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Oil prices and economic activity: An asymmetric cointegration approach

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Abstract

The aim of this paper is to study the long-term relationship between oil prices and economic activity, proxied by GDP. To account for asymmetries existing in the links between the two variables, we propose an approach based on asymmetric cointegration. Our empirical analysis concerns the U.S. economy, but also the G7, Europe and Euro area economies. Results indicate that, while standard cointegration is rejected, there is evidence for asymmetric cointegration between oil prices and GDP.

JEL classification: C22; E32; Q43

Keywords: Oil prices; GDP; Asymmetric cointegration

1. Introduction

The aim of this paper is to study the long-term relationship between oil prices and economic activity. Unlike most of the existing literature, which focuses on the U.S. economy, we also analyze the effects of oil prices on GDP in the G7, Europe and Euro area countries.[1] In order to apprehend this relationship, various papers have considered the usual cointegration framework (for a survey,

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[1] The G7 includes the following countries: Canada, France, Germany, Italy, Japan, the United Kingdom and the United States. The European countries considered in the paper are: Austria, Belgium, Finland, France, Germany, Ireland, Italy, The Netherlands, Norway, Portugal, Spain, Sweden and the United Kingdom. Finally, the Euro area includes the following countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, The Netherlands, Portugal and Spain.

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see Jones and Leiby, 1996; Brown and Yücel, 2002). These studies generally put forward an inverse relationship between the two variables (Hamilton, 1983; Burbidge and Harrison, 1984; Gisser and Goodwin, 1986). However, by the mid-1980s, the estimated linear relationship between oil prices and GDP began to lose significance: the declines in oil prices occurred over the second half of the 1980s were found to have smaller positive effects on economic activity than predicted by usual linear models. At the same time, evidence of asymmetries in the link between the two variables has been established in some papers (see e.g. Mork, 1989; Mory, 1993; Olsen and Mysen, 1994; Ferderer, 1996; Brown and Yücel, 2002). More specifically, it has been evidenced that economic activity responds asymmetrically to oil price shocks. Indeed, rising oil prices appear to retard aggregate economic activity by more than falling oil prices stimulate it. In order to account for these characteristics, we propose to use the asymmetric cointegration framework, developed by Balke and Fomby (1997), Enders and Dibooglu (2001), Enders and Siklos (2001) and Schorderet (2004) among others. As in Schorderet (2004), we distinguish positive and negative increments of time series allowing us to break down a series into its initial value and its negative and positive cumulative sums. Asymmetric cointegration comes from the analysis of multivariate combinations arising from this decomposition.

The paper is organized as follows. Section 2 recalls the main transmission channels through which oil prices may have an impact on GDP. Section 3 briefly presents the econometric framework. Section 4 is devoted to an empirical analysis of the link between oil prices and GDP in the long run in the G7, U.S., Europe and Euro area countries. Section 5 concludes the paper.

2. Oil prices and economic activity: transmission channels

Oil prices may have an impact on economic activity through various transmission channels. First, there is the classic supply-side effect according to which rising oil prices are indicative of the reduced availability of a basic input to production, leading to a reduction of potential output (see, among others, Barro, 1984; Brown and Yücel, 1999; Abel and Bernanke, 2001). Consequently, there is a rise in cost production, and the growth of output and productivity are slowed. Second, an increase of oil prices deteriorates terms of trade for oil-importing countries (see Dohner, 1981). Thus, there is a wealth transfer from oil-importing countries to oil-exporting ones, leading to a fall of the purchasing power of firms and households in oil-importing countries. Third, according to the real balance effect (Pierce and Enzler, 1974; Mork, 1994), an increase in oil prices would lead to increase money demand. Due to the failure of monetary authorities to meet growing money demand with increased supply, there is a rise of interest rates and a retard in economic growth (for a detailed discussion on the impact of monetary policy, see Brown and Yücel, 2002). Fourth, a rise in oil prices generates inflation. The latter can be accompanied by indirect effects, called second round effects, given rise to price–wages loops. Fifth, an oil price increase may have a negative effect on consumption, investment and stock prices. Consumption is affected through its positive relation with disposable income, and investment by increasing firms' costs. Sixth, if the oil price increase is long-lasting, it can give rise to a change in the production structure and have an impact on unemployment. Indeed, a rise in oil prices diminishes the rentability of sectors that are oil-intensive and can incite firms to adopt and construct new production methods that are less intensive in oil inputs. This change generates capital and labor reallocations across sectors that can affect unemployment in the long run (Loungani, 1986). For all these reasons, oil prices can affect economic activity.

