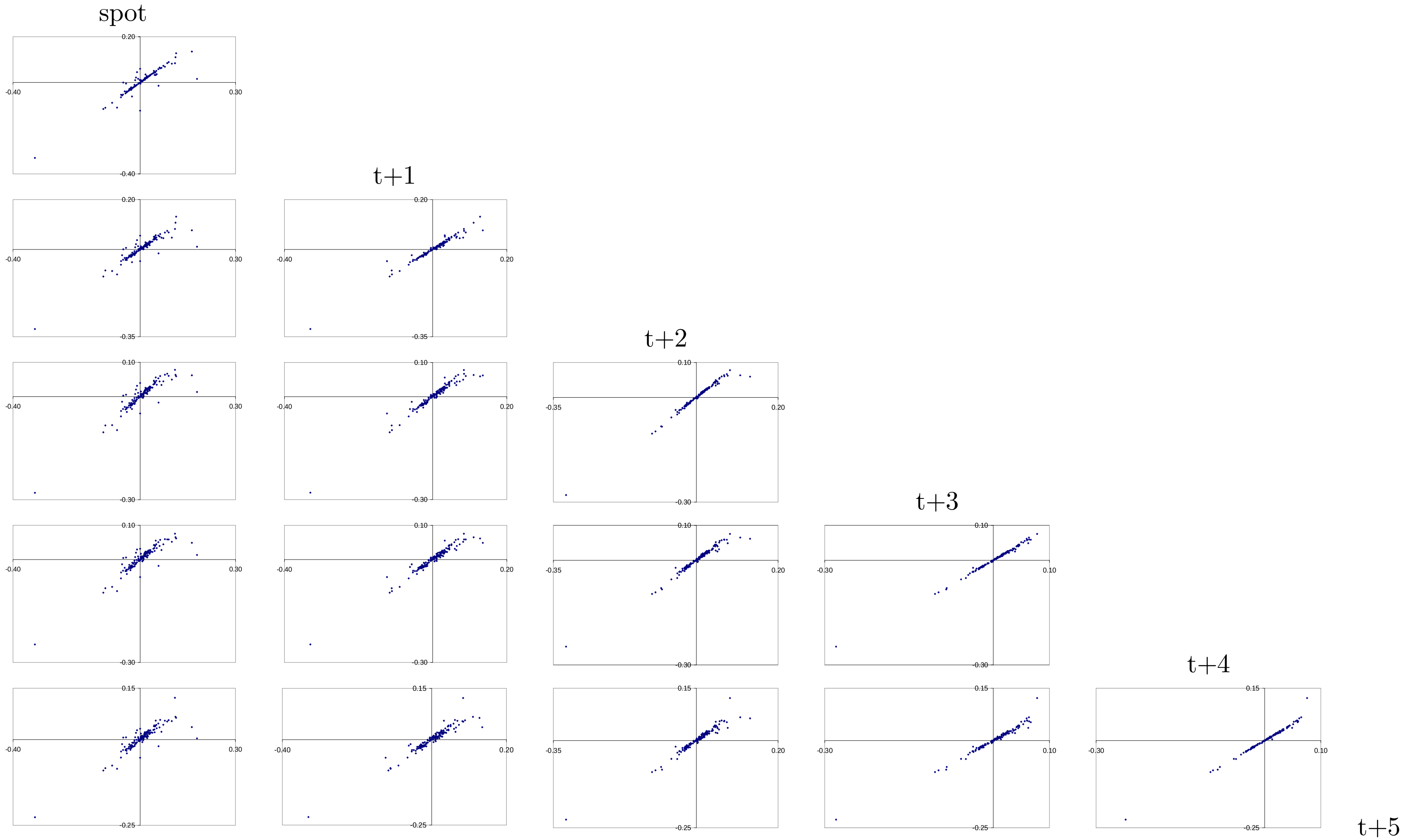
ruary 1986 oil spot prices collapsed by 11 per cent following OPEC's failure to agree production. In late 1990 the Kuwait invasion and the subsequent Gulf War caused sharp changes in oil prices with some increases of more than 10 per cent followed by an abrupt fall of more than 30 per cent when the US decided to supply strategic reserves (SPR) to the market. When hurricane Katrina hit the Gulf of Mexico at the end of August 2005, oil price spiked by "just" 4 per cent.

Overall, Figure 1 provides a reasonable picture of the time pattern of oil price surprises as defined in this work and it bears a good resemblance to the measure of exogenous oil production shortfall recently constructed by Kilian (2006a).

The scatter plot of price changes at various horizons (Figure 2) shows that the linearity assumption implicit in our identification procedure seems well satisfied: the relative size of event-day change at various horizons is the same, but for a scale factor representing the sign and the size of the shock.

We regress the oil market event-day changes in the contracts for horizons 1–5 months on the oil price change. We take the impulse responses as the coefficient estimates from these regressions. They are listed in Table 2, along

Figure 2: Oil price surprises and futures price changes



with standard errors in parentheses. In the first month, the effect diminishes to almost 90 per cent of the impact effect; it gradually dissipates to about 60 per cent of the initial value over the next five months. Looking at standard errors, it can be seen that all effects are strongly significant.

One may think that since OPEC decisions may be, at least in part, driven by market developments, oil price changes happening on the day of OPEC meetings do no represent truly exogenous supply shocks. In the second column of Table 2 we show that even if we drop the events related to OPEC decisions, the responses remain the same. Furthermore, the responses remain the same if we use only the events of the recent oil price build up (third column) or, finally, if we use only events that caused oil price changes larger than 5 per cent (in absolute terms) in a single day (last column).

