THE INSTABILITY OF THE OIL PRICE-MACROECONOMY RELATIONSHIP

FRANÇOIS LESCAROUX*
I.F.P.

ABSTRACT

The aim of this paper is to draw attention to the short-run instability of the oil price-macroeconomy relationship resulting from the oversimplification characterizing the compact models used in most recent studies.

Taking the U.S. unemployment rate as an example, we briefly review the theoretical mechanisms whereby oil price movements affect this measure and explain why the influence of the oil price shows high-frequency variations. Our central point is that oil price changes not only affect value-added and overall prices but also the distribution of value-added in employee compensation, nonlabor costs and profits, and hence unemployment.

We first propose a simple structural econometric model of the U.S. unemployment that confirms the great influence exerted on the labor market by the distribution of value-added using the unit profits and nonlabor costs as explanatory variables. Oil price movements have no influence on unemployment in this model. We interpret this result as a proof that they do not affect unemployment directly but via our explanatory variables.

To model the oscillations of the unemployment rate, we rely on an endogenous equilibrium-reverting mechanism based on an original spring force that differs from the one used in error-correction models. Rather than consider the value of the error correction term in \((t-1)\), we think it necessary to consider the value of the accumulated disequilibrium in \((t-1)\). In this way, we account for the amplitude of the disequilibrium and its duration. This is tantamount to applying to the unemployment gap an autoregressive model whose number of lags varies continually. This formula allows us to represent a significant portion of the nonlinear influences exerted on the unemployment gap.

We next estimate an unconstrained VAR model intended to represent the mechanisms of propagation on the labor market and its interactions with energy market. It appears that a great uncertainty is associated with the response of this system to an oil price shock. We explain this inconstancy as the consequence of a varying response of unit profits to oil price innovations over the business cycle. Indeed economic conditions do not offer the same capabilities of response at times when economic activity is accelerating and in deceleration phases, which is the source of a short-run instability in the oil price-macroeconomy relationship.

François Lescaroux
I.F.P. PhD student
Institut Français du Pétrole
228-232 avenue Napoléon Bonaparte
92852 Rueil-Malmaison
France
Email: francois.lescaroux@ifp.fr

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INTRODUCTION

Oil price movements are often considered as a major source of business cycle fluctuations since the mid-'seventies. A Granger-causal relationship has been identified (Hamilton, 1983) between oil price changes and variations in macroeconomic indicators such as GNP (negative correlation) and the unemployment rate (positive correlation) in the United States, with Granger-causality running from the former to the latter. As more data has become available, however, this relationship appears weaker; it is no longer statistically significant at conventional levels. A possible interpretation of this evolution is that a disruption occurred in the relationship between the oil price and the macroeconomy (Hooker, 1996, 1999), due either to changes in the conduct of monetary policy in the Volcker-Greenspan era, or to the declining importance of oil in the production of wealth (as a consequence of the structural evolution of the Western economies into post-industrialized societies and the optimization of energy use).

While we are sympathetic to this last idea of a long-term weakening of the relationship between the oil price and the macroeconomy, the aim of this paper is to draw attention to the short-term instability resulting from the oversimplification characterizing the compact models used in most recent studies.

Taking the U.S. unemployment rate as an example, we briefly review the theoretical mechanisms whereby oil price movements affect this measure and explain why the influence of the oil price shows high-frequency variations. Our central point is that oil price changes not only affect value-added and overall prices but also the distribution of value-added in employee compensation, nonlabor costs and profits, and hence unemployment.

We propose a simple structural econometric model of the U.S. unemployment that confirms the great influence exerted on the labor market by the distribution of value-added. Over time, the level of unemployment equilibrium fluctuates following long-term profitability trends, but our view is that shifts in the natural level only explain a relatively small share of unemployment rate variations; imperfections in the labor market lead to imbalances in the distribution of wealth, which cause short-term fluctuations in employment in the spirit of Goodwin’s “Growth Cycle” model (1967). Monetary policy and the intensity of demand also affect out-of-equilibrium unemployment rate variations. To model these oscillations, we rely on an endogenous equilibrium-reverting mechanism based on an original spring force that differs from the one used in error-correction models (Granger and Engle, 1987). Oil price movements have no influence on unemployment in this model. We interpret this result as a proof that they do not affect unemployment directly but via our explanatory variables.

We further estimate an unconstrained VAR model intended to represent the mechanisms of propagation on the labor market and its interactions with energy market. It appears that a great uncertainty is associated with the response of this system to an oil price shock. On one hand, oil price variations affect unemployment through a monetary channel, acting first on inflation, which induces an interest rate response. On the other hand, they exert an unstable influence on unit profits and hence on unemployment. We explain the inconstancy of the latter relationship as the consequence of a varying response of unit profits to oil price innovations over the business cycle. Indeed economic conditions do not offer the same capabilities of response at times when economic activity is accelerating and in deceleration phases, which is the source of a short-run instability in the oil price-macroeconomy relationship.

A SHORT HISTORY OF THE ANALYSIS OF OIL PRICE VARIATIONS EFFECTS

Since the mid-'seventies, oil price movements have been considered by many economists and econometricians as a major source of business cycle fluctuations. Since then, great efforts have been made to analyze the mechanisms whereby oil price shocks affect the macroeconomy and to measure the impact of these shocks on economic growth (see the series of detailed reviews by Jones, Leiby and their co-workers, 1996, 1997, 2002, 2004).
Initial studies aimed to elucidate the new oil shock phenomenon by exploring its effects on demand: a hike in the crude price was taken as an exogenous inflationary shock (Pierce and Enzler, 1974) or as a transfer of wealth from importing to exporting countries (Hickman, Huntington and Sweeney, 1987). By and large, these analyses identified the following consequences: a slowdown in domestic demand (hence higher unemployment and lower GDP) and inflationary pressure (a risk of tighter monetary policy).

