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Guro Børnes Ringlund, Knut Einar Rosendahl and Terje Skjerpen

Does oilrig activity react to oil price changes?

An empirical investigation

Abstract:
In this paper we analyse how oilrig activity in different Non-OPEC regions is affected by the crude oil price. Oilrig activity outside OPEC is an important indicator for production in the near future, and is more sensitive to the oil price than production from existing fields. We estimate relationships between oilrig activity and crude oil prices using Equilibrium Correction Models (ECM) augmented with a stochastic time trend. The results generally show a positive relationship between oilrig activity and the crude oil price, but the strength of the relationship differs across regions. Rig activity in the US seems to react much faster and stronger to oil price changes compared to other regions. In the long-run the price elasticity in the US is above 1.5. Half the effect is observed after six months. In other regions the long-run elasticity is mainly between 0.5 and 1. Overall, it seems to be a clear relationship between the oil industry structure in the region and the reaction to price changes.

Keywords: Oilrig Activity, Oil Prices, Equilibrium Correction Model, Stochastic trend

JEL classification: C22, Q31, Q41

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Address:

Guro Børnes Ringlund, Statistics Norway, Research Department. E-mail: gurori@hotmail.com

Knut Einar Rosendahl, Statistics Norway, Research Department. E-mail:
knut.einar.rosendahl@ssb.no

Terje Skjerpen, Statistics Norway, Research Department. E-mail: terje.skjerpen@ssb.no

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Statistics Norway
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Telephone: +47 62 88 55 00
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1 Introduction

The oil market has for decades been characterised by fluctuating prices, partly due to the role of OPEC, making it difficult for other oil producers to determine the profitability of investments in exploration and development. As a market leader, OPEC has applied different strategies over the years. Holding back production has lead to higher prices, but also to a gradual decline in oil demand and enhanced oil supply from producers outside OPEC. Since oil production in Non-OPEC countries is rather inflexible in the short-run, due to low operating costs and long investment time, this effect only shows up after some years.

Although it is an established fact that higher crude oil prices eventually bring about higher Non-OPEC supply, there is little consensus about the size and speed of this effect. For instance, Ramcharran (2002) estimated supply functions for 9 Non-OPEC countries, finding long-run price elasticities between -0.1 and 2.4. In an earlier study Griffin (1985) found somewhat higher elasticities for most countries (with a range between -0.1 and 3.4). Alhajji and Huettner (2000) focused on OPEC behaviour, but also arrived at supply elasticities of Non-OPEC as a group. In a dynamic dominant firm model, these elasticities became 0.01 in the short-run and 1 in the long-run.

The large difference between short- and long-run elasticities in the latter study reflects that most oil production worldwide comes from fields that have been developed years before, where operating costs are rather small. Thus, for most of the fields in Non-OPEC countries the production level is to a certain extent insensitive to price changes, see e.g. Thomson (2001).[1] On the other hand, the oil price obviously affects investment decisions, related to developing new fields and exploring new areas, as price expectations to a large degree seem to be adaptive (cf. Farzin, 2001). Consequently, oil price changes influence production mainly through its impact on field development, which in turn is affected in the long-run by exploration activity. Bringing a new field on stream takes up to several years, and since the lifetime of a field may be several decades, careful analyses of future market conditions are needed. That is, a sudden price hike is not sufficient to bring a large, new field into production.

For OPEC it is of vital importance to understand how Non-OPEC oil production is affected by the crude oil price. It is also essential for other market participants and analysts, who need to comprehend the functioning of the oil market. Thus, it is wise to take a closer look at the relationship between oilrig

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[1] Note that Black and LaFrance (1998) question this rigidity based on an empirical investigation of oil fields in Montana.

activity and crude oil prices in different parts of the world. Rig activity is a good indicator of the exploration effort and field development taking place, and frequent records of this activity exist for various regions and countries (see below). Alternative indicators could be investment spending or the number of wells drilled, but such data does not exist at the same detailed level (wrt. temporal distribution and geographical coverage). Therefore, in this study we investigate empirically how the world market price of crude oil affects the oilrig activity in the short- and long-run in different geographical areas. As far as we know, similar studies have not been undertaken before.[2] Iledare (1995) estimates the effects of wellhead gas prices (and other variables) on natural gas drilling effort, measured as total footage drilled, in West Virginia. He applies annual data for 1977-1987, and pools cross-sections of times series data to allow a sufficient number of observations. The estimations indicate a price elasticity of around 1 over this period. Farzin (2001) examines empirically how the oil price affects reserve additions (i.e., development and extension) of known fields in the US, using annual data from 1951-95. He obtains short- and long-run elasticities of 0.11 and 0.16. OGJ (2003a) notes that rig mobilization in the US tends to follow, but lag, oil price fluctuations, and shows a figure of movements in the West Texas Intermediate (WTI) price and rig mobilization since 1993. Moreover, Abraham (2000) claims that the oil industry has been quick to boost drilling activity whenever oil prices have stayed high for at least six months, except in the beginning of 2000.

To estimate the relationship between crude oil prices and oilrig activity, we use an Equilibrium Correction Mechanism (ECM) model, enabling us to consider both short- and long-run effects of price changes. Moreover, we introduce a stochastic time trend (Harvey et al., 1986) in order to capture unobservable factors such as technological change and resource availability (see e.g. Hunt and Ninomiya, 2003, who applies stochastic trends in energy demand estimation). In our view, this is preferable to introducing deterministic proxy variables, which may be unable to describe the complex influences of these factors.

Our estimation results clearly show a positive relationship between oilrig activity in Non-OPEC regions and crude oil prices in the long-run. However, there are large differences across geographical areas, reflecting that the oil industry structure is very diverse in different regions. The strongest relationship is undoubtedly found in the US, where the long-run price elasticity is above 1.5.

The next section dwells more upon the background of estimating econometric relations for oilrig activity, and presents the data we use. Section 3 then discusses the estimation method applied, whereas

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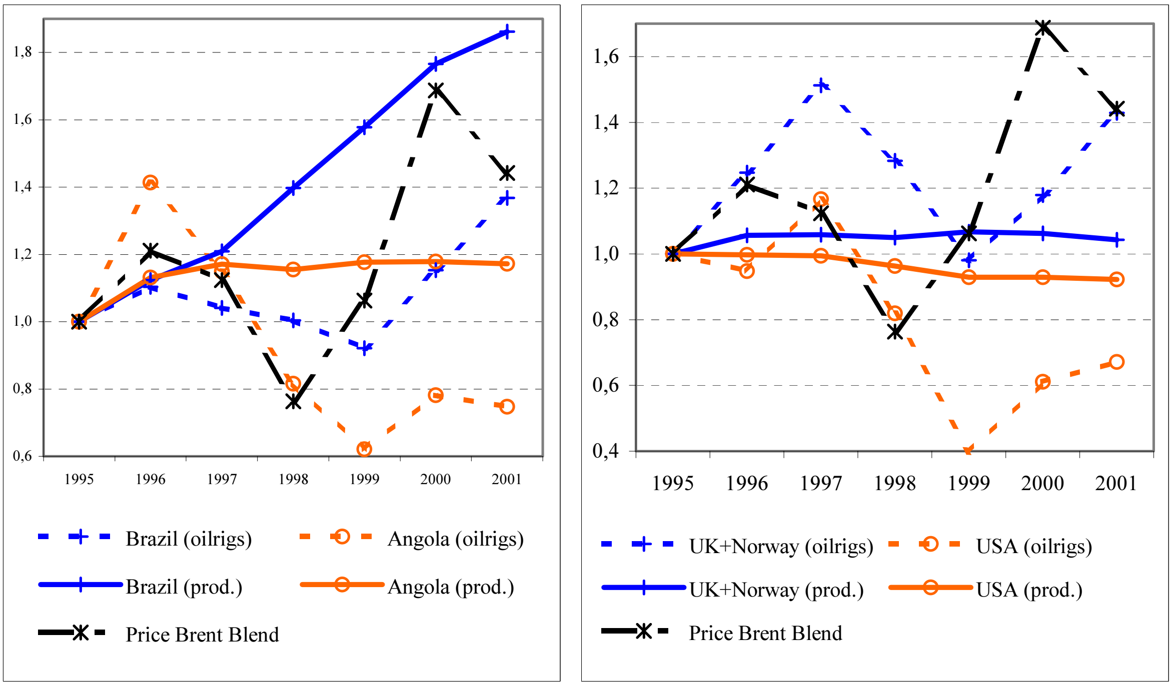
[2] Except for a simple regression analysis of annual rig activity in the US presented in a short paper by Renshaw (1989).

Section 4 presents the results. A brief summing up and conclusions are given in Section 5. Supplementary results and calculation methods are found in the appendices.