Our measure of oil supply shock may – at least to some extent – be contaminated by other important information that hit the market. We found that on nine dates in our sample of oil supply shocks there were releases of US industrial production and capacity utilization data, two variables that according to Pagano and Pisani (2006) capture the evolution of world oil demand and are useful in predicting oil prices. Therefore, to assess whether our results are driven by these dates, we drop them from the sample of oil shocks and re-estimate the regression equation (1). The results are similar to those with the whole sample reported in Table 2. Furthermore, if we impose that the impulse responses of the oil price in the VAR match this new response measured from the restricted sample of futures data, the impulse responses of the other variables remain the same.[7]

Our approach to the identification relies on the assumption that the futures market provides an efficient forecast of the change in the time path of the oil price or, at least, that risk premia in oil futures do not change. The small interval we concentrate on permits us to assume confidently that risk premia do not change. We test the assumption that at horizons 1-5 months ahead oil futures provide efficient forecasts of subsequent oil prices by regressing the log of average oil price (the variable we have in the VAR) on the log of the forecast for month t implicit in oil futures at month $t-1, ..., t-5$. The test that the slope coefficient (β) is equal to 1 is supported in all five cases (see Table 3). All estimates of the intercepts are not statistically different from zero, but a joint test fails to reject the assumptions of the intercepts being equal to zero and slopes equal to one. This is not a problem as long as a non-zero intercept can be traced to a constant risk premium: only risk premia varying in response to the event shock would undermine our identification procedure. Yet, the small interval over which we measure the shock should limit the latter possibility. Therefore, we assume constant risk premia and conclude that this does not affect our identification strategy.[8]

[7]Results are available upon request.

[8]Pagano and Pisani (2006) show that risk premia on oil futures are correlated with

Table 2: Impulse responses of oil price to an oil news

h	*Jan.1986-Dec.2005*	*No OPEC*	*Apr.2003-Dec.2005*	*Big events*
1	0.89 (0.03)	0.87 (0.03)	0.82 (0.07)	0.88 (0.06)
2	0.76 (0.03)	0.77 (0.04)	0.73 (0.07)	0.74 (0.06)
3	0.67 (0.03)	0.65 (0.03)	0.67 (0.07)	0.64 (0.05)
4	0.62 (0.02)	0.59 (0.03)	0.63 (0.06)	0.59 (0.05)
5	0.59 (0.03)	0.57 (0.03)	0.59 (0.06)	0.57 (0.05)
obs	125	80	47	25

Notes: OLS estimates, standard errors in parentheses. The regression is the percentage change in the futures price contracts at date $t+h$ on the surprise percentage change in the spot price.

Table 3: Forecast efficiency tests for oil price futures

h	constant (α)	slope (β)	p-value ($\beta = 1$)
1	0.08 (0.05)	0.98 (0.017)	0.15
2	0.12 (0.08)	0.97 (0.025)	0.18
3	0.14 (0.09)	0.96 (0.031)	0.24
4	0.14 (0.11)	0.97 (0.036)	0.36
5	0.14 (0.13)	0.97 (0.041)	0.48

Notes: OLS estimates, standard errors in parentheses. The regression is the log spot price at date $t+h$ on the log futures price contract at date t expiring h months later.

3 VAR analysis

In this section we report the results of the structural VAR estimation. Our monthly data set consists of six variables from January 1965 to December 2005. They are the oil spot price (WTI quality), the Fed funds rate, the monetary aggregate M2, nominal wages, the CPI and industrial production. All the variables, except the interest rate, are in logs and the VAR includes a constant and dummy variables for seasonality. We implicitly assume that there is enough co-integration so that they are jointly covariance stationary.[9] We choose 14 lags according to the Akaike Information Criterion. In monthly data, 12 lags are usually enough to eliminate autocorrelation of residuals. In our case a specification search performed with Akaike information criterion suggests that the correct specification is between 12 and 14 lags (depending on the lags included in the test). We use 14 lags as a benchmark because, even if 12 lags may be enough, two more lags may capture the remaining seasonality even after the inclusion of seasonal dummies. Moreover, many macroeconomic time series are well approximated by second-order AR models (Kim, 1999). Anyway, our results are robust regardless of 12-lag, 13-lag or 14-lag specification.

The industrial production index is our measure of economic activity. The remaining variables capture some of the most important transmission channels through which oil prices may affect economic activity indirectly. Effects of oil prices on CPI induce changes in real economic activity through changes in relative prices. There can be a monetary channel, given that short-term interest rates and M2 can react to inflationary pressures. Finally, a labour market channel is introduced by using a nominal wage index.[10]

In what follows we initially perform an impulse response analysis to understand how macroeconomic variables react to an oil price shock. Subsequently, we also analyse the effects of oil price shocks on the historical path of industrial production growth and CPI inflation.

3.1 Impulse responses

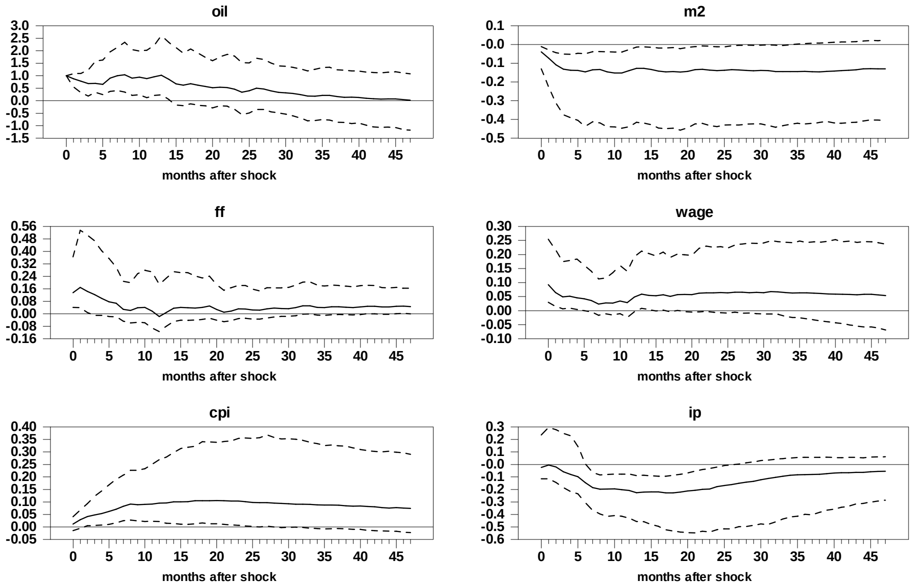
In Figure 3 we report the responses to 1 per cent oil price shock. The bands correspond to the 68 per cent confidence level and are computed using Montecarlo integration. After the oil shock, the price level increases and industrial production decreases. These movements are what we refer to as stagflation. Oil price response is rather persistent: in the first 12

business-cycle indicators, such as the degree of capacity utilization in manufacturing, at horizons longer than four months. More importantly, they show also that at short horizons (1-6 months) the assumption of constant risk premia produces forecasts of oil prices statistically not distinguishable from those obtained under the hypothesis of time-varying risk premia.