An increase in the oil price also affects supply (Rasche and Tatom, 1977 a, b and c), because energy is one of the basic inputs in the production process. When the cost of energy goes up, industry must pay more to produce goods at the same level of quantity as before the shock.

The Real Business Cycle theory developed in the 1980s reinforced the interpretation whereby oil shocks were supply shocks and most subsequent studies were situated within this theoretical framework.

A granger-causal relationship was identified (Hamilton, 1983) between oil price changes and variations in macroeconomic indicators such as GNP (negative correlation) and the unemployment rate (positive correlation) in the United States, with granger-causality running from the former to the latter. Burbidge and Harrison (1984) came to the same conclusion but using a slightly different approach. As more data has become available, however, this relationship appears weaker; it is no longer statistically significant at conventional levels.

Mork (1989) was the first to propose filtering the oil price signal to restore the causal relationship identified by Hamilton. With data running through 1988, he showed that only oil price increases granger-caused GNP variations, while oil price collapses had no significant impact on economic activity.

This asymmetry was soon justified by indirect influences, which affect economic growth whenever prices go up or down. The direct and indirect effects are supposed to amplify in the first case and to compensate each other in the second. The main explanations rely on adjustment costs resulting from the reallocation of factor input across sectors and plants (Lilien, 1982, Hamilton, 1988, Loungani, 1986, Loungani and Mahidhara, 1997, Davis and Haltiwanger, 2001). The postponement of irreversible investment decisions (Ferderer, 1996) when oil price variations create uncertainty is another mechanism that might contribute to this asymmetry.

In the nineties, however, oil price increases no longer granger-caused GNP or unemployment variations (Hooker, 1996). Three interpretations are possible. First, oil prices have never been an important source of fluctuations in economic activity and their impact was overstated (Tobin, 1980, Darby, 1982, 1984, Bohi, 1991). Second, oil price shocks affect the macroeconomy directly or indirectly through various channels. As a result, this relationship is more complex might have been expected, considering the relatively short experience of the ‘eighties. This line of research tries to identify a method of filtering the oil price signal that expresses asymmetrical and non-linear influences such as adjustment costs resulting from sectorial imbalances or the postponement of irreversible investment decisions when oil price variations generate uncertainty. Lee, Ni and Ratti (1995) and Hamilton (1996) proposed specifications intended to integrate the "surprise factor" associated with oil price changes while Ferderer (1996) used the price volatility of some petroleum products. But Hooker (1999) showed that these measures did not granger-cause GDP variations in the 1980-1998 sample. The third interpretation is that, in about 1980, a break occurred in the relationship between the oil price and the macroeconomy (Hooker, 1996, 1999), due either to changes in the conduct of monetary policy in the Volcker-Greenspan era, or to the structural evolution of Western countries into post-industrialized societies.

**THE SHORT TERM INSTABILITY OF THE OIL PRICE-UNEMPLOYMENT RELATIONSHIP**

We are sympathetic to this last idea of a long term weakening of the oil price-macroeconomy relationship corresponding to the declining importance of oil in the production of wealth (as a consequence of the structural evolution of the economies and the
optimization of energy use), but our goal here is to highlight a short-run instability resulting from the oversimplification of the compact models used in most of the recent studies.

In point of fact, the analysis of mechanisms governing the impact of an oil price hike on employment and prices systematically relies on a "decision tree" structure.

From the standpoint of the supply of goods and services, any increase in the cost of an intermediate product adds to production costs. Businesses are tempted to pass this increase on to customers by raising their selling prices to protect their profits. This creates inflation, especially if wages are indexed to prices, which produces an inflationary spiral. Furthermore, if one business enterprise practices such a risk transfer but its competitors do not, then its competitiveness – hence its sales and hiring – may be adversely affected. Businesses can also compensate for this increase by lowering wages, which leads to higher unemployment, or by accepting narrower profit margins, in which case they are likely to invest less and eventually reduce their hiring.

When consumer energy prices rise, household purchasing power tends to fall, with a slackening of demand for goods and services. However, the impact depends on prior choices and varies according to the responses of inflation, real wages, margins and taxes to the oil price hike.

For Real Business Cycle theoreticians, the mechanisms differ slightly but the chain of events remains the same. When the price of an intermediate product rises, real wages increase more slowly or fall and the labor supply contracts, because workers prefer to wait to obtain a greater return on their labor.

The observation in space and over time of the strategies selected reveal major differences between countries (Criqui and Percebois, 1988), and in a country between shocks (Bohi, 1991). In France, producers were hardest hit by the first oil shock in 1973, reporting lower profits, whereas consumers footed most of the bill for the second shock (Chaouat, 1982, 1983).

Different factors are involved in decision-making. Past experience is useful in eliminating choices with the most adverse effects (e.g. price control in the 1970s). In addition, government economic policy is shaped by ideological convictions and sometimes by political considerations in the run-up to elections. Above all, current economic conditions determine future trajectories. When demand accelerates, labor negotiates from strength and is ill inclined to accept a wage reduction whereas it is easier for producers, whose sales are growing, to agree to lower the share of their profits in value-added. Inversely, in a sluggish economy, businesses cannot accept any further narrowing of their margins, but are in a good position to impose lower labor costs by reducing wages and/or employment.

Short- and long-term national budget trends also influence how employment and the economy at large respond to an increase in oil prices. A country whose government budget is not running a deficit (or at least not too large a deficit) can substantially offset heavier energy bills, at least temporarily, by adjusting energy and/or corporate taxation or granting subsidies to the most oil-dependent sectors.

Theoretical and historical analysis shows the various possibilities as to how an oil price hike can propagate through an economic system and affect it and the various immediate consequences implied by the choices made. It also clearly establishes that the problem involves two phases: initially, the price variation affects value-added and inflation as well as the distribution of wealth, then these factors determine unemployment. Now the estimated models combine these two stages into one and attempt to directly measure the impact of crude price variations on employment or GDP growth, especially the bivariate models used recently by Hamilton (2003). Hooker (1999) also interprets the instability of the relationship between oil prices and either output or unemployment as a sign of an indirect influence via other macroeconomic variables. He mentions interest rates and inflation.