From an empirical viewpoint, studies generally consider a linear cointegration framework to analyze the link between oil prices and GDP (see Brown and Yücel, 2002 for a recent survey). However, as was previously mentioned, economic activity tends to react asymmetrically to oil price shocks: an increase in oil prices seems to retard aggregate economic activity by more than a fall in oil prices stimulates it. In order to apprehend this fact, it seems necessary to go further than the usual cointegration concept which is rather restrictive. Let us now describe a general framework — asymmetric cointegration — devoted to account for these asymmetries.

3. Definitions and general methodology

Following Schorderet (2004), the starting point consists in the decomposition of a time series X_t into its positive (X_t^+) and negative (X_t^-) partial sums:

$$X_t^+ = \sum_{i=0}^{t-1} 1\{\Delta X_{t-i} \geq 0\} \Delta X_{t-i} \text{ and } X_t^- = \sum_{i=0}^{t-1} 1\{\Delta X_{t-i} < 0\} \Delta X_{t-i} \tag{1}$$

Consider two integrated time series X_{1t} and X_{2t} and define X_{jt}^+ and X_{jt}^- for $j=1, 2$, according to Eq. (1). Suppose that there exists a linear combination z_t between X_{jt}^+ and X_{jt}^- such that:

$$z_t = \beta_0 X_{1t}^+ + \beta_1 X_{1t}^- + \beta_2 X_{2t}^+ + \beta_3 X_{2t}^- \tag{2}$$

Then, as stated by Schorderet (2004), X_{1t} and X_{2t} are asymmetrically — or directionally — cointegrated if there exists a vector $\beta' = (\beta_0, \beta_1, \beta_2, \beta_3)$ with $\beta_0 \neq \beta_1$ or $\beta_2 \neq \beta_3$ (and β_0 or $\beta_1 \neq 0$ and β_2 or $\beta_3 \neq 0$) such that z_t in Eq. (2) is a stationary process. The idea is that the relationship between the variables might not be the same whenever they increase or decrease. To simplify and without loss of generality, suppose that only one component of each series appears in the cointegrating relationship (2), that is[2]:

$$z_{1t} = X_{1t}^+ - \beta^+ X_{2t}^+ \text{ or } z_{2t} = X_{1t}^- - \beta^- X_{2t}^- \tag{3}$$

Due to the nonlinear properties of z_{jt}, $j=1, 2$, OLS on Eq. (3) are likely to be biased in finite sample. For this reason, Schorderet (2004) suggests to estimate by OLS the auxiliary models[3]:

$$\varepsilon_{1t} = X_{1t}^- + \Delta X_{1t}^+ - \beta^- X_{2t}^- \text{ or } \varepsilon_{2t} = X_{1t}^+ + \Delta X_{1t}^- - \beta^+ X_{2t}^+ \tag{4}$$

As proved by West (1988), since the regressor has a linear time trend in mean, the OLS estimate of Eq. (4) is asymptotically normal and usual statistical inference can be done. In order to test the null hypothesis of no cointegration against the alternative of asymmetric cointegration, the traditional Engle and Granger procedure can be applied to Eq. (4).

4. Empirical results

Our aim is to study the long-term relationship between oil prices and GDP not only in the U.S., but also in G7, Europe and Euro area countries. Oil prices and GDP series are naturally expressed

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[2] See Granger and Yoon (2002). Note that various extensions of these models are possible. For example, an intercept can be added.

[3] These auxiliary regressions can be obtained by defining z_{jt} ($j=1, 2$) as the outcome of some underlying disturbance denoted ε_{jt} ($j=1, 2$). Defining: $z_{1t} = \begin{cases} \max \quad [-\beta^+ \Delta X_{2t}^+; \varepsilon_{1t}] \text{ for } t=1 \\ \max \quad [X_{1t-1}^+ - \beta^+ X_{2t}^+; \varepsilon_{1t}] \text{ for } t=2, \ldots, T \end{cases}$ and letting $\Delta X_{1t}^- = \begin{cases} \min \quad [0; \beta^+ \Delta X_{2t}^+ + \varepsilon_{1t}] \text{ for } t=1 \\ \min \quad [0; \beta^+ X_{2t}^+ - X_{1t-1}^+ + \varepsilon_{1t}] \text{ for } t=2, \ldots, T \end{cases}$ then, under some conditions, this last equation can be rewritten as $\Delta X_{1t}^- = \varepsilon_{1t} - z_{1t}$. Combining the latter equation with Eq. (3), one obtains the auxiliary model.