[9] See Sims (1990).

[10] This variable may also capture second-round effects of the oil price increase.

Figure 3: Estimated impulse responses and Montecarlo intervals (futures-based identification)



months after the shock, oil prices remain significantly above the baseline, to which they return only 14 months after the shock. Industrial production decreases, albeit with some lag: starting from the seventh month it becomes negative and, roughly one year after the shock it reaches a minimum value. Subsequently, it remains below the baseline, for two years after the shock. The effects of the shock die out almost completely after three years. The CPI is persistently above the baseline, reaching a statistically significant peak roughly one year after the shock. Subsequently, it remains above the baseline in a very persistent way.

Nominal wages increase on impact, then slowly decreases towards the baseline. The increase is rather small and not significant. The Fed funds rate significantly increases on impact, it reaches a peak and then it slowly decreases, until it reaches the baseline one year after the shock. The Fed funds rate response is consistent with a systematic reaction, in a restrictive direction, of the monetary authority: M2 persistently decreases, displaying a classical liquidity effect, in correspondence to the Fed funds rate increase.

3.2 Historical decomposition

To gauge the contribution of oil price shocks to the historical path of industrial production growth and CPI inflation we focus on the last five US recessions, as dated by the NBER.

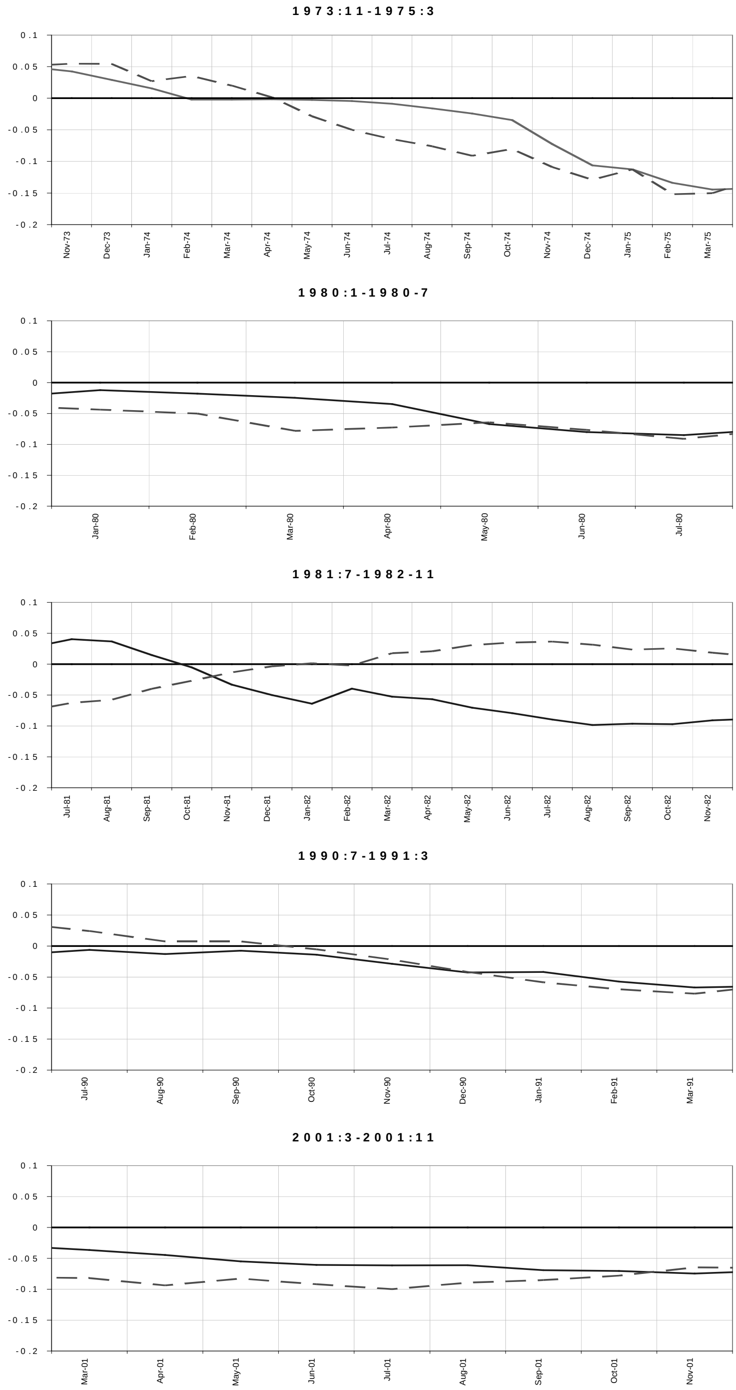
In Figure 4 we consider industrial production year-on-year growth. In each graph the blue (solid) line is the difference between actual data on industrial production growth and the data obtained from the estimated VAR under the assumption that no shocks hit the economy (the baseline projection), while the red (dashed) line reports the contribution of oil price shocks. Cumulative innovations in oil prices explain almost all the gap between the baseline and the actual data. The only exception is the second part of the double-dip recession of the 1980s – after Volcker engineered the famous and painful disinflation – which does not seem to be associated with any oil shock.

Oil supply shocks contribute not only to real activity, but also to CPI inflation, as shown in Figure 5. In particular, oil shocks help to close the gap between the baseline and actual data during the two episodes of stagflation of 1973-75 and 1980, when CPI strongly deviates from the baseline. Note also that in 1981-82, as for industrial production, CPI inflation is not strongly associated with oil price movements and that, in the last two recessions, actual inflation is much closer to the baseline, notwithstanding the positive effects of cumulative oil shocks. This result has two complementary explanations. First, other shocks having negative effects on CPI may have counterbalanced the positive effect of oil shocks: for example, positive supply-side shocks due to increases in productivity or in the degree of competition among firms. Second, in the 1970s oil shocks led to an increase in inflation largely because of the monetary policy response: in 1974 and in 1979, thinking perhaps that the shocks were temporary, the policy response was initially to ease monetary conditions. Inflation expectations rose in periods when inflation was already on the rise, and this required, later, a stronger monetary tightening. The increased credibility and transparency of monetary policy after the early 1980s may thus have contributed to stabilize inflation expectations, reducing the impact of oil shocks on CPI dynamics. Indeed, as is evident from the baseline, average inflation declined substantially.