2 Background and data

Rig activity is preparation for future production of oil (or gas), either through exploration for new fields or development of existing fields. Thus, the current rig activity is an important signal for the future level of oil production. Moreover, rig activity outside OPEC is more flexible than oil production, and can therefore react quicker to changes in market conditions than actual production does. To illustrate this, take a look at Figure 1. Here we have plotted annual oil production and average oilrig activity for some important Non-OPEC countries, together with the average yearly price of Brent Blend (1995-level=1 for all variables). It is clearly seen that oil production is following a certain trend over time, with only small deviations. On the other hand, oilrig activity is much more volatile. This is most evident around 1998, when the oil price plummeted.

Figure 1. Oil production and oilrig activity in selected Non-OPEC countries, 1995-2001. 1995=1



Source: BP Statistical Review of World Energy June 2003 and Baker Hughes

Expectations about future profitability of producing oil will clearly be important for the level of rig activity in most regions of the world. For Non-OPEC countries, which are more or less price takers in the oil market, expectations about profitability depend crucially on price expectations. Furthermore, price expectations often seem to be more or less adaptive (cf. the reactions to the oil price plunge in 1998 displayed in Figure 1), although long-term expectations of the market balance is important, too. Consequently, we formulate a dynamic relationship between oilrig activity and a smoothed oil price. We investigate how these are related both in the short- and the long-run and how robust the relationship is with respect to the smoothing assumption. Another issue we address is speed of adjustment, i.e., how fast does oilrig activity change when the oil price changes. As our data are separated into several oil producing regions, focus is also on heterogeneity in oilrig responses across regions. There are lots of reasons why the response could differ, such as different terrain (e.g. offshore vs. onshore) and field size, and different regulations of the oil industry.

The relationship between oilrig activity and crude oil prices in OPEC-countries is more complicated, as these countries are able to affect the price of oil through coordinated production or quota levels. On the other hand, OPEC consists of heterogeneous countries, and some of these countries seem to behave more or less independently of the rest of the producer group (cf. e.g. Alhajji and Huettner, 2000).Thus, we will also estimate the relationship between oil price and oilrig activity for some of the OPEC countries (not for the Middle East countries).

The formal model put forward below is partly based on Farzin's (2001) study of reserve additions in the US. Iledare (1995) derives a theoretical model framework for estimating natural gas drilling in West Virginia. However, concluding that the model is too complex to be estimated, he rather puts forward a log-linear function quite similar to the one we develop below (not dynamic, however).

Let $C(Y_t^*)$ denote the total cost of oilrig activity (Y_t^*), and let $F(Y_t^*, Z_t, A_t)$ denote the number of barrels developed (or discovered) as a function of rig activity, the technological level (Z_t) and accumulated rig activity (A_t) in the area. Furthermore, let X_t^e be the expected value per barrel of the developed (discovered) resource. Then, an oil producer faces the following profit maximization problem:[3]

$$(1)\ \underset{Y_t^*}{Max}\left\{X_t^e\, F(Y_t^*, Z_t, A_t) - C(Y_t^*)\right\}.$$

This yields the following first order condition:

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[3] Note that we ignore intertemporal behaviour here, i.e., postponing the development until the present value of extraction is maximized (see Farzin (2001) for a justification of this choice).

$$(2)\, X_t^e F'_{Y^*}(Y_t^*, Z_t, A_t) = C'(Y_t^*).$$

It is reasonable to believe that C'>0 and C">0, at least at the regional level. If rig activity in a region increases, the day rates of hiring rigs tend to increase. In the short-run the rig fleet is limited, and the utilization rate seems to be an important determinant for the day rate.[4]

The relationship between the number of barrels developed (discovered) and rig activity is complex, of course. There are different sorts of rigs used in different terrains, and new sorts of rigs are developed as technology improves. We assume that $F'_{Y*}>0$ and $F''_{Y*}\leq 0$, i.e., the number of barrels developed (discovered) is an increasing but concave function of rig activity.

Technological progress may reduce the time needed to develop a certain field, but it may also make development in new areas possible.[5] Thus, the impact on the rig count is ambiguous. Both Farzin (2001) and Iledare (1995) account for accumulated reserve additions or accumulated drilling effort as a proxy for resource depletion, leading to higher costs when an area is mature. However, in our context this is more dubious as accumulated rig activity may have the opposite effect if it mainly relates to exploratory activity, which may increase the prospects of developing new fields. Consequently, both technological development and accumulated rig activity are important factors, but their impact may change over time. We therefore prefer to introduce a stochastic time trend (μ_t) to capture these effects (see below), and suggest the following specification of F() (with $\rho \leq 1$):

$$(3)\, F(Y_t^*, Z_t, A_t) = \left(Y_t^*\right)^{\rho} e^{\mu_t}.$$

Farzin (2001) introduces a partial adjustment process for his reserve additions, reflecting various constraints and an option value of waiting for higher prices. Moreover, he applies an adaptive expectations formation hypothesis for expectations of future prices.[6] His cost function is specified as a Cobb Douglas function. From these assumptions he arrives at an equation (in logarithmic form) where

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[4] ODS-Petrodata presents monthly day rates for selected regions and rig segments, see http://www.ods-petrodata.com.

[5] The most advanced deep-water drilling rigs are able to drill at water depths of up to 10,000 feet, with drilling depths up to 35,000 feet, see http://www.coltoncompany.com/shipbldg/worldsbldg/rigbldg/dwrigs.htm. See also OGJ (2003c) for a discussion of new offshore drilling techniques.

[6] Farzin (2001) argues that futures prices are highly susceptible to speculative factors, and that there is strong statistical support for adaptive price expectation formation. Moreover, Pindyck (2001) argues that the expected future spot price usually will exceed the futures price due to risk adjustment. We have tested the effects of using futures prices (8 weeks) instead of spot prices for Brent Blend in the case of Europe. The statistical results were poorer compared to our main estimation results based on spot prices.

reserve additions is a linear function of the last period price, the two last periods reserve additions and cumulative reserve additions, and a time trend. It is straightforward to arrive at a similar equation within our context, where the accumulated effect and the (deterministic) time trend are replaced by a stochastic time trend. However, wechoose to estimate a more general function, allowing for more time lags for both the oil price (X_t) and the rig activity (Y_t):[7]

$$(4) \ln Y_t = \sum_{i=1}^{p} \alpha_i \ln Y_{t-i} + \sum_{i=0}^{q} \delta_i \ln X_{t-i} + \kappa M_t + \mu_t + \varepsilon_t .$$

In Eq. 4 M_t is an m×1 vector of deterministic variables, μ_t is a stochastic trend and ε_t is an error term. The logarithmic form implies that the short-, intermediate- and long-run price elasticities are functions of the α_i and the δ_i parameters, and do not depend on the value of the variables at a given point of time.

As indicated by the M_t vector, other factors than prices may also affect rig activity. There may be seasonal variations,[8] due to weather conditions or companies' spending patterns - high activity at the end of a year to fulfil commitments, and correspondingly low activity at the start of a year before new contracts are made. Furthermore, there are a limited number of available rigs, and these are used both for oil and natural gas. Thus, if many of the rigs are being utilised in developing natural gas fields, this leaves fewer rigs available for developing oil fields. Factors like taxation, political unrest and government sanctions may also have an impact.

Although we test for some of these factors, the crude oil price is the main explanatory variable in our estimation. We have computed smoothed prices on logarithmic form ($\ln X_t$ in Eq. 4) over the last 3, 6, 12, 24, 30 and 36 months[9] and have estimated models under these alternative smoothing assumptions for all the regions. For each region we have picked a maintained model, characterised by a specific smoothing choice, based on the statistical results. We present the results for this model in the main text. In addition results based on two alternative smoothing assumptions are reported in Appendix B (to ease the comparison, results for the model with prices smoothed over 12 months are reported for all regions, even if the merits of this model vary somewhat over the regions).

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[7] This is consistent with less restricted adjustment processes wrt. rig activity and the formation of price expectations. One consequence of this is that we cannot compute the values of the underlying structural parameters, as Farzin (2001) does. However, these are very much dependent on the specification of the functional forms.

[8] Seasonal variations could of course also be stochastic rather than deterministic (see Section 3).

[9] Prices are computed by taking the mean of the logarithm of the prices for the last 3, 6, etc. months, including the current month.

The data for active oilrigs have been collected from the Baker Hughes Rig Counts[10]. The Rig Count is a census of the number of drilling ('rotary') rigs actively exploring for or developing oil or natural gas. Very small rigs like truck mounted rigs are not included, but the rigs are of course very heterogeneous. Onshore rigs are for instance usually smaller and more flexible than offshore rigs. In the US weekly census a rig is counted as active if it is actually drilling. In the monthly international rig count, which has figures for most regions and countries outside the US, the rig is active if it is drilling at least 15 days during the month counted. This makes the Baker Hughes counts a good indicator of current rig activity, and thus exploration and development effort, rather than a measure of the rig fleet's capacity.

For the US, we have weekly activity data from July 1987 until July 2002, which have been converted into monthly data. For some other regions in the world (Europe, the Middle East, Latin America, Africa and Asia Pacific), we have monthly oilrig activity data for the period January 1995 until July 2002. This may seem as a short time period for estimating a dynamic relationship, but the results as well as earlier indications (cf. eg. Abraham, 2000, and OGJ, 2003a) indicate that 7 years may be sufficient to obtain valid results. The most important region not covered by the census is the former Soviet Union, but also onshore China is lacking. In addition, the Canadian data do not separate between gas- and oilrig-activity, and Canada is therefore neither included in our empirical analysis.