Table 1
Unit root tests on individual series

	Series in logarithms		Series in first differences	
	ADF	PP	ADF	PP
LOIL	−2.7052 (2)	0.5419 (1)	−8.8703** (1)	−8.7975** (1)
LGDP G7	−2.8579 (3)	9.3293 (1)	−7.3504** (2)	−3.9069** (1)
LGDP U.S.	−3.6674* (3)	7.9608 (1)	−8.5653** (2)	−6.5410** (1)
LGDP Europe	−2.9368 (3)	−3.3049 (3)	−8.5653** (2)	−8.7986** (2)
LGDP Euro area	5.8736 (1)	−2.2557 (2)	−3.7314** (1)	−5.8436** (1)

ADF: Augmented Dickey–Fuller test. PP: Phillips–Perron test. (1): Model without constant, or deterministic trend. (2): Model with constant, without deterministic trend. (3): Model with constant and deterministic trend. * (resp. **): Rejection of the null hypothesis at the 5% (resp. 1%) significance level.

in real terms, *i.e.* they have been deflated by consumer price indexes.[4] We use quarterly data from 1970:1 to 2004:3.

4.1. Unit root and standard cointegration tests

We apply standard unit root tests on each individual series (in logarithms). Results are given in Table 1. All series are I(1) at the 1% significance level.

We now proceed to standard cointegration tests between the two considered series, by regressing GDP on oil prices and an intercept. Results are displayed in Table 2. We consider three tests of the null hypothesis of no cointegration: Augmented Dickey–Fuller (ADF), Phillips–Perron (PP) and the trace statistic developed by Johansen (1988) and Johansen and Juselius (1990). On the contrary, the Shin (1994) test — which is an extension of the Kwiatkowski et al. (1992) test to cointegration — is based on the null of cointegration.

In order to perform the ADF test, we have to specify the number of lagged differenced terms of the dependent variable to be added to the test regression. We include here a number of lags sufficient to remove dependence in the residuals. We retain the number of lags that minimize information criteria. Concerning the application of the PP test, one has to choose the bandwidth parameter needed for estimating the residual spectrum at frequency zero. To this end, we use the Newey–West (1994) data-based automatic bandwidth parameter method. The same procedure is retained for the application of the Shin test. In all cases, the aim of these choices concerning the truncation parameters (the number of lags for ADF and the bandwidth parameter for PP and Shin) is to remove serial correlation in the residuals.

Results in Table 2 are somewhat mitigated. Indeed, according to ADF, PP and Shin tests all residual series are nonstationary, meaning that oil prices and GDP series are not cointegrated. Results of the Johansen test indicate that the two considered series are cointegrated at the 5% significance level in Europe and in the Euro area. It thus seems difficult to conclude. This problem may be due do the fact that the usual concept of cointegration is too restrictive. Indeed, as was previously mentioned, there exists a strong evidence in favor of an asymmetric relationship between oil prices and GDP series. To investigate this possibility, it is necessary to go further than the usual concept of cointegration in order to allow for asymmetric cointegration.

[4] The oil price series is the U.S. dollar per barrel market price of crude oil taken from the IMF database.

Table 2
Unit root tests on residual series

	ADF	PP	Shin	Johansen
G7	−0.6048 [1]	−1.0609	1.4698**	11.3297
U.S.	−0.4357 [1]	−0.6284	1.4886**	7.7305
Europe	−0.4168 [2]	−1.2325	1.4569**	16.0511*
Euro area	−0.4863 [2]	−1.3025	1.4897**	15.6718*

ADF: Augmented Dickey–Fuller test. PP: Phillips–Perron test. Johansen: trace statistic. The number of lags in the ADF regressions is given into brackets. * (resp. **): Rejection of the null hypothesis at the 5% (resp. 1%) significance level.

4.2. Tests for asymmetric cointegration between oil prices and GDP series

In order to test for asymmetric cointegration between oil prices and GDP, we estimate the two following auxiliary models according to Eq. (4):

$$\mathrm{LGDP}_t^- + \Delta\mathrm{LGDP}_t^+ = \alpha^- + \beta^-\mathrm{LOIL}_t^- + \varepsilon_{1t} \tag{5}$$

$$\mathrm{LGDP}_t^+ + \Delta\mathrm{LGDP}_t^- = \alpha^+ + \beta^+\mathrm{LOIL}_t^+ + \varepsilon_{2t} \tag{6}$$

We then test if the residuals ε_{1t} and ε_{2t} are or not stationary. As mentioned by Schorderet (2004), standard unit root and cointegration tests can be applied to the residuals of these auxiliary models. Consequently, we apply the same tests as for the standard cointegration case: ADF, PP, Shin and Johansen tests.