The lower incidence of oil shocks on output and inflation is particularly evident in the last few years. Yet, to investigate more systematically the effects of the latest oil price build-up, we re-estimate the historical decomposition starting from March 2003 – the eve of the latest oil price increase. The results of these analyses for industrial production and CPI are displayed, respectively, in Figures 6 and 7.

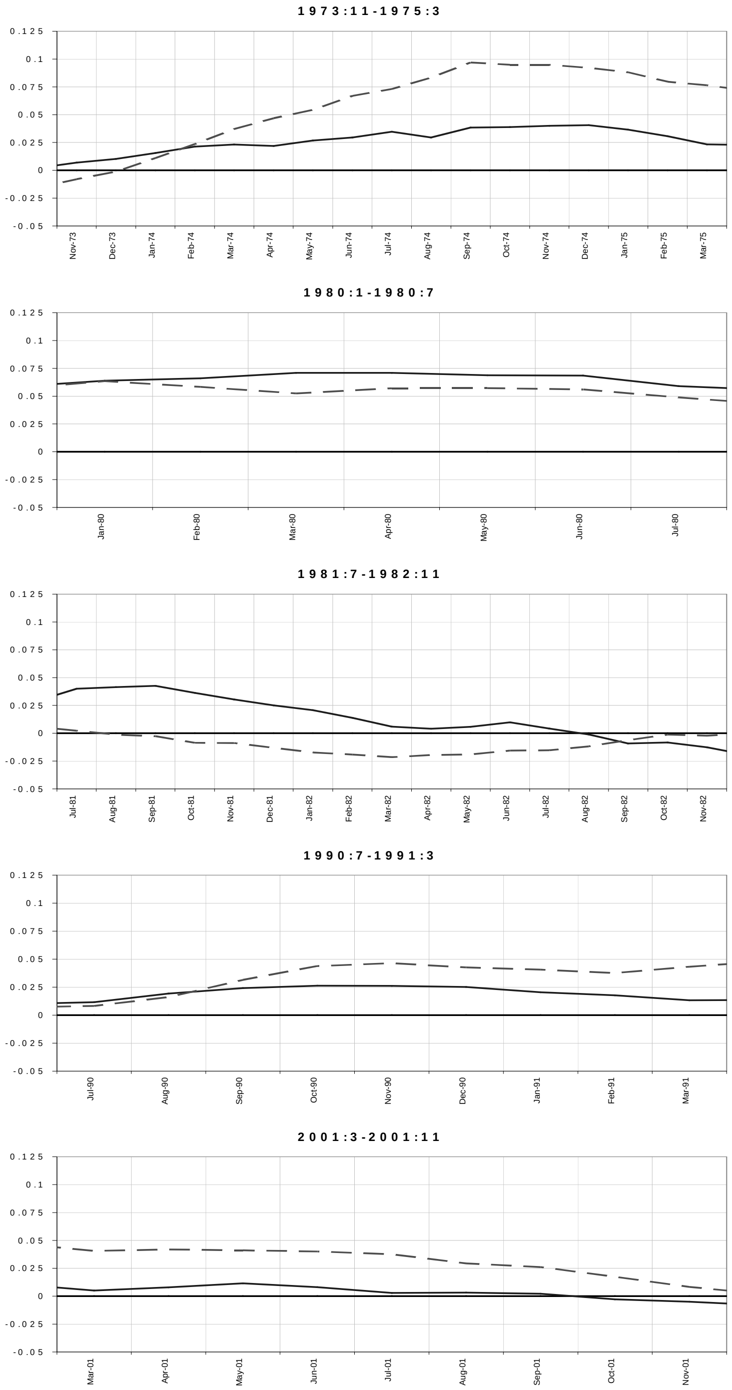
As is evident in the two graphs, the oil price increase seems to have affected neither inflation nor industrial production. Presumably other shocks

Figure 4: Historical decomposition of industrial production growth in the last five US recessions



Notes: periods are NBER recessions. Blue (solid) lines are the difference between actual and baseline industrial production growth, red (dashed) lines are the contribution of oil shocks

Figure 5: Historical decomposition of CPI inflation in the last five US recessions



Notes: periods are NBER recessions. Blue (solid) lines are the difference between actual and baseline CPI inflation, red (dashed) lines are the contribution of oil shocks hit the economy, more than offsetting the effect of the oil shock on output. For instance, still expansionary monetary conditions, strong productivity growth, wage moderation and globalization may have contributed to counterbalance that shock.

4 Robustness

In this section we test the robustness of our results by conducting several sensitivity exercises. They relate to the identification scheme, the measure of the shock and the sample size. Finally, we also re-estimate impulse responses using the methodology recently proposed by Jordà (2005, 2006). This experiment is meant to support the VAR lag specification and allows us to perform a formal test on the significance of the oil shock on CPI and industrial production.

As a first exercise we evaluate the effects on impulse responses of the parameter uncertainty in the construction of the matrix R (see equation (A7) in Appendix A) by holding the reduced-form parameters fixed at their ML-OLS point estimates. Figure 8 reports impulse responses to an oil price shock normalized to 1 per cent. More precisely, we perform Montecarlo integration drawing parameters from the posterior distribution, but we use only ML-OLS point estimates to construct the matrix R so that it is constructed once and for all and not at each parameter draw. The responses of CPI and industrial production are virtually unchanged but, as expected, the confidence bands around them are much tighter.

To understand how results would change when using a standard identification scheme, we report in Figure 9 responses to the oil price shock identified using a Choleski decomposition. We order variables as in Burbidge and Harrison (1984), that is oil price, interest rate, M2, wage index, CPI level and industrial production. Hence, the oil price shock affects immediately all the other variables.[11] The results are similar to those obtained under our futures-based identification (and to those in Burbidge and Harrison, 1984).[12] The variables react in a similar way under both identification schemes. The size of the industrial production median response is greater using futures. A possible explanation is that our identification captures oil supply shocks, which have a relatively strong effect on industrial production, while the Choleski ordering confounds oil demand and supply shocks. Note that while results are very similar, the results we get using futures are

[11] Note that in this case the responses would correspond also to the generalized impulse responses described in Pesaran and Shin (1998).