Models that integrate wage fluctuations as well as inflation obtain results that offer more scope for interpretation and analysis. However, a few key aspects of the problem are not addressed, since compact models systematically fail to integrate some of the basic fundamentals that influence how unemployment responds to an oil shock including, among
others, tax trends (both energy-related and in general), the subsidization of business enterprises and mostly corporate profits.

The importance of profit margin fluctuations was recently heavily documented in analysis of price-unemployment interactions using a Phillips curve (Brayton, Roberts and Williams, 1999, Gordon, 1998, Eller and Gordon, 2002).

However, with the exception of the Markov state-switching models (Raymond and Rich, 1997), the compact empirical models that we found recently in the literature proceed as if the impact of oil price variations on unemployment was systematic and constant. In other words, these evaluations yield a sort of average impact that, in all likelihood, has no economic significance for the past and no predictive power for the future.

In analyzing the influence of the oil price using a Markov state-switching model, Raymond and Rich (1997) concluded that oil price hikes mainly influence economic activity during low-growth periods and do not trigger a shift from a state of high or normal growth to one of low growth. The authors do not offer an economic interpretation of these results. It seems to us that their conclusion can be explained, at least partially, by seeing how the distribution of value-added between profits, wages and nonlabor costs evolves during the business cycle.

We consider this point to be especially important, in light of the recent increase in the oil price. The rise in the barrel price caused by expanding world crude demand does not seem to have affected American economic growth. In our opinion, the reason is that the increase occurred when the U.S. economy was in a phase of acceleration, which leaves plenty of leeway for determining how to pay the oil bill. In contrast, the first two oil shocks were caused by a contraction of supply and occurred during a period of decelerating economic activity when profits were falling (their level and as a percentage of total value-added).

A SIMPLE STRUCTURAL ECONOMETRIC MODEL OF UNEMPLOYMENT

a. Presentation of the data and preliminary processing

To back up our opinion, we built a simple structural econometric model of adjustments in the U.S. labor market. Our goal is to show that the manner in which created wealth is distributed among different economic agents is a major cause of variations in unemployment and its level and that, consequently, various possible scenarios must be evaluated when analyzing the consequences of a high-frequency shock on employment.

To Real Business Cycle theoreticians, we grant that the level of unemployment equilibrium fluctuates over time, following long-term profitability trends. This being said, we think it preferable to approach the problem from the standpoint of labor demand and that the shifts in natural level only account for a relatively low share of variations in the unemployment rate. Looking at the short term and remaining within neoclassical tradition, we favor a more conventional interpretation of employment movements: wage rigidities and the inertia of nonlabor costs lead to imbalances in the distribution of wealth which, in the short-run, lead to adjustments in the labor market. This part of our model also reflects the influence of Goodwin’s “Growth Cycle” model (1967) and expresses the effects of the repartition of value-added on unemployment. We also borrow the idea from Keynes that demand for goods and services influences employment and that fluctuations in labor demand originate when economic activity accelerates and decelerates in successive phases. For the short term, we consider "classical" unemployment (due to insufficient supply) as well as "keynesian" unemployment (due to insufficient demand).

The variable whose variations we seek to clarify is the unemployment rate for civilians aged 16 and over. To eliminate the influence of demographic trends and the effects of women entering the labor market from our time sample, we estimated a measure of the unemployment rate that neutralizes changes in the composition of the labor force (see Brayton, Roberts and Williams, 1999, p. 36). We used an average that was weighted for the rates of unemployment associated with five categories of workers: aged 16-19, men aged 20 to 24,
women aged 20 to 24, men over 25 and women over 25. For weights, we used the respective
shares of 1993 employment reported for these groups.

The unemployment rate for workers aged 16 and over and its measurement, adjusted
for demographic factors, are shown in Figure 1. Subsequently, we use the adjusted
unemployment rate but do not always say so (in equations, the exponent "adj" serves as a
reminder).

Our explanatory variables are as follows. We used the shares of corporate profits
(with inventory valuation and capital consumption adjustments), compensation of employees
and nonlabor costs (consumption of fixed capital, net interest and miscellaneous payments,
taxes on production and imports less subsidies plus business current transfer payments) per
unit of real gross value-added of non-financial domestic corporate business (Figure 2). We
also used the implicit price deflator for GDP, the Treasury bills rate at constant, fixed
maturity (5-year) and a variety of oil price measures (PPI for crude oil, PPI for petroleum
products, Mork oil price variable). See the data appendix for further details.

b. Estimation of the long-run equilibrium unemployment rate

We first verified the existence of a long-run equilibrium relationship between the
unemployment rate and the unit profits from current production. The estimated relationship
expresses in an equation what Chancellor Schmidt said in words: “Today’s profits allow
tomorrow’s investment which will create employment the day after tomorrow”. The
cointegration relationship is:

\[ U_{\text{adj}} = -19.23 - P_{\text{profits}} + 7.99, \]

Sample: 1948:1-2004:1; R² = 0.27; ADF Test on the residuals (without trend and intercept,
one lagged difference): τ = -4.5301 (critical value at the 5% level tabulated by McKinnon, 1991: -3.9475).

where \( U_{\text{adj}} \) is the demographically adjusted unemployment rate and \( P_{\text{profits}} \) are the unit
profits (t-stats are reported in brackets although their distribution is not standard for this kind
of equation). Subsequently, we consider the residual of this equation to be the unemployment
gap (notation: \( U^g \)).