As before, we choose the truncation parameters — the number of lags for ADF and the bandwidth parameter for PP and Shin — to remove dependence in the residuals. Thus, the residuals of Eqs. (5) and (6) are not autocorrelated. Before interpreting the results, it should be noted that, since we work on residuals — *i.e.* on series that are estimated and not observed — we have to use other critical values than those tabulated when tests are applied on "true" series. Performing a Monte Carlo experiment, Schorderet (2004) showed that critical values for the ADF test are essentially the same as those given by Fuller (1976) when serial correlation has been correctly removed. Thus, for the PP test, we use the same critical values as for the ADF test. Concerning the Shin and Johansen tests, we use the critical values respectively tabulated by Shin (1994) and Johansen and Juselius (1990).

Results are displayed in Tables 3 and 4. Considering first tests on ε_{1t}, it appears that for Europe and Euro area countries, oil prices and GDP appear to be asymmetrically cointegrated according

Table 3
Unit root tests on residual series: tests for asymmetric cointegration

	Tests on ε_{1t}			
	ADF	PP	Shin	Johansen
G7	−2.4527 [8]	−2.3186	0.1687	12.3012
U.S.	−2.0627 [11]	−2.1175	0.1613	12.3197
Europe	−3.3164* [1]	−4.2109***	0.2117	9.7710
Euro area	−2.8677 [1]	−4.4729***	0.2022	10.4515

ADF: Augmented Dickey–Fuller test. PP: Phillips–Perron test. Johansen: trace statistic. The number of lags in the ADF regressions is given into brackets. * (resp. **, ***): Rejection of the null hypothesis at the 10% (resp. 5%, 1%) significance level.

Table 4
Unit root tests on residual series: tests for asymmetric cointegration

	Tests on ε_{2t}			
	ADF	PP	Shin	Johansen
G7	−3.0092 [1]	−2.6828	0.1462	21.2999**
U.S.	−2.9284 [1]	−2.6711	0.1528	16.4379**
Europe	−3.2186* [1]	−2.8586	0.1466	19.9343**
Euro area	−3.1219* [1]	−2.7929	0.1459	15.7197**

ADF: Augmented Dickey–Fuller test. PP: Phillips–Perron test. Johansen: trace statistic. The number of lags in the ADF regressions is given into brackets. * (resp. **, ***): Rejection of the null hypothesis at the 10% (resp. 5%, 1%) significance level.

to the unit root tests (except for ADF in the Euro area case). On the contrary, the Johansen test indicates no evidence of cointegration between the two variables. Tests on ε_{2t} indicate that oil prices and GDP in Europe and Euro area are asymmetrically cointegrated according to all tests, but PP. For U.S. and G7 countries, and according to ADF and PP, asymmetric cointegration is rejected. This hypothesis is however not rejected using the Shin test (and the Johansen test on ε_{2t}). It is noteworthy that ADF and PP approaches allow us to test the existence of a mean reversion towards a long-term target. The rejection of the unit root hypothesis would imply a constant rate of adjustment, the deviation observed at a given time being proportional to the deviation observed at the preceding period. As we have seen, such a hypothesis seems difficult to support due to asymmetry phenomena. The Shin test is a non-parametric one: it allows to test non-parametrically the degree of persistence of the deviations. The Shin test is thus less restrictive and more general than ADF and PP tests, since it is not based on a linear parametrization. Using this argument, our results indicate evidence of asymmetric cointegration, specially in the form (6) meaning that the two series are cointegrated when there is a rise in the oil prices.

Tables 5 and 6 report the long-term relationship estimates. Note that we have calculated consistent estimates of the standard errors of coefficients. Following West (1988), these estimates are obtained by scaling the OLS statistics by $\sqrt{\hat{s}/\hat{\sigma}_{\varepsilon}^{2}}$, where $\hat{\sigma}_{\varepsilon}^{2}$ denotes the usual OLS standard error of regression and $\hat{s}$ is an estimate of the spectral density at frequency zero of the regression disturbance. The results indicate that regressions are generally better for Eq. (6), especially for U.S. and G7 countries.