[12] Following Sims (1990) we estimate the system with the variables in log-levels, while Burbidge and Harrison (1984) estimate their VAR in log-differences. When we estimate the VARs in log-differences (besides interest rate), results do not change. To save on space we do not report them. They are available from the authors upon request.

Figure 6: Historical decomposition of industrial production growth (Mar.2003-Dec.2005)

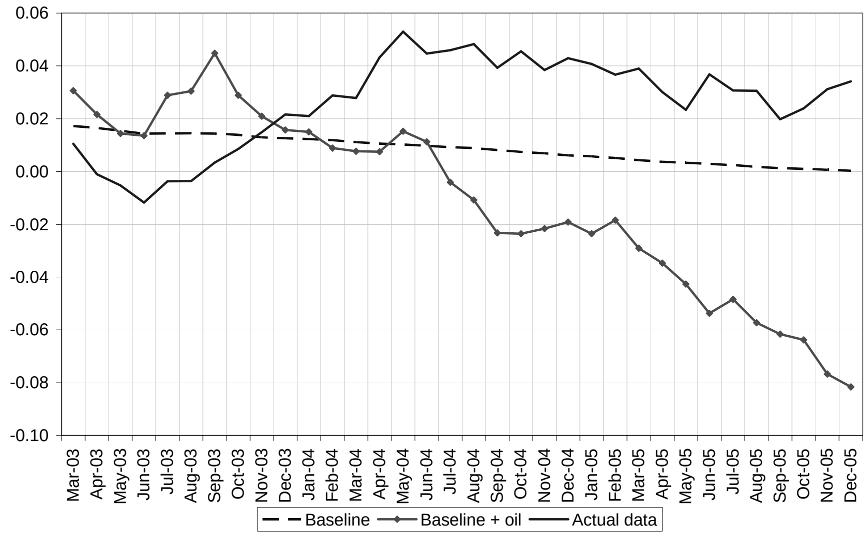


Figure 7: Historical decomposition of CPI inflation (Mar.2003-Dec.2005)

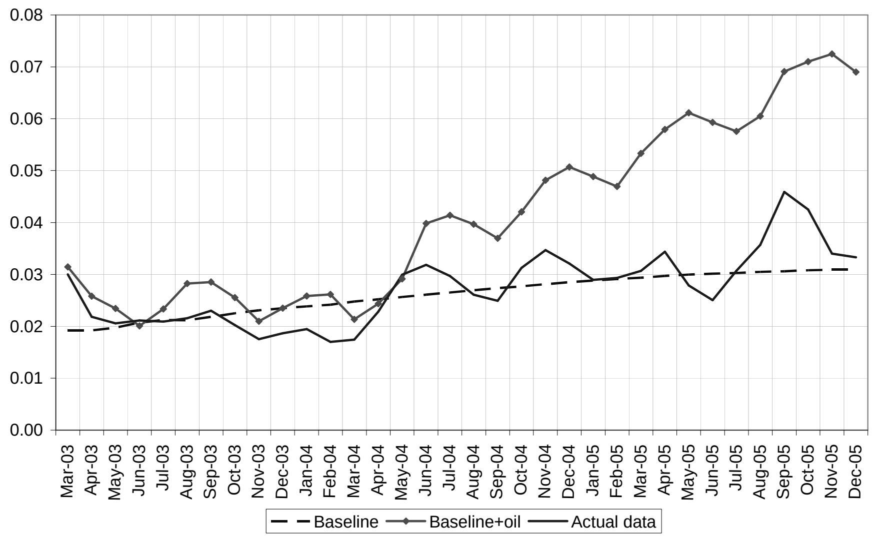
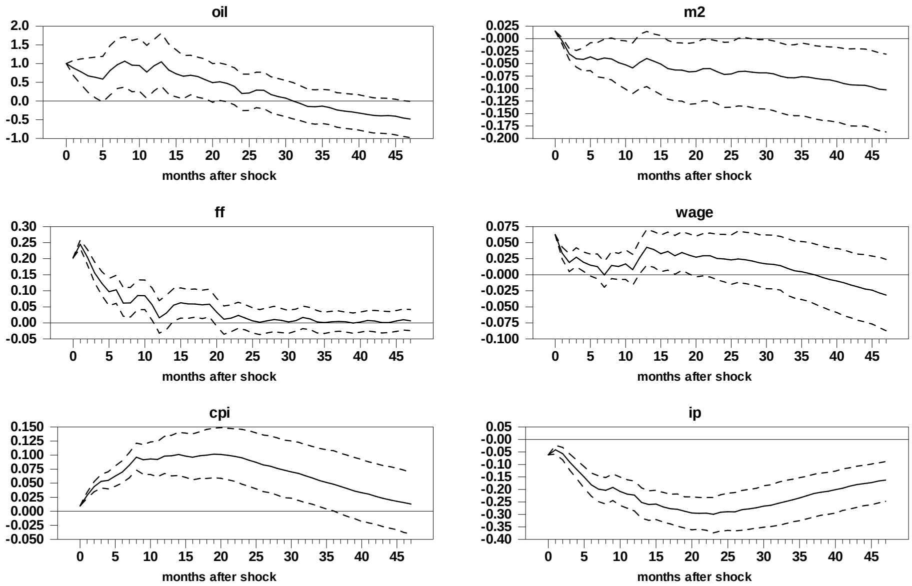


Figure 8: Estimated impulse responses and Montecarlo intervals whitout parameter uncertainty (futures-based identification)