This cointegration relationship defines the trajectory of equilibrium unemployment. It
expresses the mechanisms explained above in the long-run. Let us suppose, for example, that
the economy experiences a rise in productivity (driven by a reduction in production costs or
nonlabor costs, efficiency gains, demand acceleration, etc.). The value-added will rise and the
share of profits will increase first. Businesses will soon hire for two reasons: because the
economic conditions encourage them and their benefits allow them to do so. But the rise in
employment results in the appearance of hardness on the labor market. The growing scarcity
of resources induces a rise in their rental price, i.e. wages and salaries. The increase in
employment and in wages and salaries cause the unit labor cost to rise in turn. So, unit nonlabor costs decrease. If the rise in productivity is temporary, unit profits will decline to a lower level than before and businesses will reduce employment to restore their margins by cutting unit labor costs; this results in a cycle in the spirit of Goodwin (1967). But if the productivity gains are permanent, the unit profits will be higher after this rise in profitability than before, despite the subsequent increase in unit labor costs, and the equilibrium unemployment rate will shift to a lower level. Symmetrically, if the economy enters a period of low profitability, real wages might grow faster than productivity and unit profits would decline first; then businesses would restore their margins by cutting labor costs and the equilibrium unemployment rate would shift to a higher level. The equilibrium unemployment rate estimated with the previous relationship shows high-frequency variations. We applied a low-pass filter (Hodrick-Prescott, parameter of 1600) to extract its trend (Figure 3).

Figure 3: Long-run relationship expressing the U.S. unemployment rate as a function of unit profits. Right scale: adjusted unemployment rate (fine line), estimated equilibrium level (short, bold dotted line), filtered equilibrium level (long, bold dotted line). Left scale: residual (bold line).

Our estimation of the equilibrium unemployment rate approximately fits the one proposed by Gordon (1997). He derives different estimates of the NAIRU (Non-Accelerating-Inflation Rate of Unemployment) corresponding to different values of its supposed volatility. The results are shown graphically, which makes it difficult to make a direct comparison. But the trajectories of his measures are similar to ours and the values are close to Gordon’s preferred specification (for a standard deviation of 0.20). Our result differs with his only at the beginning of the sample. His estimate of the NAIRU falls from 6% in 1955 to about 5.40% at the beginning of the ’60s and then comes up again. For this short period, our estimate is closer to that made by Staiger, Stock and Watson (1997). The mathematical expectation of their NAIRU declines from the beginning to the middle of the ’60s, then goes up till the early ’70s (we find ourselves in the low part of their 95% confidence interval).

c. Explaining the short-run variations of the unemployment gap

We then estimated a model of the short-run variations in the unemployment gap. Our modeling of short-run variations relies on an equilibrium-reverting mechanism with a original spring force, different from the lagged error-correction term used in the error-correction models (ECM; Granger and Engle, 1987). We think that this approach neglects one of the most important factors in a dynamic system: time. According to us, the spring force $S_t$ which
pushes an economic variable toward its equilibrium value at time $t$ is not proportional to the value of the desequilibrium observed “yesterday”, at time $(t-1)$, but to the value of the accumulated desequilibrium observed since the departure from equilibrium, i.e. the integral of the residual from the long-run relationship between the last date when it was zero and $(t-1)$:

$$S_t = -k \cdot \int_{t_0(t)}^{t-1} \frac{U^s(\tau)}{d\tau} = -k \cdot \sum_{i=t_0(t)}^{t-1} U^s(i) = -k \cdot \text{Int}(U^s, t),$$

where $t_0(t)$ is the last date before $t$ such that:

$$U^s(t_0(t)) = 0.$$

Of course, this measure is strongly correlated with the lagged value of the error-correction term used in the ECM. The introduction of such a spring force in a model of unemployment gap variations is tantamount to using an autoregressive model of the unemployment gap with the length of the AR structure varying continually.

Our estimation procedure is as follows. The exogenous variables of interest were the unit profits and nonlabor costs (in first differences), the real Treasury bill rate (in first differences), the GDP deflator (in variation rate) and an oil price measure. We first estimated unconstrained models of unemployment gap variations. All used lagged variations of the endogenous variable (from 1 to 4 quarters) and the spring force. To this basic structure, we added the lagged values (and occasionally the instantaneous ones) of one, some or all of the exogenous variables and the lagged value (1 quarter) of the error-correction term used in an ECM. We step-by-step eliminated the non-significant variables in each of these models.

The results were converging toward the relationship:

$$D(U^s) = -0.0025 \cdot \text{Int}(U^s, t) + 0.42 \cdot D(U^s)_{t-1} - 8.33 \cdot (\pi_t - \pi_{t-2})$$

$$+ 0.03 \cdot D(TBill)_{t-2} + 42.69 \cdot D(Cost^NL)_{t-1}$$

[2]

Sample (adjusted): 1954:4-2003:3; $R^2 = 0.56$; S.E.E. = 0.219;

Durbin-Watson stat = 1.92; $Q_{LB}(8) = 11.28$ (proba = 0.19).

where $Cost^NL$ is unit nonlabor costs, $TBill$ is the 5-year real Treasury bill rate and $\pi$ is the quarterly rate of variation of the GDP deflator. (The $R^2$ is relatively small because this equation explains the variations of the unemployment gap. If we estimate the level of the unemployment gap instead, using its lagged value as an explanatory variable, the $R^2$ is 0.97 and the value of the coefficient associated with the lagged value of the endogenous variable is 1.)

The residuals and the squared residuals show no sign of autocorrelation. Furthermore, the BDS test does not reject the null hypothesis of independence for all tested values of the embedding dimension; our spring force seems to enable us to catch a significative part of the non-linear influences that act on the unemployment gap. The usual tests (recursive estimates, CUSUM, CUSUM Square, rolling Chow test) indicate that the relationship is not unstable (see the Statistical Appendix for the BDS test and the rolling Chow test; other tests were not reported to save space but are available on request).

The signs of the coefficients are as expected: a rise in nonlabor costs or interest rates induces a rise in the unemployment gap and an acceleration in prices means a reduction in the unemployment gap.