It should be noted that these relationships are expressed in terms of partial sums, not in terms of the variables themselves. Consequently, the coefficients β^{+} and β^{-} cannot be interpreted in the usual way. What is important to note is that the coefficients are significantly different, meaning that an asymmetry phenomenon exists. Indeed, the fact that β^{+} is generally higher than β^{-} means

Table 5
Long-term relationships

	Relation (5)			
	α^{-}	β^{-}	Adj. R^2	Std. error
G7	−0.02 (−5.41)	0.01 (8.84)	0.61	0.01
U.S.	−0.04 (−8.00)	0.02 (10.77)	0.70	0.02
Europe	−0.01 (−5.79)	0.01 (15.37)	0.80	0.01
Euro area	−0.08 (−5.50)	0.28 (18.56)	0.86	0.01

Adj. R^2 is the adjusted R-squared. Std. error is the residual standard error resulting from OLS estimation of the two equations. t-statistics have been corrected as in West (1988) and are displayed in parentheses.

Table 6
Long-term relationships

	Relation (6)			
	α^+	β^+	Adj. R^2	Std. error
G7	0.01 (0.36)	0.16 (37.15)	0.97	0.05
U.S.	−0.03 (−1.31)	0.18 (35.97)	0.96	0.06
Europe	0.03 (2.28)	0.13 (44.44)	0.98	0.04
Euro area	0.04 (3.05)	0.13 (45.16)	0.98	0.04

Adj. R^2 is the adjusted R-squared. Std. error is the residual standard error resulting from OLS estimation of the two equations. t-statistics have been corrected as in West (1988) and are displayed in parentheses.

that an oil price increase has more impact on GDP than an oil price decrease.[5] Thus, our results put forward that GDP asymmetrically reacts to the evolution of the oil prices.

We have shown that economic activity reacts asymmetrically to oil price shocks. What are the possible explanations for this phenomenon? As argued by Brown and Yücel (2002), classical supply-side effects cannot explain asymmetry. Possible explanations are: monetary policy, adjustment costs, adverse effects of uncertainty on the investment environment (Ferderer, 1996), and asymmetry in petroleum product prices (especially gasoline). Consider first monetary policy. Assume that prices are nominally sticky downward. Then, an increase in oil prices leads to important GDP losses if the monetary authorities do not maintain nominal GDP constant through unexpected inflation. On the contrary, after a decline in oil prices, real wages must grow to clear the markets. Thus, monetary policy can have asymmetric effects (Bernanke et al., 1997). According to the cost adjustment explanation (Hamilton, 1988), the costs induced by changing oil prices can retard economic activity. Such costs could arise from sectoral imbalances[6] (Lilien, 1982; Hamilton, 1988), coordination failures between firms (Huntington, 2000), or because the energy-to-output ratio is embedded in the capital stock (Atkeson and Kehoe, 1999). Finally, various studies (e.g. Bacon, 1991; Balke et al., 1998) evidenced that petroleum product prices respond asymmetrically to crude oil prices: gasoline prices increase more quickly when crude oil prices are rising than they fall when crude oil prices are decreasing. Thus, there exist various reasons for explaining the evidenced asymmetric cointegration relationship between oil prices and GDP.

5. Conclusion

The majority of empirical work that addresses whether or not a long-term relationship exists between oil prices and GDP is based on the usual linear cointegration framework. However, it is well known that asymmetries exist in the links between the two variables. Indeed, rising oil prices seem to retard aggregate economic activity by more than falling oil prices stimulate it. To account for these characteristics we have proceed to an asymmetric cointegration analysis. Our empirical study indicates that such a behavior is relevant since asymmetric cointegration can be found, while standard cointegration was rejected. Moreover, this phenomenon is not only present in the U.S., but also exists in the other considered countries: the G7, Europe and Euro area countries.

[5] This can be clearly seen by regressing LGDP on $LOIL^+$ and $LOIL^-$: the estimated coefficient on $LOIL^+$ is generally twice as high as that on $LOIL^-$, especially for Europe and Euro area.

[6] This is the so-called dispersion hypothesis of Lilien (1982) according to which an increase in oil prices would lead to a contraction in sectors that make use of oil in the production process.

References

Abel, A.B., Bernanke, B.S., 2001. Macroeconomics. Addison Wesley Longman Inc.

Atkeson, A., Kehoe, P.J., 1999. Models of energy use: putty–putty vs. putty–clay. American Economic Review 89, 1028–1043.