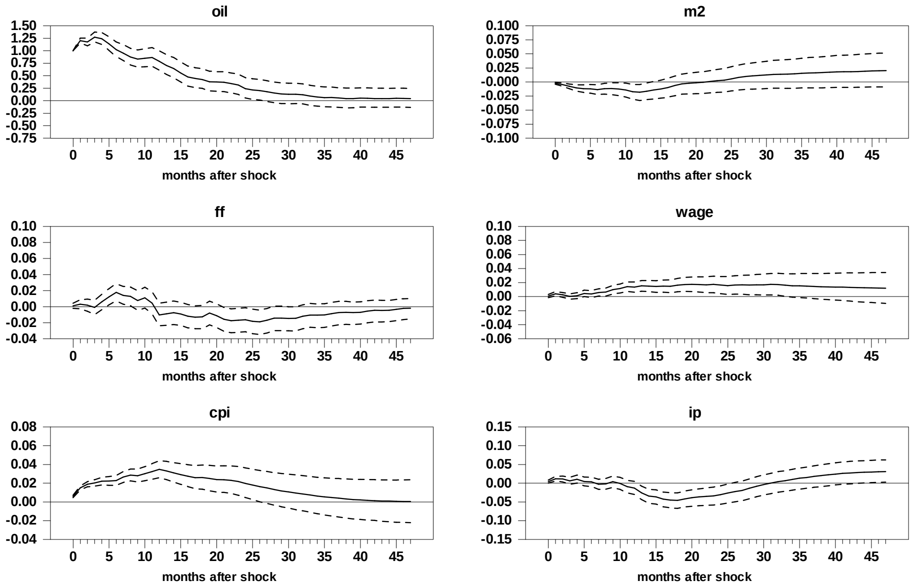
invariant to the Wold order of the variables.[13]

Since our sample period for measuring the oil price responses to change in futures prices is limited to the period starting in January 1986, we check whether results change when we estimate the VAR on that period. This check is also important because the historical decomposition shows a lower response of consumer price inflation from the early-1980s on. Figure 10 reports the responses to a 1 per cent oil price shock. Remarkably, the shapes of the impulse responses are similar across samples. In particular, responses of industrial production are almost indistinguishable, peaking roughly in the same period and with the same intensity. Consistently with the results of the historical decomposition, the size of the increase in the CPI is less in the smaller sample. The CPI is above the baseline and reaches a statistically significant peak seven months after the shock, but returns quite rapidly to the baseline.

We also estimate the VAR on the whole sample, but allowing for a deterministic change in trend CPI inflation in 1981, the period in which Hooker

[13] We have also tried alternative orderings of the variables. Results are not greatly affected.

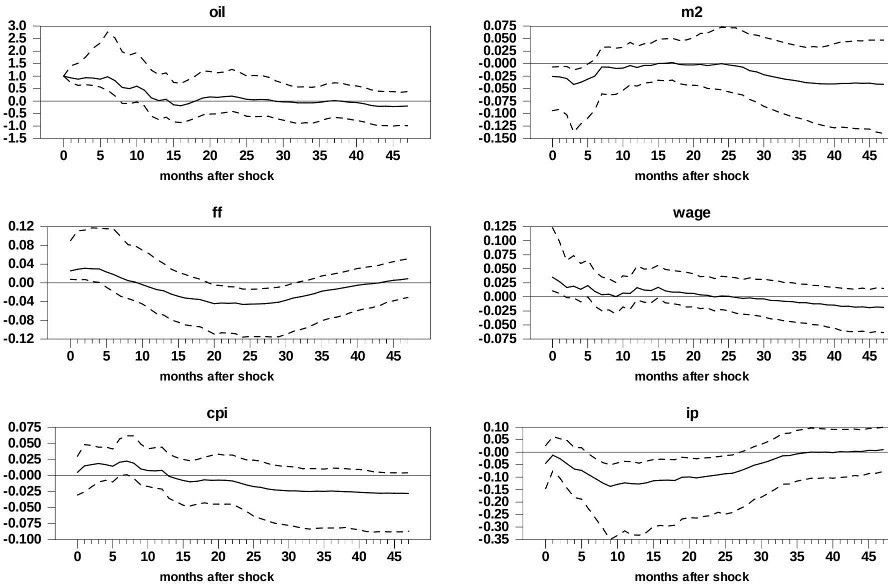
Figure 9: Estimated impulse responses and Montecarlo intervals (Choleski identification)



(2002) finds a structural break in the US Phillips curve, or in 1984, the break date envisaged by Bernanke and Mihov (1998). According to the results, not reported for brevity, the response of industrial production to the oil shock is literally unchanged. Indeed, the median response of the CPI seems more front-loaded and less significant with respect to that obtained without allowing for lower average inflation, but given the wide uncertainty around the median impulse responses, a formal test would reject the hypothesis of different sensitivity of the CPI to oil shocks.

Finally, we are aware that all estimated VARs suffer from a lag-truncation bias, which in some cases may be extremely severe. Intuitively, this bias arises because any specification forces to be zero some terms which should not. The OLS estimator would then adjust the estimates of the included lags to compensate for those that have been wrongly excluded. In a recent paper Jordà (2005) has argued that better multi-step predictions – denoted as "local projections" – can be found by direct estimations of different forecasting models, one for each step ahead, instead of iterating on a single

Figure 10: Estimated impulse responses and Montecarlo intervals, sample 1986-2005 (futures-based identification)