As indicated above, the census also reports the number of active rigs divided by countries within regions. We have chosen to mainly focus on regions, partly because the number of rigs in individual countries usually are too small to give good statistical results, but also because rigs may be moved around within a certain geographical area. Of course, this does not mean that the regional distribution reflects distinct rig markets - rigs may be moved both within and between regions, but only at a cost. Another aspect that we do not take into account is that several oil producers are operating in different regions, and may therefore not consider rig activity in different regions independently. For some of the regions that include OPEC members, we will estimate models based on data both with and without the OPEC countries. The classification of the regions is given in Appendix A.

We have price data from Petroleum Intelligence Weekly (PIW) since 1986 for the crude oil prices Brent Blend, US WTI, Nigeria and Dubai. These are on weekly basis, but have been converted into monthly data before the smoothing operation. We use the US WTI for the US and Latin America, Brent Blend for Europe, Dubai for the Middle East and Asia Pacific, and Nigeria for Africa. These prices are all nominal, and reported in dollars. The real prices have been calculated using a price index for all

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[10] See http://www.bakerhughes.com/investor/rig/index.htm

manufacturing industries in the US from the American Bureau of Labor Statistics[11]. This is a monthly index starting in January 1992. Hence, the sample period is 1992:1-2002:7 for the US region, and 1995:1-2002:7 for the other regions.

When modelling US oilrig activity the price of natural gas was included as an explanatory variable, as we believed that this would have some second-order effect on the demand for oilrigs through the demand for gasrigs (see above). However, this variable turned out to be highly insignificant. As it neither affected the other coefficients in any substantial way, it has not been included in the reported results.

3 Modelling framework

Our point of departure is the following equilibrium correction model augmented with a stochastic trend:

$$(5)\ \Delta y_t = \sum_{i=1}^{p} \beta_i \Delta y_{t-i} + \sum_{i=0}^{q} \gamma_i \Delta x_{t-i} + \lambda y_{t-1} + \theta x_{t-1} + \kappa M_t + \mu_t + \varepsilon_t .$$

This relation is based on a reparameterisation of Eq. (4). In Eq. (5) y_t denotes the log of oilrigs, x_t is the mean of the log of the oil price over a certain number of periods and M_t is an m×1 vector of deterministic variables (dummy variables). The symbol Δ denotes the first difference operator, μ_t is a stochastic trend to be specified later and ε_t is an error term. As we will elaborate on later it is reasonable to assume that the y_t and x_t are non-stationary variables integrated of order 1, i.e. I(1). An important question is whether these two variables is a part of a long-run relationship, which also involves the stochastic trend. This can be investigated by testing the significance of y_{t-1} in Eq. (5). Since we are dealing with non-stationary variables, the estimator of λ will have a non-standard distribution. We utilise the critical values provided by Banerjee et al. (1998), even if these are related to the special case when the stochastic trend degenerates to an intercept and a linear trend. Thus the critical values only have an approximate status. The estimator of θ will also have a non-standard distribution, whereas the usual t- and F-distribution can be used for testing lag orders of the differenced variables. Our inference, that is estimation and testing, is based on the assumption that x_t is weakly exogenous with regard to the parameters in Eq. (5). When we later consider calculations of elasticities we will also need that x_t is strongly exogenous. For these two concepts consider Engle et al. (1983). In determining the lag orders p and q we emphasise that the error term, ε_t, should be a white noise process such that Eq. (5) can be

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[11] Taken from http://data.bls.gov/cgi-bin/surveymost.

considered to be well-behaved from a statistical point of view. To ensure normality, irregular interventions (corresponding to dummy variables for outlier values) were necessary in a couple of the regions. A level intervention, corresponding to a step dummy variable, was included for Africa.[12] The basis for the different dummies is explained in the relevant subsections of section 4. We use the following notation for dummies: By $DIy.m_t$ we define an impulse dummy taking on the value 1 in month m of year y and zero otherwise. Correspondingly $DSy.m_t$ is a step dummy taking on the value 0 until the month preceding month m in year y. Thereafter the value is one. If there is cointegration the long-run elasticity of rig activity with regard to oil price is given by $E_\infty=-\theta/\lambda$. This follows from the Bårdsen (1989) formula and the fact that the model is formulated in logarithms. In Appendix C we provide formulae for how elasticities with respect to non-smoothed prices can be calculated after an arbitrary number of periods. We report long-run elasticities without providing t-values. Since the estimator of the long-run elasticity involves the ratio of two estimators, the usual way to proceed is to use the delta method (cf. e.g. Kmenta, 1986, p. 486). However, this requires knowledge of the covariance between the two estimators, which unfortunately is not reported by STAMP 6.2 (cf. Koopman et al., 1999), which we use in our empirical analysis.

For the long-run solution to exist the model has to be dynamically stable, which requires that all the roots of the characteristic function of y_t exceed one.[13] We have checked this condition in each case and it is fulfilled for all models for which we report results. When $\beta_i = 0$ for all i (i=1,...,p), dynamic stability requires that λ is between -2 and 0. This implies that the coefficient of the lagged dependent variable, when the equation is written on level form, has an absolute value less than one, cf. Johansen (1995, p. 46). If the coefficient is between 0 and -1, the disequilibrium term will go monotonically towards zero. With a coefficient between -2 and -1, there will be an "overshooting" effect. The disequilibrium will cycle around zero, with gradually smaller cycles, until reaching zero. Of course "overshooting" effects may also be present in the more general case when lagged differences of the left-hand side variable are included on the right hand side in Eq. (5).

As already noted a special case emerges when the stochastic trend degenerates to an intercept and linear trend. This specification was tried initially. However, it turned out that in this type of model it was hard to obtain cointegration and the residuals were not well-behaved. The assumption of deterministic trend is an increasingly contested area in time series econometrics (see e.g. Harvey et al. (1986)), and the concept of a stochastic time trend, and also stochastic seasonal effects, is increasingly being used. As

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[12] See Harvey (1989, p. 397) for further discussion on intervention analysis.

[13] For the close relationship between cointegration and dynamic stability see Boswijk (1994).

described in Section 2, there are several more or less unobservable factors that affect oilrig activity in different directions, substantiating the need for a stochastic trend rather than a deterministic one.

In light of this we included a stochastic trend in Eq. (5). We also considered stochastic seasonality, but since we generally were unable neither to detect any stochastic nor any fixed seasonality, we have not included such terms in Eq. (5). The inclusion of a stochastic trend, i.e. a time trend that is allowed to change over the sample period, did provide us with a significant cointegration coefficient in most of the modelled regions.

The stochastic trend is generally specified as a random walk with drift (see Harvey, 1989)

(6) $\mu_t = \mu_{t-1} + \beta + \eta_t; \quad \eta_t \sim \text{NIID}(0, \sigma_\eta^2)$.

Our model to be estimated thus consists of Eqs. (5) and (6). Since lagged values of y_t occur on the right hand side of Eq. 5, the stochastic trend will not be independent of the observed right hand side variables. However, since the estimation algorithm utilises the conditional distribution of Δy_t given past information, this does not prevent consistent estimation, but as Harvey (1989,p. 385) remarks the asymptotic standard errors may be influenced. The error terms ε_t and η_t are assumed to be uncorrelated and Gaussian distributed with zero means and variances σ_ε^2 and σ_η^2 respectively. μ_t is the level (the actual value) of the stochastic trend, hence η_t allows the level of the trend to change over time. β represents the slope of the trend, and is assumed to be constant. The stochastic trend may have different specifications, β may for example be restricted to zero if it turns out to be insignificant. When σ_η^2 equals zero, the model degenerates to an intercept and a linear trend. Thus testing the null of a zero variance in the stochastic trend equation against the alternative with a positive variance is of interest. However, since this test involves parameters on the boundary of the admissible parameter set, non-standard inference is involved. For instance the usually χ^2-distributed LR statistic cannot be used. Harvey (2001) considers different tests for a deterministic trend, but in a framework without observed covariates on the right hand side of the equation. Quite recently Bailey and Taylor (2002) has provided a locally best invariant (LBI) test for a random walk component in a model with deterministic regressors and non-orthogonal unobserved components. However, even if we assume orthogonal components our model is in another respect more complicated since lagged values of the endogenous variable occur on the right hand side of the equation. In light of this we do not carry out any formal test of the above hypothesis. However, it is fairly clear in our model specification that the inclusion of a stochastic trend provides us with a model that works, by ensuring significant coefficients for the equilibrium correction term. Thus all the econometric relations presented in the next section have been estimated with a stochastic trend. Since we in none of the regions obtain a significant estimate of β, the stochastic trend is represented by a random walk.