Bacon, R.W., 1991. Rockets and feathers: the asymmetric speed of adjustment of U.K. retail gasoline prices to cost changes. Energy Economics 13, 211–218.

Balke, N.S., Fomby, T.B., 1997. Threshold cointegration. International Economic Review 38, 627–645.

Balke, N.S., Brown, P.A., Yücel, M.K., 1998. Crude oil and gasoline prices: an asymmetric relationship? Federal Reserve Bank of Dallas Economic Review 2–11.

Barro, R.J., 1984. Macroeconomics. John Wiley and Sons.

Bernanke, B.S., Gertler, M., Watson, M., 1997. Systematic monetary policy and the effects of oil price shocks. Brookings Papers on Economic Activity 91–157.

Brown, S.P.A., Yücel, M.K., 1999. Oil prices and U.S. aggregate economic activity. Federal Reserve Bank of Dallas Economic Review 16–53.

Brown, S.P.A., Yücel, M.K., 2002. Energy prices and aggregate economic activity: an interpretative survey. Quarterly Review of Economics and Finance 42, 193–208.

Burbidge, J., Harrison, A., 1984. Testing for the effects of oil price rises using vector autoregression. International Economic Review 25, 459–484.

Dohner, R.S., 1981. Energy prices, economic activity and inflation: survey of issues and results. In: Mork, K.A. (Ed.), Energy Prices, Inflation and Economic Activity. Cambridge Mass., Ballinger.

Enders, W., Dibooglu, D., 2001. Long-run purchasing power parity with asymmetric adjustment. Southern Economic Journal 68 (2), 433–445.

Enders, W., Siklos, P.L., 2001. Cointegration and threshold adjustment. Journal of Business and Economic Statistics 19, 166–176.

Ferderer, J.P., 1996. Oil price volatility and the macroeconomy: a solution to the asymmetry puzzle. Journal of Macroeconomics 18, 1–16.

Fuller, W.A., 1976. Introduction to Statistical Time Series. Wiley, New York.

Gisser, M., Goodwin, T.H., 1986. Crude oil and the macroeconomy: tests of some popular notions. Journal of Money, Credit and Banking 18, 95–103.

Granger, C.W.J., Yoon, G., 2002. Hidden cointegration. University of California, San Diego, Department of Economics Working Paper 2002–02.

Hamilton, J.D., 1983. Oil and the macroeconomy since World War II. Journal of Political Economy 91, 228–248.

Hamilton, J.D., 1988. A neoclassical model of unemployment and the business cycle. Journal of Political Economy 96, 593–617.

Huntington, H.G., 2000. Oil Security as a Macroeconomic Externality. Stanford University.

Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 231–254.

Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inferences on cointegration with application to the demand for money. Oxford Bulletin of Economics and Statistics 52, 169–210.

Jones, D.W., Leiby, P., 1996. The Macroeconomic Impacts of Oil Price Shocks: A Review of the Literature and Issues. Oak Ridge National Laboratory. xerox.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P., Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root. How sure are we that economic time series have a unit root? Journal of Econometrics 54, 159–178.

Lilien, D., 1982. Sectoral shifts and cyclical unemployment. Journal of Political Economy 90, 777–793.

Loungani, P., 1986. Oil price shocks and the dispersion hypothesis. Review of Economics and Statistics 58, 536–539.

Mork, K.A., 1989. Oil and the macroeconomy when prices go up and down: an extension of Hamilton's results. Journal of Political Economy 97, 740–744.

Mork, K.A., 1994. Business cycles and the oil market. The Energy Journal 15, 15–38.

Mory, J.F., 1993. Oil prices and economic activity: is the relationship symmetric? The Energy Journal 14, 151–161.

Newey, W., West, K., 1994. Automatic lag selection in covariance matrix estimation. Review of Economic Studies 61, 631–653.

Olsen, O., Mysen, H., 1994. Macroeconomic responses to oil price increases and decreases in seven OECD countries. Energy Journal 15, 19–35.

Pierce, J.L., Enzler, J.J., 1974. The effects of external inflationary shocks. Brookings Papers on Economic Activity 1, 13–61.
Schorderet, Y., 2004. Asymmetric cointegration. Working Paper. Department of Econometrics, University of Geneva.
Shin, Y., 1994. A residual-based test of the null of cointegration against the alternative of no cointegration. Economic Theory 10, 91–115.
West, K.D., 1988. Asymptotic normality when regressors have a unit root. Econometrica 56, 1397–1417.