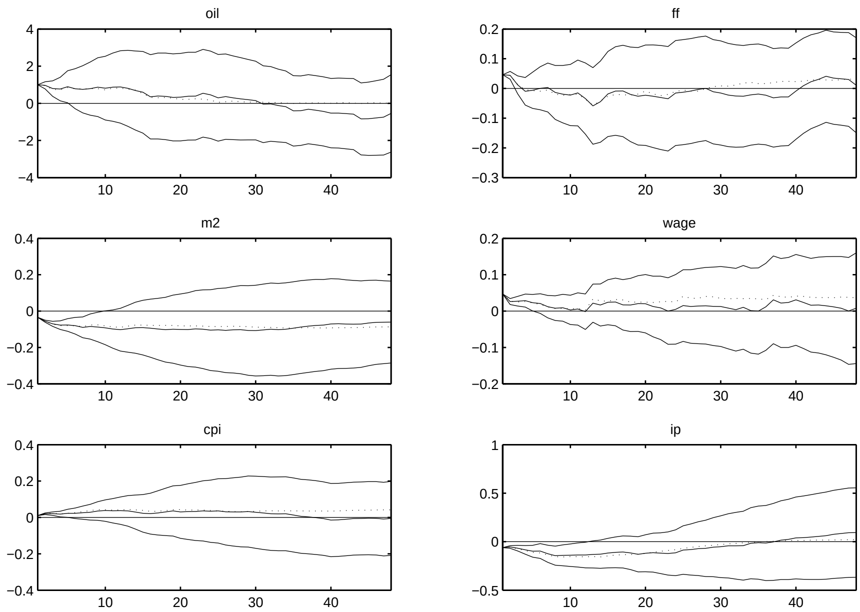


estimated model as in the VAR literature.[14] In Figure 11 we report – together with the impulse responses recovered using our VAR representation (dotted lines) – the impulse responses estimated using local projections and the corresponding 95 per cent confidence time-profile bands (solid lines) as in Jordà (2005).[15] Eyeball econometrics suggests that both approaches provide almost indistinguishable results, therefore supporting the assumption that our VAR does not suffer from the lag-truncation bias. Furthermore, by using local projections, we are also able to test the hypotheses that the oil shock has no effects on CPI and on industrial production. The value of the two $Wold - statistics$ are respectively 79.22 and 59.56, strongly rejecting the null of no effects.

[14] Indeed VAR procedure is optimal if the postulated model correctly represents the data generating process.

[15] Following Jordà (2006) we assume normality of residuals. Matlab codes to perform local projection estimates are available on the web on the Oscar Jordà homepage: http://www.econ.ucdavis.edu/faculty/jorda/index.html

Figure 11: Estimated impulse responses using linear projections



5 Concluding remarks

High frequency data have been used in VAR analyses to identify monetary policy shocks. In this paper we study the effects of daily oil supply news on US inflation and output. We use this high frequency data to form identifying restrictions for a monthly VAR. In particular, we impose that the response of oil prices to actual or potential oil supply shocks match that estimated in the futures market.

The analysis suggests the following main findings. First, news about possible oil supply disruptions have a stagflationary effect on the US economy. Impulse responses suggest that the CPI level increases in the first few months after the shock, while the industrial production level decreases in the first year. Second, oil shocks contributed to the stagflation episodes of the 1970s and to recent recessions. Historical decomposition shows that the oil shocks greatly help to fill the gaps between the baseline and the actual value of both inflation and industrial production growth. Third, over time, the effect of oil shocks on CPI inflation has diminished. This can be due to various factors: the smaller size and persistence of recent shocks (compared with the shocks in the 1970s), structural changes in the US economy and the different reactions of the monetary authorities.

Our results are subject to two caveats. First, given that we rely on linear relationships between variables, we do not explicitly take into account possible non-linearities. As suggested by some scholars (for example Mork, 1989, Hooker, 2002, Hamilton, 2003) the effect of oil shocks may be rather asymmetric, with oil price increases affecting the economy and price reductions not. Second, macroeconomic time series may be not stable over time and there can be structural breaks in the relationship between them (e.g. Hamilton, 1983, 1996, Hooker, 1996, Barsky and Kilian, 2004). We share these caveats with many of the contributions of the VAR literature we mention. Non-linearities could explain possibly different contributions of news about oil supply to output growth and CPI inflation in different periods. The analysis of such possible non-linear effects is beyond the scope of this paper and is left for future research.

Appendices

A Identifying oil supply shocks using futures

We identify oil supply shocks using information contained in futures contracts. The initial step is rather standard. From the estimated reduced-form VAR we get the structural form by relating the reduced-form residuals to the structural disturbances.

Consider the reduced form VAR:

$$A(L)\,Y_t = u_t, \tag{A1}$$

where Y_t is $G \times 1$, $A(L) = \sum_{j=0}^{\infty} A_j L^j$ and $A_0 = I$. We assume that $A(L)$, which is a $G \times G$ matrix, is invertible. Hence, the system can be written as:

$$Y_t = B(L)\,u_t, \tag{A2}$$

where $B(L) = A(L)^{-1}$.

We assume that the G reduced form errors u_t are related to structural disturbances ε_t as follows:

$$u_t = S\varepsilon_t, \tag{A3}$$

where S is a $G \times G$ full rank matrix. The VAR in equation (A1) can be made structural by writing it in terms of the structural shocks:

$$Y_t = B(L)S\varepsilon_t. \tag{A4}$$

Suppose the first column of S corresponds to the oil shock and call it α. The impulse response of all variables in the VAR to the oil shock is:

$$B(L)\,\alpha = \sum_{j=0}^{\infty} B_j \alpha L^j. \tag{A5}$$

The gth element of the $G \times 1$ vector of lag polynomials $B(L)\,\alpha$ traces out the response of the gth variable to the oil supply shock. The Bs are known, because they are implied by the reduced-form estimates. Hence, identifying the impulse response requires picking the G elements of α.

To identify oil supply shocks we use the information contained in the futures contracts in correspondence to events classified as supply shocks. The steps of the identification procedure are two: (a) we derive the response of the expected oil prices from the futures, (b) we impose the equality between the VAR impulse response of the oil prices to the oil shock and the response measured by the futures. Let us start by briefly illustrating point (b) and then point (a).

A.1 Matching responses of oil prices

Suppose that, in the case of no uncertainty, the response of the oil price at time $t+h$ to an oil price shock at time t is r_h, $h = 0, 1, ...G-1$. Hence:

$$B_{h,oil}\alpha = r_h, \tag{A6}$$

where $B_{h,oil}$ is the row of B_h corresponding to the oil price. We can stack these G equations to form:

$$R\alpha = r,$$

where the rows of R are the relevant row vectors $B_{h,oil}$ and the elements of r are the corresponding elements r_h. We get $B_{h,oil}$ from the reduced-form VAR estimates. The response of oil prices to an oil price shock, r_h, can be obtained by using the information contained in the futures.

The above system has G equations in G unknowns (the elements of α). Its solution, under the condition that R is of rank G, is:

$$\alpha = R^{-1}r. \tag{A7}$$

If R is not full rank, as happened to be the case here, it is necessary to add other restrictions in order to identify the system. In Appendix B we show that it is not possible to reject the null hypothesis of rank four.