The coefficient associated with the spring force is negative and relatively small; the return toward equilibrium is slow, which is consistent with persistent (but not permanent) effects following a shock. We compared our spring force with the error-correction term used in the ECM. The latter was not near significant and soon eliminated by the step-by-step estimation procedure (the $t$-stat associated with its coefficient is 0.19 if we introduce it into the final relationship).

Inflation was first introduced into the model as the rate of variation of the GDP deflator. The step-by-step estimation procedure eliminated all of the lagged values except the one in $(t-2)$. Its value was very close to the one for the instantaneous rate of variation of the GDP deflator, with an opposite sign. A Wald test confirmed that they sum to zero. So this
restriction was introduced in the final form of the relationship. The coefficient associated with
the variation of the inflation rate between \((t-2)\) and \(t\) is marginally significant, at the 10% level.
A rise in the inflation rate means an acceleration in prices, hence an intensification of
economic activity and a reduction in unemployment. But periods of deceleration follow
periods of acceleration and inflation is neutral in the long-run (the mean of \(\pi_t - \pi_{t-2}\) is zero for
our sample).

![Figure 4: Adjustment of the short-run model as in relation [2] (right scale: observed
values in thin line, simulated values in dotted line; left scale: residual).](image)

The distribution of the residuals is asymmetrical and shows positive extreme values
for the most part (Figure 5). Dispersive shocks could be responsible for these peaks. The dates
of these perturbations do correspond to periods of economic turmoil.

![Figure 5: Histogram and descriptive statistics for the residuals from equation [2]](image)

<table>
<thead>
<tr>
<th></th>
<th>Value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
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<tr>
<td>Median</td>
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<tr>
<td>Maximum</td>
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<td>Minimum</td>
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<td>Skewness</td>
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<td>Jarque-Bera</td>
<td>28.738</td>
</tr>
<tr>
<td>Probability</td>
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</tr>
</tbody>
</table>

But such events are too rare to test statistically if oil price variations are responsible
for them. Moreover, many influences could have systematically caused a great stir in the
economic system during these periods. What comes to mind in particular are changes in the
money supply and especially in the exchange rates. For example, the dollar devaluation on
February 12, 1973 may have contributed to the residual peak observed in the last quarter of
1974 and first quarter of 1975. And the tightening of U.S. monetary policy in the early '80s
and the subsequent take-off of the dollar may partially explain the series of small peaks
visible in the residuals.
We tried to neutralize the principal residual peaks with dummy variables: the first is zero everywhere except for first quarter 1958 (the assigned value is one) and the second is zero everywhere except for last quarter 1974 and first quarter 1975 (its value is one). The coefficients of the other variables are not affected and not statistically different from the ones reported above (results not reported to save space). But the Jarque-Bera test does not reject the hypothesis of normal distribution (value: 0.463, probability: 0.793). Moreover, the coefficient associated with the price acceleration becomes strongly significant (t-stat: -2.73, probability: 0.007). That is the reason why we retained this explanatory variable in the equation reported above.

As we expected, our different attempts to introduce oil prices into the model did not succeed. We tried various measures to catch the oil price influence (such as nominal and real refiner acquisition costs, the PPI for crude oil or for petroleum products, relative or not to the GDP deflator) and various specifications of these measures (such as the variation, the rate of variation and the year-over-year variation). No combination of these variables was significant, even marginally.

d. Dynamic simulations

In dynamic simulations (Figures 6 to 15), this equation performed well. It has to be noted that the spring force was also estimated in a dynamic manner in the simulations performed. However, the model was estimated on the whole sample so as to be able to start the simulations in the ‘60s; it isn’t thus a true out-of-sample forecast exercise.

Our forecasts clearly deviated from reality only during two time periods: the second half of the 1960s, where we were several years too early in anticipating the future rise in unemployment, and the end of the 1990s, where we again overestimated the unemployment gap. The successive dynamic simulations performed enable us to identify precisely the time periods when our model fails to explain the evolution of the unemployment gap. Figures 6 to 8 show that the forecasts overestimate unemployment from about 1966 to 1969 and figures 14 and 15 show that we again overestimate unemployment from the middle of 1997 to the end of 2000.

Our object was not to build a perfect model of U.S. unemployment, but to show that oil price fluctuations impacted it indirectly – by influencing the distribution of wealth – and not systematically, therefore we will not attempt to correct these biases. Yet we did identify possible explanations. The "Go for Growth" policy introduced by John F. Kennedy in 1962 and ‘63 may be one of the reasons causing the unemployment observed in the second part of the 1960s. The increase in government expenditure amplified demand and thereby gave a temporary boost to employment. As for the late 1990s, the strong economic growth observed then, when inflation was under control, may have cushioned the effect of the decrease in the share of profits. This interpretation is supported by the fact that, when a non-monetary variable expressing the acceleration of economic activity is included (e.g. the first difference of the stocks variations relative to the GDP level), it is possible to reduce the divergence between simulation and reality. That production costs decrease because a number of raw materials saw their price fall (in the wake of the Asian crisis) may be another contributing factor.

So this simple model seems to explain quite well the fluctuations of the unemployment gap since the mid-'fifties. It relies on an endogenous equilibrium-reverting mechanism and uses as explanatory variables the variations of nonlabor costs, to catch a short-run “classical” unemployment, an indicator of the intensity of economic activity, to catch a short-run “keynesian” unemployment and the variations of an interest rate, to catch the influence of the Fed. Oil price fluctuations certainly affect these explanatory variables by various channels but they have no direct influence on the unemployment gap.
A COMPACT UNCONSTRAINED MODEL WITH OIL PRICES

A look at trend graphs for the respective proportions of profits, nonlabor costs and employee compensation per unit of value-added would be quite enough to confirm our previous criticism of attempts to measure a direct impact. There is nothing to indicate that crude price variations have a systematic influence on the distribution of wealth.