Let $\hat{\mu}_{t|T}$ denote the smoothed estimate of the stochastic trend at time t, i.e. this is the estimate of the unobserved component using all available data. This estimate is obtained by running the Kalman smoother. In the next section we report graphs of the smoothed trend estimates related to the maintained model for each region. Furthermore let q be defined by $q = \left(\sigma^2_{\eta\eta} / \sigma^2_{\varepsilon\varepsilon}\right)$, which in the simpler framework with no observed right hand side variables, is labelled the 'signal to noise ratio'. The smaller the estimate of $\sigma^2_{\eta\eta}$, the smoother the estimated trend will be. We report the estimated q-ratios for all the estimated models.

Models of the type described above can be estimated by the STAMP-program.. To estimate the parameters of the models, that is, the hyperparameters and the regression parameters, STAMP maximizes the diffuse log-likelihood function taking properly care of the initial diffuse conditions. For a technical discussion of maximum likelihood estimation in a situation with diffuse initial conditions, cf. De Jong (1991).

4 Empirical analysis

4.1 Unit root tests

As mentioned in Section 3, the assumption of I(1)-variables is crucial for the ECM-method to be valid. For the oilrig activity, our assumption was supported by unit root tests. Neither the ADF-test (Augmented Dickey-Fuller) nor the Phillips and Perron (1988) test (PP) rejected the null hypothesis of I(1) for the level series, whereas the null hypothesis of I(2) against the alternative hypothesis of I(1) was rejected in all regions. This latter test is implemented using the differenced series. However, for the smoothed prices, the picture is not as clear. The (PP-)tests of the level series do not reject the hypothesis of I(1), but the tests of the differenced series are somewhat unclear as to the rejection of the hypothesis of I(2), in most situations. It might seem like the smoothing of the prices influences the stationarity properties of the variables, giving the series I(2)-characteristics[14].

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[14] Note that the I(2)-hypothesis is rejected when using the original "non-smoothed" prices.

However, there are a number of different tests for stationarity, and numerous papers (e.g. Leybourne and Newbold, 1999) have shown that these can give different results for the same sample. Both the ADF-test and the PP-test have been criticised for having low power, tending not to reject H_0 when this in fact should be rejected. These tests have H_0: non-stationarity versus H_1: stationarity. There are also tests that have the opposite hypotheses, with stationarity as the null hypothesis. Among these is the KPSS-test (Kwiatkowski et al., 1992). This test was performed on one of the series (US 12months), which neither ADF nor PP could reject as being I(2). The KPSS-test gave the result that it could not be rejected as I(1). Thus, different tests give different conclusions about the stationarity properties for the same series. In addition to the inaccuracy of the tests, the small size of our sample might influence the stationarity tests, making it more difficult to draw conclusions about the stationarity of the series. The upshot of all this is that, due to the uncertainty of the tests, we have chosen to base our analysis on the assumption that the variables are I(1), even though the unit root tests do not fully substantiate such an assumption.

4.2 Empirical results in relation to the equilibrium correction models

Estimations have been done in STAMP 6.2, a program that allows estimation with a stochastic trend. As mentioned in the preceding section, the variables need to be cointegrated to obtain valid estimates. We test this by inspecting the t-value for the lagged dependent variable in the model. At a 5% (1%) level of significance, the critical value is about 3.75 (4.35), cf. Banerjee et al. (1998,Table 1b). For all regions, the estimations show t-values of the estimator of λ with satisfactory level for at least one of the price variables. Thus, when allowing for a stochastic trend, it seems that oilrig activity and oil price are in fact cointegrated.

We report different diagnostics, and especially test for absence of autocorrelation, heteroskedasticity and non-normality. Generally we find the residuals to be well-behaved, but will comment on the cases where we find that this is not so.

Table 1 reports the estimation results for the different regions. The number of months indicated refers to over how many months the price is smoothed. For each region we only report the results of the maintained model; the alternative model results that are based on other smoothing assumptions are presented in Appendix B. Note that the short-run or immediate price elasticity is given by the coefficient for Δx_t, divided by the number of months on which the smoothing of the price is based. The table also reports long-run price elasticities, derived from the estimation results.[15] Figure 2 is informative about the

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[15] Calculated as $E_\infty = -\theta / \lambda$, where θ is the lagged price coefficient, and λ is the lagged oilrig activity coefficient.

speed of adjustment, since it provides calculated price elasticities in the short-, intermediate- and long-run (the underlying formulae are presented in Appendix C). Figure 3 shows the smoothed estimate of the stochastic trend for the maintained model in each region. Below we discuss the estimation results for each region separately. In Section 5 we sum up our conclusions.

Table 1. Econometric relations for oilrig activity in different regions. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	US (6 months)		Europe (6 months)		Latin America (12 months)		Non-OPEC Latin America (12 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable	-0.204[a]	-3.941			-0.173[b]	-4.028		
Δy_{t-1}	0.420	5.386	-0.503	-5.735				
Δy_{t-2}	-0.334	-4.068						
Δx_t	0.885	2.428	0.768	1.519	0.083	0.194	0.608	0.785
Δx_{t-1}	-0.556	-1.492	-1.144	-2.232			-1.047	-1.342
y_{t-1}	-0.202	-3.555	-0.212	-3.299	-0.422	-4.964	-1.022	-9.415
x_{t-1}	0.343	4.207	0.144	2.676	0.359	3.195	0.969	3.939

	US	Europe	Latin America	Non-OPEC Latin America
Long-run elast.	1.698	0.679	0.851	0.948
Variance ratio and diagnostics:				
q[c]	0.130	0.006	0.304	3.932
Std. Error	0.054	0.088	0.046	0.060
Normality[d]	1.912	1.425	0.075	0.648
H[e]	0.630	1.116	1.389	1.444
DW	1.984	2.074	2.077	2.008
Q[f]	2.501	5.471	3.445	2.987
R^2	0.452	0.408	0.285	0.156

[a] The included dummy-variable is DI00.1.

[b] The included dummy-variable is DI02.4.

[c] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[d] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[e] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p.119). The degrees of freedom in both the numerator and denominator are 39 for the US and 29 for the other regions.

[f] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom, where q is p+1 less the number of estimated variance components. p is 9 for US and 8 in the other regions.

Table 1 (continued). Econometric relations for oilrig activity in different regions. Smoothing assumption with regard to oil prices in parenthesis

Slope coefficients related to indicated variables	Asia Pacific (24 months)		Non-OPEC Middle East (24 months)		Africa (12 months)		Non-OPEC Africa (12 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable					0.663[a]	6.398		
Δx_t	0.835	0.667	3.309	2.571	0.810	0.753	0.182	1.371
Δx_{t-1}	-2.230	-1.776			-3.028	-2.770	-0.083	-0.643
y_{t-1}	-0.749	-7.301	-1.228	-11.568	-1.051	-11.483	-0.857	-7.664
x_{t-1}	0.357	1.881	0.745	2.278	0.757	2.779	0.190	1.626
Long-run elast.	0.478		0.607		0.720		0.222	
Variance ratio and diagnostics:								
q[b]	0.526		4.975		1.616		0.189	
Std. Error	0.050		0.053		0.103		0.472	
Normality[c]	0.156		6.057		9.347		29.570	
H[d]	0.662		0.679		0.518		0.850	
DW	1.906		2.014		2.170		1.959	
Q[e]	5.479		8.772		5.494		6.061	
R^2	0.113		0.185		0.403		0.201	

[a] The included dummy-variable is DS02.2.

[b] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[c] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[d] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p.119). The degrees of freedom in both the numerator and denominator are 29.

[e] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

Figure 2a. **Price elasticities for oilrig activity at different number of months after a sustained price increase for US, Europe, Latin America and Non-OPEC Latin America**

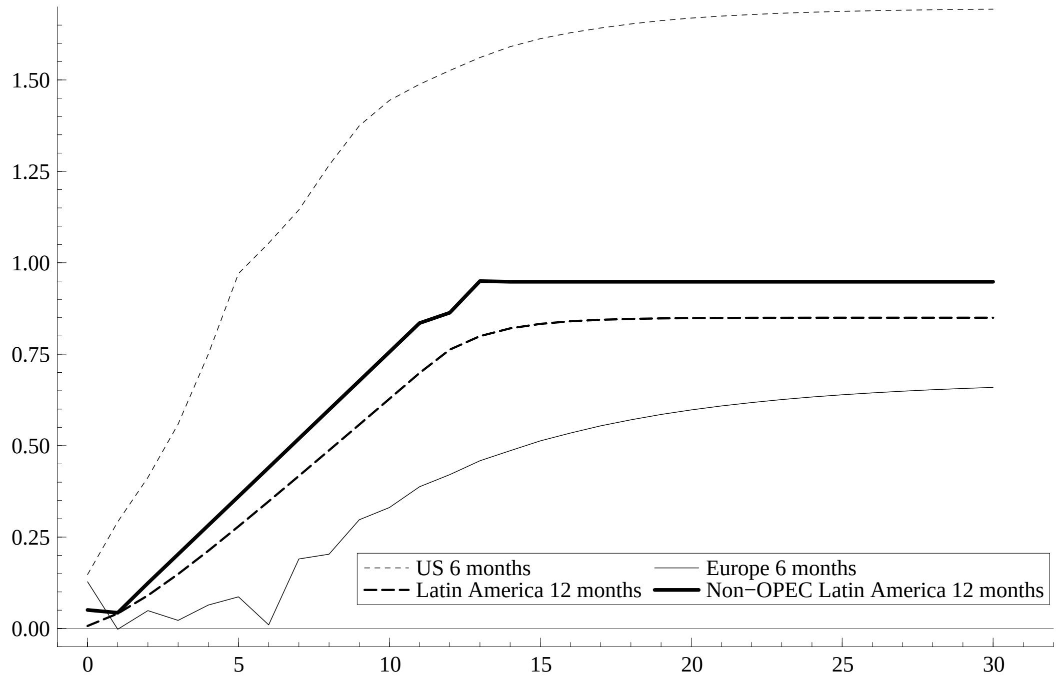


Figure 2b. **Price elasticities for oilrig activity at different number of months after a sustainedprice increase for Asia Pacific, Non-OPEC Middle East, Africa and Non-OPEC Africa**