In the next section we show how the response r_h of oil prices to oil price shocks can be measured directly from the oil price futures market and explain what further restrictions we impose on the system.

A.2 Measuring oil price shocks using futures

The oil price futures contract f for date $t+h$ is a bet on the oil spot price s on date $t+h$. Parties to the h-period contract agree in t on a price f_{t+h} for oil to be delivered at $t+h$. The standard no-arbitrage condition implies that:

$$0 = E_t\left[m_{t+h}\left(s_{t+h} - f_{t+h}\right)\right] \tag{A8}$$

where m is the stochastic pricing kernel.

The previous equation can be rewritten as:

$$f_{t+h} = E_t s_{t+h} + \frac{cov(s_{t+h}, m_{t+h})}{E_t\left(m_{t+h}\right)} \tag{A9}$$

This condition says that the futures rate is equal to the expected future funds rate plus a risk term.

We will focus on the change in oil futures prices $\Delta_{d_t} f_{t+h}$ $\left(\text{equal to } f_{t+h}^{d_t} - f_{t+h}^{d_{t-1}}\right)$ on the day d_t of events that we classify as oil supply shocks. Hence, as long as the risk term in equation (A9) does not change on the day of the event, we can write:

$$\Delta_{d_t} f_{t+h} = E_{d_t} s_{t+h} - E_{d_{t-1}} s_{t+h} \equiv \Delta_{d_t}^e s_{t+h} \tag{A10}$$

where $\Delta_{d_t} E s_{t+h}$ is the change in expectations about the spot price on the date $t+h$ due to the unanticipated event that has perturbed the oil market on date d_t.

In the VAR the expected oil price at $t+h$ conditional on information in the dataset at time t is:

$$E_t s_{t+h} = \sum_{i=0}^{\infty} B_{h+i,oil} S \varepsilon_{t-i} \tag{A11}$$

The change in the expectation $\Delta^e_{d_t} f_{t+h}$ from day d_{t-1} to d_t is due to changes in the expectations of shocks εs over this day, $\Delta^e_{d_t}\varepsilon_t$, given that all the past εs (ε_{t-1},ε_{t-2},...) are known at the beginning of the day. In order to single out the changes in expectations due to oil shock ε_{1t}, we can use equation (A10) and write:

$$\Delta_{d_t} f_{t+h} = B_{h,oil}\alpha\Delta^e_{d_t}\varepsilon_{1t} + B_{h,oil}S^*\Delta^e_{d_t}\varepsilon_t \tag{A12}$$

where α is the first column of S and the matrix S^*, is equal to S with the first column replaced by zeros. We assume that the second term is zero: news does not lead the market to reassess its view of the other shocks. We obtain:

$$\Delta_{d_t} f_{t+h} = B_{h,oil}\alpha\Delta^e_{d_t}\varepsilon_{1t} \tag{A13}$$

Combining equations (A13) and (A6) we get:

$$\Delta_{d_t} f_{t+h} = r_h\Delta^e_{d_t}\varepsilon_{1t} \tag{A14}$$

where $r_h = B_{h,oil}\alpha$ is the impulse response of the oil price to the oil price shock at horizon h. Since this equation holds for every h, we substitute out the unobserved quantity $\Delta^e_{d_t}\varepsilon_{1t}$ with $\Delta_d f_t/r_0\,(=\Delta_d s_t/r_0)$ to get:

$$\Delta_{d_t} f_{t+h} = \frac{r_h}{r_0}\Delta_{d_t} s_t \tag{A15}$$

The latter equation states that each day an oil supply shock hits the market, futures prices at different horizons should change proportionally. The factor of proportionality is the same for each shock, while the magnitude of the shock can obviously be different. We estimate this factor of proportionality from the data on futures contracts and use the normalization $[r_0 = 1\%]$ to obtain the estimated $\hat{r}_h$ in our identification strategy.

The above steps allow us to recover the point estimate of α. Given that α depends non-linearly, through R, on the reduced-form parameter estimates, the uncertainty surrounding the latter translate into large uncertainty about the former. Following Canova and De Nicolò (2002), we mitigate this large uncertainty by imposing the signs of the impact responses of three (Fed funds, money and wages) out of the six variables in the VAR. In particular, we impose that the signs cannot be the opposite of the OLS point estimates of α. Imposing such sign restrictions is consistent with the possibility of R not being full rank and is in line with Faust *et al.* (2004), who use high frequency futures data to identify a monetary policy shock in a VAR.[16] Importantly, we do not restrict the variables of interest (CPI and industrial production).

[16] Indeed, Faust *et al.* (2004) impose stronger restrictions than ours, since they impose ranges (and not just signs) for all the impact responses of the variables.

B Testing the rank of R

In order to test the rank of the matrix R, we rely on Cragg and Donald (1997).

The null hypothesis is $\rho(R) = k$ against the alternative that $\rho(R) > k$. The matrix is a non-linear function of the reduced-form parameter vector ϑ. Since we have an estimate $\hat{\vartheta}$ of ϑ, applying the same non-linear function we obtain $\hat{R}$.

Assuming that

$$T^{\frac{1}{2}} \left(\vartheta - \hat{\vartheta} \right) \longrightarrow_d N(0, V_\vartheta)$$

and knowing that

$$T^{\frac{1}{2}} \left(vec(\hat{R}) - vec(R) \right) \longrightarrow_d N(0, V_R)$$

where vec is for vectorization and

$$V_R = VCV(vec(R)) = \frac{dvec(R)'}{d\vartheta} V_\vartheta \frac{dvec(R)}{d\vartheta},$$

to test the rank of R we can use the statistic:

$$S(L) = T \min_{P \in \pi(k)} \left(vec(R) \hat{-} vec(P) \right)' \hat{V}_R^{-1} \left(vec(R) \hat{-} vec(P) \right)$$

where $\pi(k)$ is the space of all conformable matrices of rank k.

Under the null hypothesis, the test has a limiting χ^2. We therefore test the hypothesis $H_0 : \rho(R) = k$ rejecting the null if hypothesis if $S(k)$ exceeds the critical level α of a chi-square. Iterating the test for increasing values of k, we were not able to reject the hypothesis of $k = 4$, with $S(4) = 6.12$ and $p - value = 0.19$.

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