Nevertheless, we proceeded to estimate a vectorial autoregressive model (VAR) intended to represent the mechanisms of propagation on the labor market and its interactions with energy market to reinforce this criticism. Its general form is:

$$A(L) \cdot X_t = \varepsilon_t,$$

where $A(L)$ is a matrix polynomial in the lag operator, $X$ is the column vector of endogenous variables and $\varepsilon$ is a vector of innovations.

The endogenous variables of interest in $X$ are the logarithm of the PPI for refined products relative to the GDP deflator ($\text{Oil}$), the real Treasury bills rate ($\text{TBill}$), the logarithm of the GDP deflator ($P$), the unit profits ($\text{Profits}$) and labor cost ($\text{Comp} L$) and the adjusted unemployment rate ($\text{U}^{\text{adj}}$). All are in first differences.

We refrained from introducing the spring force in the model as an exogenous variable, because it would have affected the impulse analysis.

The autoregressive structure chosen by maximizing the log-likelihood includes 7 lags. Generally, the VAR models estimating the oil price-macroeconomy relationship include between 4 and 8 lags with quarterly data.

We performed pairwise Granger-causality tests (Granger, 1969) to identify interactions between our variables (Table 1).

**Table 1: Pairwise Granger-causality tests**

<table>
<thead>
<tr>
<th>Dependent variable: $D(\log\text{(Oil)})$</th>
<th>Exclude</th>
<th>$\chi^2$</th>
<th>df</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(\text{TBill})$</td>
<td>9.87</td>
<td>9</td>
<td>0.36</td>
<td></td>
</tr>
<tr>
<td>$D(\log\text{(P)})$</td>
<td>18.92</td>
<td>9</td>
<td>0.03**</td>
<td></td>
</tr>
<tr>
<td>$D(\log\text{(Profits)})$</td>
<td>10.38</td>
<td>9</td>
<td>0.32</td>
<td></td>
</tr>
<tr>
<td>$D(\text{Comp} L)$</td>
<td>9.35</td>
<td>9</td>
<td>0.41</td>
<td></td>
</tr>
<tr>
<td>$D(\text{U}^{\text{adj}})$</td>
<td>10.87</td>
<td>9</td>
<td>0.29</td>
<td></td>
</tr>
<tr>
<td>All</td>
<td>50.24</td>
<td>45</td>
<td>0.27</td>
<td></td>
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</tbody>
</table>

<table>
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<th>Dependent variable: $D(\text{TBill})$</th>
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<th>$\chi^2$</th>
<th>df</th>
<th>Prob.</th>
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</thead>
<tbody>
<tr>
<td>$D(\log\text{(Oil)})$</td>
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<td>9</td>
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<tr>
<td>$D(\log\text{(P)})$</td>
<td>23.84</td>
<td>9</td>
<td>0.01***</td>
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</tr>
<tr>
<td>$D(\log\text{(Profits)})$</td>
<td>10.91</td>
<td>9</td>
<td>0.28</td>
<td></td>
</tr>
<tr>
<td>$D(\text{Comp} L)$</td>
<td>9.61</td>
<td>9</td>
<td>0.38</td>
<td></td>
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<tr>
<td>$D(\text{U}^{\text{adj}})$</td>
<td>16.68</td>
<td>9</td>
<td>0.05**</td>
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</tr>
<tr>
<td>All</td>
<td>95.13</td>
<td>45</td>
<td>0.00***</td>
<td></td>
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</table>

<table>
<thead>
<tr>
<th>Dependent variable: $D(\log\text{(P)})$</th>
<th>Exclude</th>
<th>$\chi^2$</th>
<th>df</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(\log\text{(Oil)})$</td>
<td>21.77</td>
<td>9</td>
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<tr>
<td>$D(\text{TBill})$</td>
<td>6.74</td>
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<tr>
<td>$D(\log\text{(Profits)})$</td>
<td>15.36</td>
<td>9</td>
<td>0.08*</td>
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<tr>
<td>$D(\text{Comp} L)$</td>
<td>15.79</td>
<td>9</td>
<td>0.07*</td>
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</tr>
<tr>
<td>$D(\text{U}^{\text{adj}})$</td>
<td>23.42</td>
<td>9</td>
<td>0.01***</td>
<td></td>
</tr>
<tr>
<td>All</td>
<td>96.94</td>
<td>45</td>
<td>0.00***</td>
<td></td>
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</table>
Dependent variable: $D(P_{profits})$

<table>
<thead>
<tr>
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<th>df</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(log(Oil))$</td>
<td>14.07</td>
<td>9</td>
<td>0.12</td>
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<tr>
<td>$D(TBill)$</td>
<td>19.02</td>
<td>9</td>
<td>0.03**</td>
</tr>
<tr>
<td>$D(log(P))$</td>
<td>6.87</td>
<td>9</td>
<td>0.65</td>
</tr>
<tr>
<td>$D(Comp_f)$</td>
<td>22.29</td>
<td>9</td>
<td>0.01***</td>
</tr>
<tr>
<td>$D(U_{adj})$</td>
<td>39.27</td>
<td>9</td>
<td>0.00***</td>
</tr>
<tr>
<td>All</td>
<td>131.71</td>
<td>45</td>
<td>0.00***</td>
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</table>

Dependent variable: $D(Comp_f)$

<table>
<thead>
<tr>
<th>Exclude</th>
<th>$\chi^2$</th>
<th>df</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(log(Oil))$</td>
<td>6.54</td>
<td>9</td>
<td>0.69</td>
</tr>
<tr>
<td>$D(TBill)$</td>
<td>10.99</td>
<td>9</td>
<td>0.28</td>
</tr>
<tr>
<td>$D(log(P))$</td>
<td>4.09</td>
<td>9</td>
<td>0.91</td>
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<tr>
<td>$D(P_{profits})$</td>
<td>11.77</td>
<td>9</td>
<td>0.23</td>
</tr>
<tr>
<td>$D(U_{adj})$</td>
<td>25.54</td>
<td>9</td>
<td>0.00***</td>
</tr>
<tr>
<td>All</td>
<td>82.96</td>
<td>45</td>
<td>0.00***</td>
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</tbody>
</table>