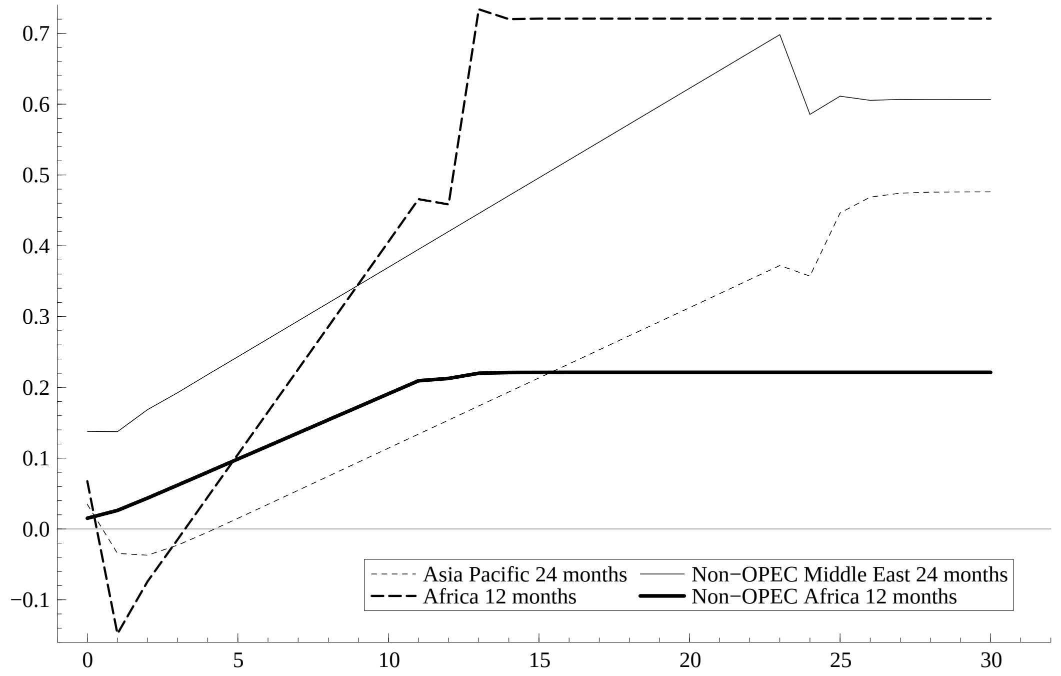
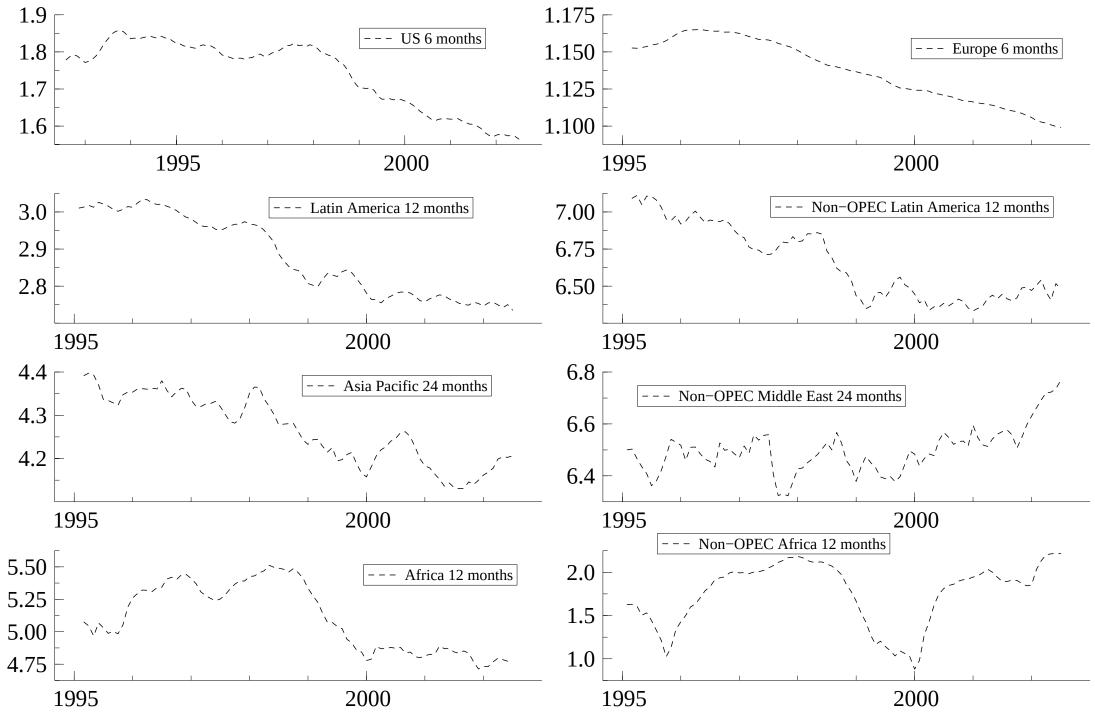


Figure 3. Extracted stochastic trend ($\hat{\mu}_{t|T}$) in different regions



4.3 United States

For the US initial estimations revealed a high degree of autocorrelation in the residuals. To correct for this, two lags of the differenced oilrig activity variable were included. Inspection of the residuals revealed a very significant outlier value for January 2000 (cf. Abraham, 2000), which was corrected for by introducing a dummy variable for this month, as this improved the diagnostic test for normality substantially. A possible explanation for this outlier may be that insecurity surrounding the new millennium (Y2K) led to decreased activity. The maintained model for the US is the one based on 6 months prices, whereas the two alternative models are based on 3 and 12 months prices, respectively. For the two latter models the lag-specification of the differenced independent variable differs somewhat from that used for the maintained model.

We see from Table 1 that the short-run price effect is clearly significant, with an immediate price elasticity of about 0.15. Thus, the US region seems to react fairly quickly to price changes. The results indicate that the US rig market is very flexible - oil producers are able to rent rigs on a short notice. This partly reflects that the rig market in the US is competitive, and that oil producers face few regulations compared to other regions. Another important explanation may be that the drilling activity is to some degree occurring in so-called marginal and smaller fields. Although the smallest rigs are not included in the Rig Count, most of the US rigs are used onshore and are more flexible than e.g. offshore rigs mostly used in Europe. Producers in the US market seem to expect that price changes will last at least long enough to affect the profitability of extraction. We observe that the dummy variable for January 2000 is, as expected, highly significant and negative.

The estimated long-run price elasticity is 1.7. From Figure 2 we see that more than half of the long-run price effect seems to be realised after only 6 months. This confirms the impression above of a fast reaction in the US rig market when oil prices are changing. According to the results for the US in Table B.1, the estimated long-run elasticity based on the two other specifications does not deviate much from the one obtained using the maintained model.

Figure 3 shows the smoothed estimate of the stochastic trend for the US. We observe that in the period 1993-1998 there seems to be no clear positive or negative trend, but from 1998 the trend is clearly negative. This may reflect that most of the US region is fairly mature - oil production in the US has been more or less declining since 1985. One important exception is the deep waters of Gulf of Mexico, where exploration and development increased rapidly during the first part of the 1990's. In this period, then, this expansion to a new area was able to counteract the general tendency of decline. The negative trend is also observed in several other regions, which indicates that there may be a common explanation for this trend, e.g. technological progress reducing the time to drill a well. The drop in 1998-1999, which is seen for several regions, could be connected to the dramatic price reduction in this period.

4.4 Europe

As for the US we use 6 months prices in the maintained model for Europe. It was necessary to include one lag of the differenced dependent variable to get rid of autocorrelation in the residuals.