Dependent variable: $D(U_{adj})$

<table>
<thead>
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<th>Exclude</th>
<th>$\chi^2$</th>
<th>df</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(log(Oil))$</td>
<td>9.80</td>
<td>9</td>
<td>0.37</td>
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<tr>
<td>$D(TBill)$</td>
<td>22.78</td>
<td>9</td>
<td>0.01***</td>
</tr>
<tr>
<td>$D(log(P))$</td>
<td>10.41</td>
<td>9</td>
<td>0.32</td>
</tr>
<tr>
<td>$D(P_{profits})$</td>
<td>22.07</td>
<td>9</td>
<td>0.01***</td>
</tr>
<tr>
<td>$D(Comp_f)$</td>
<td>20.34</td>
<td>9</td>
<td>0.02**</td>
</tr>
<tr>
<td>All</td>
<td>88.40</td>
<td>45</td>
<td>0.00***</td>
</tr>
</tbody>
</table>

The asterisks denote the rejection of the null hypothesis of no Granger-causality at the 10% (*), 5% (**) and 1% (***) levels.

These statistics confirm that, when the model takes wealth distribution into account, the energy price no longer serves to forecast the unemployment rate. On the other hand, it would appear that the distribution of value-added as well as the Treasury bill rate yields information about the unemployment variations.

We interpret the acceptance of the null hypothesis of no Granger-causality running from oil price fluctuations to unit profits variations with a probability equal to 0.12 as the sign of an unstable influence. We estimated the same model with a time trend to catch a potential weakening of the relationships between oil price variations and the other variables. The results were not affected by this modification so we do not think that this inconstancy results mainly from a long-run structural evolution. The response of unit profits to oil price innovations varies over the business cycle and consequently the standard deviations of the coefficients associated with oil price variations are too large, leading to non significant coefficients. On the other hand, oil price variations Granger-cause significantly inflation (as well as unemployment variations and, marginally, the distribution of value-added). This is consistent with Hooker’s results (1999): he finds that the direct impact of oil price changes on U.S. all-items CPI inflation is stable over a sample included in ours. So oil price affects unemployment through its varying impact on unit profits and through a monetary channel, acting first on the general price level, which induces a monetary response and a change in interest rates.

Job destruction seems to be the best way for businesses to modify the distribution of value-added to their advantage, thus reducing labor costs and raising their profits. The consequences of hiring are just the opposite.

We performed impulse analysis using the Generalized Impulses method (Pesaran and Shin, 1998) proposed by Eviews so that the results do not depend on the VAR ordering. These simulations show that a great uncertainty is associated with the response of the system to an oil price shock (Figures 16 to 20). It does not propagate in a systematic manner and so its impact is inconsistent.
In particular, the very broad confidence intervals associated with unit profits and employee compensation responses show that although, on average, an increase in the price of petroleum products tends to exert downward pressure on these two variables (for about 6 quarters for the former and 2 years for the latter), other evolutions are possible (and have already occurred). Consequently, the indirect effect on unemployment is also highly variable. Unemployment tends to increase, principally between the fourth and the sixth quarters after the shock: the mathematical expectation of the impact reaches its peak the fifth quarter at a value of 0.05 but the 95% confidence interval lies between -0.01 and 0.10.

Inflation accelerates (the response is statistically significant the third quarter after the shock).

However, various trajectories are possible depending on the will and ability of businesses to restore margins and whether they try to do it by raising selling prices or cutting wage costs. Some of these trajectories are characterized by a decrease in unemployment or the beginning of a period of disinflation.

On the other hand, an innovation to the profits (Figures 21 to 25) clearly impacts the unemployment rate, which shows a statistically significant downtrend for one year. The impact on prices is not as clear and depends presumably on a trade-off between unemployment and inflation. On average, a rise in corporate profits exerts upward pressure on employee compensation for two years (significant the fifth and sixth quarters after the shock), which can result in a rise in unemployment or be passed on to selling prices the following year.
Figures 21 to 25: Responses (from left to right and top to bottom) of $D(T\text{Bill})$, $D(\log(P))$, $D(\pi_{\text{final}})$, $D(\text{Comp})$, $D(U_{\text{adj}})$ to a Generalized One Standard Deviation innovation on $D(\pi_{\text{final}})$ (x-axis: quarters since shock; dotted lines: 95% confidence interval).

Similarly, an innovation to labor costs (Figures 26 to 30) leads to a (non significant) reduction in the share of profits for one year. Profit margins are restored by means of a rise in unemployment (significant for one year). Since prices also tend to escalate (significantly the second, third and fourth quarters after the shock), inflation eats into real salaries and wages. The share of employee compensation per unit of value-added falls off after about four quarters (significantly the fifth and the sixth). Then employment stabilizes, with later an easing of the tension affecting prices (at the end of the third year after the shock).
Regarding the influence of the oil price on unemployment, this compact model corroborates the idea that energy price variations influence unemployment indirectly, notably through the distribution of value-added and a monetary channel. Various strategies have been implemented over time in response to oil price fluctuations; dominant theories change and the knowledge acquired thus far makes it possible to eliminate certain inappropriate responses. This might help explain the progressive weakening in the relationship between the crude price and economic activity substantially documented by Hooker. But this influence is also unstable in the short-run, over the business cycle, because economic realities dictate which alternatives are feasible.

**SUMMARY AND CONCLUSIONS**

This article focuses on the short-term instability of the relationship between oil price variations and the unemployment rate. Starting with theoretical analysis of the mechanisms whereby a variation in the oil price is propagated, thereby modifying macroeconomic conditions, we would like to emphasize the importance of the choices made concerning profit, wage and price trajectories. These variables account for a large part of high-frequency unemployment variations. However, economic conditions do not offer the same capabilities of response over the business cycle, which is the source of the short-term instability in this relationship shown statistically by Raymond and Rich (1997). At times when economic activity is accelerating, a higher barrel price leads to a reduction in value-added growth that can be equitably distributed between the various agents. Inversely, in deceleration phases, it can lead to a reduction in value-added, which represents a problem of a very different nature, with much more adverse effects.

As for modeling, this article puts forward a new formula for the spring force exerted on unemployment rate out-of-equilibrium. Rather than consider the value of the error correction term in \((t-1)\), we think it necessary to consider the value of the accumulated disequilibrium in \((t-1)\). In this way, we account for the amplitude of the disequilibrium and its duration. This is tantamount to applying an autoregressive model whose number of lags varies continually to the unemployment gap. This formula allows us to represent a significant proportion of the non-linear influences exerted on the unemployment gap.
We tried to take into account President Nixon’s Economic Stabilization Program (ESP) in our estimation. The imposition of price control in the U.S. in August 1971 led to an increasing overstatement of real GDP (and understatement of the deflator) through the first quarter of 1973. This policy was then relaxed in phases through the third quarter of 1974 with progressive elimination of the biases (see Darby, 1982, especially p. 741-744). We modeled this influence with the PC dummy variable which Darby uses (1982, 1984). This type of variable is also used by Gordon in all of his studies since 1982 and by Brayton, Roberts ans Williams (1999). It starts at 0 in 1971:2, rises smoothly to 1 in 1973:1, and then falls smoothly to 0 again in 1974:4. The variations of this variable (multiplied by an unknown coefficient) were added to the inflation rates in the unconstrained models before the step-by-step elimination procedure (for $i = 0, \ldots, 4$, $D(\log(P^*)_{h_i}) = D(\log(P))_{h_i} + \alpha D(PC)_{h_i}$). But the coefficient $\alpha$ was not significant. We tried the dummy variable used by Brayton, Roberts and Williams, which differs from Darby’s on the starting date of bias elimination, but it was also not significant. The oversimplification of these variables may be responsible for the large standard deviations associated with their coefficients, but we prefer to bias the test against accepting the price control hypothesis so as to be as objective as possible.

2 We checked for the sensibility of our results to the choice of this index by using other energy price measurements (same index not corrected for inflation, nominal and real refiner acquisition costs). These modifications do not affect the model’s estimation and the conclusions remain the same.

3 The variations of the unit profits, employee compensation and nonlabor costs are related thus: $D(P_{\text{profits}}) + D(\text{Comp}_{L}) + D(\text{Coût}_{NL}) = 0$. As a result, we retained only two of these variables.

REFERENCES


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JONES Donald W., BJORNSTAD David J. and LEIBY Paul N., The findings of the DOE workshop on economic vulnerability to oil price shocks: Summary and integration with the previous knowledge, Oak Ridge National Laboratory, xerox, 1997.


DATA APPENDIX

The data used in this study, the source and any transformations are as follows. Data is available on request. The samples specified were the samples available at the time we downloaded the series.

Unemployment rates: Civilian aged 16 and over unemployment rate from the U.S. Bureau of Labor Statistics (ID: LNS14000000), aggregated from monthly to quarterly using average value, unemployment rate for civilian males and females aged 16-19 from BLS (ID: LNS14000012Q), unemployment rate for civilian males aged 20-24 from BLS (ID: LNS14000037Q), unemployment rate for civilian females aged 20-24 from BLS (ID: LNS14000038Q), unemployment rate for civilian males aged 25 and over from BLS (ID: LNS14000049Q) and unemployment rate for civilian females aged 25 and over from BLS (ID: LNS14000050Q); all series were seasonally adjusted. (Sample: 1948:1-2004:2)


GDP Deflator: Implicit Price Deflator for Gross Domestic Product, seasonally adjusted, as reported in the Table 1.1.9 from the series of NIPA Tables (National Income and Product Accounts) compiled by the U.S. Bureau of Economic Analysis. (Sample: 1947:1-2004:2)

Price, Costs, and Profit Per Unit of Real Gross Value Added of Non-financial Domestic Corporate Business: Table 1.15 from the series of NIPA Tables compiled by the BEA. (Sample: 1947:1-2004:1)

PPI for Petroleum Products Refined: from the BLS (ID: WPU057), not seasonally adjusted, converted from monthly to quarterly using average value. (Sample: 1947:1-2003:4)

PPI for Crude Oil: from the BLS (ID: WPU0561), not seasonally adjusted, converted from monthly to quarterly using average value. (Sample: 1947:1-2003:4)

Mork Oil Price (also called the Refiner Acquisition Cost, see Mork, 1989): through 1972, 400 times the log first difference of the Crude Oil PPI. From 1974, the same transformation applied to the DOE (Department of Energy) composite domestic first purchase price (http://tonto.eia.doe.gov/merquery/mer_date.asp?table=T09.01). Intervening quarters are the growth rate of the PPI multiplied by 1.095 (see Mork, 1989). (Sample: 1947:2-2003:1)

Treasury bills rate: Yields on Treasury securities at constant, fixed maturity (5-year), not seasonally adjusted, converted from monthly to quarterly using average value (FedStats, http://www.federalreserve.gov/releases/h15/data/m/tcm5y.txt), deflated with the annualized rate of the GDP deflator. (Sample: 1954:1-2003:1)
Figure 31: Probability associated with the F-statistic of the Chow test (x-axis: proposed date for a structural break in relation [2], 5% threshold in dotted line).

Table 2: BDS Test for the residuals of the short term equation [2]

<table>
<thead>
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<th>Sample: 1954:4 2003:3</th>
</tr>
</thead>
<tbody>
<tr>
<td>Included observations: 196</td>
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<table>
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<td>0.3804</td>
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| Raw epsilon | 0.3026039 |
| Pairs within epsilon | 27034.000 | V-statistic | 0.7037172 |
| Triples within epsilon | 3985562.0 | V-statistic | 0.5293237 |