The immediate price elasticity is slightly lower than the one found for the US, but contrary to in the US it was not significant, cf. Table 1. Moreover, the estimated long-run price elasticity is 60 per cent lower than the US elasticity. As can be seen from Table B1 the long-run elasticity for Europe is somewhat higher in the two alternative models based on prices smoothed over 12 and 24 months, but still the main impression is that the long-run elasticity is higher for the US than for Europe. We also see from Figure 2 that the speed of adjustment is much slower in the latter region. Thus, it seems that oil producers are less price sensitive in Europe than in the US. One reason may be that oil exploration and development are more regulated in Europe, particularly in Norway. Another explanation could be that in Norway at least the net tax rate is very high (78%), so that the government is in fact bearing most of the risk of e.g. price reductions. Drilling in Europe is mainly taking place offshore, with large installations, and in Norway there is a predominance of long-term contracts (OGJ, 2003b), which may also reduce the price sensitivity.

Another explanation could be that the rig fleet in Europe generally has less idle capacity than the fleet in the US, reducing the ability to increase rig activity when the oil price is rising. According to ODS-Petrodata[16], the rig utilization rate in the North Sea has been higher than in both the US and elsewhere over the last years. Moreover, in the US a larger portion of rigs are used for gas exploration and development. This increases the flexibility when it comes to oilrig activity, as rigs used for gas drilling may be converted to oilrig activity.

The estimated stochastic trend for Europe is showing a clear (though weak) tendency of decline from 1996-1997 (see Figure 3). This is consistent with the development of the now matured North Sea - UK and Norwegian oil production reached a peak level around 1996, consistent with a reduced demand for drilling afterwards.

Since Europe consists of two main oil-producing countries, which are quite different in various aspects, we also estimated the oilrig activity in the UK and Norway separately. The results, which are reported in Appendix B, suggest as expected that oil producers in Norway are considerably less price responsive than those operating in the UK.

4.5 Latin America

Regarding Latin America we consider two different definitions of the region, according to whether the OPEC member Venezuela is included or not. Venezuela, being the most important country in Latin America when it comes to oilrig activity, has played different roles within OPEC over the last decade. Sometimes it has played tough disregarding its quota (i.e., behaving like a price *taker*), whereas at other times it has been an eager supporter of restrained OPEC production (i.e., behaving like a price *maker*). In the latter case it is not reasonable to assume that the country considers the oil price as exogenous.

Starting with Venezuela included in the Latin American sample, the maintained model is the one with smoothed prices over 12 months. We have included a dummy variable for April 2002. This month there was an attempted coup d'état against president Hugo Chavez in Venezuela, which greatly affected the oil industry in the country, causing a sharp decline in the oilrig activity. With this dummy included, all

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[16] http://www.ods-petrodata.com

diagnostic tests are fairly satisfactory, apart from the somewhat low R^2 -value. The two observed variables seem to be cointegrated, cf. the high t-value for the estimator of λ. From Table B2 we see that the estimated q-ratio is zero using 3 and 6 months prices, which implies that the stochastic trend degenerates to an intercept and a linear trend. However, note that in these two cases the t-values for the estimator of λ are rather low, such that the two observed variables do not seem to be cointegrated.

Before discussing these results more fully, let us turn to Latin America without Venezuela. The maintained model is again the one with 12 months prices. Since Venezuela is removed, the significant outlier value in April 2002 naturally disappears. The diagnostic tests are satisfactory, apart from the low R^2 -value. Here, too, the two alternative models are based on 3 and 6 months prices. As opposed to when Venezuela is included, the estimated q-ratio is positive in all three cases, but varies a lot according to the smoothing assumption.

As can be seen from Table 1, the immediate price effect is much higher when Venezuela is excluded, but it is not significant. In the intermediate and long-run the price effect is still higher without Venezuela, but not much (cf. Figure 2). Nevertheless, the results indicate that Latin American countries other than Venezuela seem to react more heavily to price changes than Venezuela itself.[17] This may support the view that Venezuela is not acting like a regular price taker - its oil supply is influenced by its OPEC membership.

Compared to the other regions, Latin America without Venezuela shows less response to oil price changes than the US and Europe, despite a dominance of onshore drilling. This reflects that the region has a large share of state oil companies (especially Mexico). In the long-run the price elasticity is still significantly below the US, but much higher than in other developing regions. This result may indicate that the state oil companies in Latin America after all are fairly market oriented, but it also reflects that most of the countries (except Mexico) have widespread cooperation with international oil companies.

Oil production in Non-OPEC Latin America increased by around 20 percent from 1995 to 1998, and then by merely 5 per cent over the next four years. Thus, the demand for oilrigs was reduced over this period. This pattern is also seen in Figure 3, where we note that the estimated stochastic trend is slightly falling until 1998, and then shifts downwards to a new level.

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[17] This is also supported by estimations of Venezuela separately, where the short-run price effect is negative (but insignificant) and the long-run elasticity is 25 per cent smaller than in Non-OPEC Latin America.

4.6 Asia Pacific

For this region the maintained model is the one with prices smoothed over 24 months. In this region we do not include any lags of the differenced dependent variable. All diagnostic tests seem satisfactory, but the explanatory power of the model is somewhat low. The two alternative models for this region, cf. Table B3, are based on prices smoothed over 12 and 36 months respectively. In the latter model the estimated long-run elasticity is near 1, whereas in the former it is below 0.2. In the maintained version the elasticity is a little bit less than 0.5. Thus the results related to the long-run elasticity are not very robust across model versions. It should also be noted that some of the intermediate elasticities are negative, see Figure 2. This goes together with somewhat imprecise estimates.

In the maintained model the immediate price effect is insignificant and even smaller than in Non-OPEC Latin America (see Table 1). The estimated long-run price elasticity is also lower than in most other regions, and so is the speed of adjustment (see Figure 2). Only 30 per cent of the long-run price effect is obtained after one year. The two most important countries in this region are India and Indonesia,[18] but several other countries also have significant rig activity. As most of the drilling is taking place onshore, using relatively small rigs, the results probably reflect that the Asian countries have had a far more regulated oil industry than most Western countries, with Norway as an important exception. Oil companies therefore need much more time to prepare or get permission to explore and to develop fields (according to IPE (2003, p. 163-64) both India and Indonesia has recently moved steps towards privatisation of the oil industry). Another explanation could be that profitability concerns are less important than in the Western regions, e.g. due to some sort of capacity or credit restrictions. For instance, IEA (2003) emphasises that non-major oil producers operating outside OECD may have difficulties raising capital, e.g. due to higher risks.

Figure 3 shows that the stochastic trend in Asia Pacific is generally negative, just like in the former regions. Oil production in this region was rising steadily until 1997, when the economic downturn began in Asia. This might have affected spending in the petroleum sector, although the main oil producing countries in this region were not hit as hard as the more industrialised countries.

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[18] Being an OPEC member, it could be argued that Indonesia should not be treated together with Non-OPEC Asian countries. However, over the last years Indonesian production has not really been constrained by their OPEC membership, according to most analysts. Yet, we have tested empirically the effect of removing Indonesia from the sample, and this does not change the results for Asia Pacific (not reported).

4.7 Non-OPEC Middle East

The Middle East region comprises the most important OPEC-countries, and the assumption of exogenous price for this region is obviously troublesome. Thus, we only estimate models for this region with the OPEC-countries excluded. As for Asia Pacific smoothed prices over 24 months are used for the maintained model. Apart from some indication of non-normality the diagnostics are satisfactory. However, the explanatory power is somewhat low for this region. In the two alternative models (cf. Table B3) we consider prices smoothed over 24 and 30 months. The results seem fairly robust across the three models.

The immediate price elasticity for Non-OPEC Middle East is in fact significant, and comparable to the one in the US. This is quite surprising, inasmuch as the countries in Non-OPEC Middle East are quite heterogeneous - the main countries being Egypt, Oman and Syria. We observe that the speed of adjustment is quite slow, and that slight overshooting effect following a sustained price change is present, cf. Table 1 and Figure 2. It is difficult to give a reasonable explanation for this. One possibility could be that oil producers' price expectations are somewhat different from what we assumed in Section 2. For instance, they may expect a further price increase (decrease) when the oil price starts to rise (fall), and therefore adjust rig activity in advance of (expected) price changes. However, it is difficult to explain why this should turn up mainly in Non-OPEC Middle East. In the long-run the price elasticity for this region is in the range of the elasticities in Asia Pacific and Europe, cf. Table 1. The stochastic trend is fairly constant (see Figure 3), which reflects that oil production in this region has been fairly stable over this period.

4.8 Africa

The Africa region includes three OPEC countries, i.e., Algeria, Libya and Nigeria. As these are usually not considered among the core OPEC members, we start by estimating relations for Africa including these three countries. Subsequently we consider Africa without the OPEC-members. In the maintained model, which is based on prices smoothed over 12 months, we have no lags of the differenced dependent variable. Due to a redefinition of the variables in our dataset for Africa (i.e., for Algeria and Libya), the number of active oil rigs increases substantially in February 2002, and this increase persists for the rest of the sample. Accordingly, we have included a step dummy from 2002:2. Still, there are some sign of non-normality.

In Table B4 we have reported results for two alternative models for Africa based on 24 and 36 months prices, respectively. For all three models there is some overshooting. The long-run elasticities are considerably higher than in the maintained model, but we consider the model based on 12 months prices as the maintained model since it is somewhat more well-specified than the two other. The Box-Ljung statistic is significant for the model based on 36 months prices, and the non-normality is more severe for the two models in Table B5 than for the model in Table 1.

From Table 1 we see that the short-run elasticity in Africa is quite similar to Non-OPEC Latin America (not significant). The long-run price elasticity is lower, though, but higher than in Asia Pacific and Non-OPEC Middle East. From Figure 2 we see that the intermediate price elasticity is slightly negative for a short period initially, which we believe is due to estimation inaccuracy.

When we remove the three African OPEC countries but use the same econometric specification as for the entire region, there are still problems with the normality test, but the rest of the tests are satisfactory.[19] Both the short- and long-run elasticities are very small compared to other regions. The two alternative models for Non-OPEC Africa, which are based on prices smoothed over 24 and 36 months, respectively, yield negative long-run elasticities (cf. Table B4) and could thus be disregarded as potentially maintained models. The econometric results may reflect that African countries are fairly unstable, with regime changes, civil wars (e.g. in Angola) and other political instability that may influence oil activity. Although some African countries like Angola have become fairly open to the international oil industry (IPE, 2003, p. 115-116), it seems that the oil price only to a limited degree has determined the speed of expansion of the oil sector.

The estimated stochastic trend for Non-OPEC Africa has two drops, in late 1995 and in 1999 (see Figure 3). The last one may be due to an outbreak of the civil war in Angola in late 1998.

5 Discussion and conclusions

The estimation results described above indicate that all regions seem to experience enhanced oilrig activity when the oil price increases, at least in the long-run. Although the econometric results are somewhat mixed for some regions, there is a general picture of price-conscious oil producers in all regions considered. Moreover, the results clearly show that the US has by far the highest price elasticity, both in the short and long-run. Whereas the US shows noticeable effects on oilrig activity after just 3 months, some developing regions need more than a year before such effects occur. The US elasticities

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[19] Dummies for outliers were experimented with, but were not included in the final estimations, as they led to autocorrelation of the residuals. In addition, the 2002:2-dummy is excluded since it is related to two of the OPEC countries.

are also fairly robust across estimation models building upon different assumptions with respect to the smoothing of the oil price.

A long-run price elasticity of 1.7 in the US means for instance that a permanent price increase from $25 to $30 per barrel of oil will produce a growth in oilrig activity of about 35 per cent after some years. In the other main regions (except Non-OPEC Africa) a similar price adjustment would raise oilrig activity by between 10 and 20 per cent. If we consider all the effects on rig activity in Non-OPEC together, i.e., compute a Non-OPEC price elasticity based on the results of the individual regions, we get a long-run price elasticity close to unity.[20]

There may be several reasons for why oilrig activity in the US reacts faster and stronger to price signals in the oil market than other regions, and some of them were mentioned above in Section 4.3. First of all, production of oil in the US is carried out by private oil companies, and there are relatively few governmental restrictions on their activities, except for the ban on exploration and development in certain areas (cf. Rutledge, 2003, and IEA, 2003). In most other parts of the world the oil production is either managed directly by state oil companies, or the government controls the oil production activity more or less strictly through licences on exploration and development. Although most governments are concerned with profits of their activity, they may also be concerned with other aspects such as a balanced and stable development of the oil industry and a stable income over time. Moreover, with governmental control two decision phases are often necessary; first the company decides to invest, and next the government approves or not (or the government launches a licensing round and the companies consider whether to explore or not). This may explain why the price effect is insignificant in the short or even medium term for several regions.

Another reason may be that oilrig activity is more price sensitive in areas where the total unit costs of oil production (including field development) is high. US fields are generally more expensive than fields in other regions, but also offshore fields in Europe and Latin America are relatively costly. These are the three regions with the clearest price response. A third reason may be that the US (except for the deep Gulf area) is a mature oil province, with a large infrastructure. In such a province, owners of small and short-lived fields with close connection to existing, larger fields may react quicker to oil price changes, as they are only concerned with the near-term price levels. Moreover, most of the US drilling is

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[20] The average number of oilrigs over the last 12 months of the estimation period was used as weights here. This weighing ignores that oilrigs are not a homogeneous unit, i.e., one large oilrig may develop more oil reserves than several small ones. Moreover, note that several important Non-OPEC countries, such as the former Soviet Union, Canada and onshore China, are not included in the data.

occurring onshore, which presumably is more flexible than offshore drilling, and therefore able to react quicker to price changes.

From a methodological point of view, we would like to emphasise our econometric approach using equilibrium correction models with a stochastic trend. As pointed out in the presentation of the results, the presence of a stochastic trend seems to be crucial to the model specification. This reflects that there are other aspects than the oil price that influence the rig activity (such as technological change and resource depletion), and that this is better specified as a stochastic trend rather than a deterministic one.

Finally, it is important to stress that the results for some of the regions, such as Africa, should be considered with caution. Yet, we conclude from the empirical analysis that oil price changes can induce significant changes in oilrig activity, and hence investment in new oil production capacity in Non-OPEC. For OPEC this means that a long-lasting price level in the upper $20's for a barrel of oil will eventually bring about much more Non-OPEC oil than the price level observed in the late 1990's

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Appendix A

The regions, for which we consider econometric modelling, consist of the following countries (an asterisk* refers to OPEC countries):

- **The United States**
- **Europe:** Denmark, France, Germany, Netherlands, Hungary, Italy, Norway, Poland, Romania, Turkey, United Kingdom, Yugoslavia, "others".
- **Non-OPEC Middle East:** Egypt, Oman, Pakistan, Sudan, Syria, Yemen, "others".
- **Latin America:** Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Peru, Trinidad, Venezuela*.
- **Africa:** Algeria*, Angola, Congo, Gabon*, Libya*, Nigeria*, South Africa, Tunisia, "others".
- **Asia Pacific:** Australia, Brunei, India, Indonesia*, Japan, Malaysia, Myanmar, New Zealand, Offshore China, Papua New Guinea, Philippines, Taiwan, Thailand, Vietnam, "others".

Appendix B

Supplementary econometric results

Table B1. Econometric relations for oilrig activity in US and Europe. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	US (3 months)		US (12 months)		Europe (12 months)		Europe (24 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable	-0.209[a]	-4.266	-0.179[a]	-3.931				
Δy_{t-1}	0.352	4.665	0.366	4.551	-0.446	-4.701		
Δy_{t-2}	-0.264	-3.261	-0.149	-1.734				
Δx_t	0.486	3.423	0.708	1.209	0.264	0.261	3.079	1.649
Δx_{t-1}					1.179	0.905		
$\Delta \mathrm{x}_{t-2}$	-0.529	-3.162			-1.829	-1.778		
y_{t-1}	-0.209	-3.907	-0.554	-5.651	-0.336	-3.300	-1.523	-16.702
x_{t-1}	0.380	5.0586	0.793	4.040	0.274	2.761	1.410	2.609
Long-run elast.	1.818		1.431		0.815		0.926	
Variance ratio and diagnostics:								
q[b]	0.181		1.459		0.032		11.548	
Std. Error	0.052		0.053		0.089		0.092	
Normality[c]	0.255		1.661		1.403		2.043	
H[d]	0.591		0.779		1.016		1.100	
DW	2.038		2.046		2.025		1.952	
Q[e]	3.660		2.978		2.967		3.341	
R^2	0.487		0.465		0.378		0.358	

[a] The included dummy-variable is DI00.1.

[b] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[c] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[d] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator is 39 for US and 29 for Europe.

[e] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is 9 for US and 8 for Europe.

Table B2. Econometric relations for oilrig activity in Latin America and Non-OPEC Latin America. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	Latin America (3 months)		Latin America (6 months)		Non-OPEC Latin America (3 months)		Non-OPEC Latin America (6 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable	-0.197[a]	-4.204	-0.180[a]	-3.833				
Δy_{t-1}					-0.291	-2.598		
Δy_{t-2}					-0.298	-2.661		
Δx_t	0.220	1.793	0.295	1.560	0.165	0.921	0.187	0.425
Δx_{t-1}							-0.584	-1.294
y_{t-1}	-0.077	-2.304	-0.097	-2.248	-0.274	-2.891	-1.041	-9.517
x_{t-1}	0.124	4.650	0.131	3.714	0.263	3.179	0.705	3.400
Long-run elast.	1.610		1.351		0.960		0.677	
Variance ratio and diagnostics:								
q[b]	0		0		0.258		6.333	
Std. Error	0.042		0.043		0.060		0.061	
Normality[c]	0.127		0.691		1.144		0.352	
H[d]	1.510		1.618		1.502		1.584	
DW	2.217		2.122		1.934		1.992	
Q[e]	3.986		3.844		1.903		3.427	
R^2	0.391		0.372		0.183		0.123	

[a] The included dummy-variable is DI02.4.

[b] q is defined as $\sigma^2_\eta / \sigma^2_\varepsilon$.

[c] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[d] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator is 29.

[e] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

Table B3. Econometric relations for oilrig activity in Asia Pacific and Non-OPEC Middle East. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	Asia Pacific (12 months)		Asia Pacific (36 months)		Non-OPEC Middle East (12 months)		Non-OPEC Middle East (30 monhs)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Δx_t	-0.030	-0.047	0.012	0.005	1.171	1.769	3.801	2.243
Δx_{t-1}	-0.0740	-1.160	0.944	0.386				
$\Delta \mathrm{x}_{\mathrm{t}-2}$			-3.660	-1.545				
y_{t-1}	-0.519	-5.585	-1.038	-9.523	-1.221	-11.630	-1.230	-11.940
x_{t-1}	0.088	1.215	0.994	1.764	0.489	2.239	0.784	1.709
Long-run elast.	0.170		0.958		0.400		0.637	
Variance ratio and diagnostics:								
q[a]	0.106		6.203		7.120		10.163	
Std. Error	0.050		0.051		0.054		0.054	
Normality[b]	0.120		0.301		5.391		3.996	
H[c]	0.614		0.543		0.707		0.677	
DW	1.931		1.994		2.016		2.016	
Q[d]	4.290		6.745		5.381		5.413	
R^2	0.108		0.074		0.157		0.158	

[a] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[b] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[c] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator is 29.

[d] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

Table B4. Econometric relations for oilrig activity in Africa and Non-OPEC Africa. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	Africa (24 months)		Africa (36 months)		Non-OPEC Africa (12 months)		Non-OPEC Africa (36 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable	0.644[a]	6.225[a]	0.556	4.992				
Δx_t	2.614	1.233	2.183	0.483	-0.043	-0.345	0.052	0.434
Δx_{t-1}	-2.892	-1.373	3.971	0.871				
y_{t-1}	-1.139	-12.201	-1.149	-11.902	-0.803	-7.578	-0.792	-7.521
x_{t-1}	1.625	3.510	1.635	1.725	-0.178	-0.175	-0.008	-0.080
Long-run elast.	1.427		1.423		-0.222		-0.010	
Variance ratio and diagnostics:								
q[a]	1.260		1.777		0.205		0.188	
Std. Error	0.103		0.108		0.478		0.477	
Normality[b]	10.072		10.905		27.869		27.183	
H[c]	0.695		0.691		0.783		0.772	
DW	2.135		2.123		1.966		1.972	
Q[d]	7.569		12.389		4.656		4.627	
R^2	0.405		0.352		0.171		0.173	

[a] The included dummy-variable is DS02.2.

[b] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[c] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[d] This is a Fisher-distributed test statitic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator are 29.

[e] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

Table B5. Econometric relations for oilrig activity in Norway. Smoothing assumption with regard to oil prices in parentheses

Slope coefficients related to indicated variables	Norway (6 months)		Norway (12 months)		Norway (24 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable	-0.652[a]	-4.401	-0.695[a]	-5.285	-0.681[a]	-4.807
Δx_t	0.557	0.866	-0.080	-2.164	0.039	1.032
Δx_{t-2}	-1.241	-1.792				
y_{t-1}	-1.009	-10.142	-1.233	-14.633	-1.237	-13.820
x_{t-1}	0.378	2.507	0.010	0.223	0.015	0.354
Long-run elast.	0.375		0.008		0.012	
Variance ratio and diagnostics:						
q[b]	0.059		2.114		1.054	
Std. Error	0.148		0.158		0.164	
Normality[c]	3.522		2.084		2.122	
H[d]	0.684		0.874		0.687	
DW	1.850		1.779		1.757	
Q[e]	7.625		5.778		5.312	
R^2	0.543		0.559		0.525	

[a] The included dummy-variable is DI96.6.

[b] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[c] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[d] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator is 29.

[e] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

Table B6. Econometric relations for oilrig activity in UK. Smoothing assumption with regard to oil prices in parentheses[a]

Slope coefficients related to indicated variables	UK (6 months)		UK (12 months)		UK (24 months)	
	Estimate	t-value	Estimate	t-value	Estimate	t-value
Dummy-variable 1			-0.431[b]	-3.317	-0.335[b]	-2.692
Dummy-variable 2			-0.733[c]	-5.740	-0.728[c]	-5.792
Δy_{t-1}	-0.477	-4.115	-0.158	-1.755		
Δy_{t-2}	-0.213	-2.067				
Δx_t	0.942	1.307	1.557	1.208	1.835	0.705
Δx_{t-2}	-2.707	-3.391	-4.264	-3.336		
y_{t-1}	-0.302	-3.227	-0.764	-5.829	-1.091	-11.774
x_{t-1}	0.381	3.1962	0.640	2.239	1.533	2.482
Long-run elast.	1.262		0.838		1.405	
Variance ratio and diagnostics:						
q[d]	0.866		0.447		1.361	
Std. Error	0.145		0.118		0.126	
Normality[e]	7.934		3.025		7.451	
H[f]	1.142		1.113		1.527	
DW	1.945		1.969		1.930	
Q[g]	6.359		4.774		10.002	
R_S^2[h]	0.778		0.852		0.832	

[a] Estimates of fixed seasonal effects are not reported.

[b] The dummy-variable is DI99.7.

[c] The dummy-variable is DI99.10.

[d] q is defined as $\sigma_\eta^2 / \sigma_\varepsilon^2$.

[e] This is the Doornik and Hansen (1994) adjusted version of the Bowman-Shenton (1975) statistic for normality, which has a χ^2 distribution with 2 degrees of freedom.

[f] This is a Fisher-distributed test statistic for heteroskedasticity, cf. Koopman et al. (1999, p. 119). The degrees of freedom in both the numerator and denominator is 29.

[g] This is the Box-Ljung test statistic for autocorrelation based on the first p autocorrelations. It is χ^2-distributed with q degrees of freedom where q is p+1 less the number of estimated variance components. p is set to 8.

[h] For definition of R_S^2 cf. Koopman et al. (1999, p. 180).

Calculation of immediate, intermediate and long-run elasticities

As an example we consider the case with log-prices smoothed over six months and where in addition to the long-run terms we have included two lags of the left-hand side variables and the current and one period lagged relative change of (smoothed) oil prices. Thus this corresponds to the model used for US with smoothed prices over six months. The model is given as

$$(C.1)\Delta y_t = \beta_1 \Delta y_{t-1} + \beta_2 \Delta y_{t-2} + \gamma_0 \Delta x_t + \gamma_1 \Delta x_{t-1} + \lambda y_{t-1} + \theta x_{t-1} + \text{"}other\,terms\text{"},$$

where

$$(C.2)\, x_t = \frac{1}{6}\sum_{j=0}^{5} p_{t-j} \,.$$

In Eq. (C.2) p_t denotes the log of the non-smoothed price in period t. Inserting from (C.2) in (C.1) and reparameterising such that we can write (C.1) in levels we obtain

$$(C.3)\, y_t = \sum_{i=1}^{3} \xi_i y_{t-i} + \sum_{i=0}^{7} \zeta_i p_{t-i} + \text{"}other\,terms\text{"},$$

where

$$(C.4a)\, \xi_1 = 1 + \beta_1 + \lambda,$$

$$(C.4b)\, \xi_2 = \beta_2 - \beta_1,$$

$$(C.4c)\, \xi_2 = \beta_2 - \beta_1,$$

$$(C.4d)\, \zeta_0 = (1/6)\gamma_0,$$

$$(C.4e)\, \zeta_1 = (1/6)(\theta + \gamma_1),$$

$$(C.4f)\, \zeta_i = (1/6)\theta; i = 2,3,4\, and\, 5,$$

$$(C.4g)\, \zeta_6 = (1/6)(\theta - \gamma_0) \text{ and}$$

$$(C.4h)\, \zeta_7 = -(1/6)\gamma_1.$$

Let the two 8×1 vectors $\mathrm{y}_{\mathrm{t}}^{*}$ and $\mathrm{p}_{\mathrm{t}}^{*}$ be defined as

$$(C.5)\, y_t^* = \left[y_t,\ \ y_{t-1},\ \ y_{t-2},\ \ y_{t-3},\ \ y_{t-4},\ \ y_{t-5},\ \ y_{t-6},\ \ y_{t-7} \right]' \text{ and}$$

$$(C.6)\, p_t^* = \left[p_t,\ \ p_{t-1},\ \ p_{t-2},\ \ p_{t-3},\ \ p_{t-4},\ \ p_{t-5},\ \ p_{t-6},\ \ p_{t-7} \right]'.$$

Model (3) may now be represented by

$$(C.7)\, y_t^* = H y_{t-1}^* + K p_t^* + D_t,$$

where

$$(C.8)\, H = \begin{bmatrix} \xi_1 & \xi_2 & \xi_3 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix},$$

$$(C.9)\, K = \begin{bmatrix} \zeta_0 & \zeta_1 & \zeta_2 & \zeta_4 & \zeta_4 & \zeta_5 & \xi_6 & \zeta_7 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}$$

and D_t is an 8×1 vector with the first element corresponding to the "other terms" in (C.3) and where all the other terms are equal to zero. The elasticity after j periods (j=0,1,…) is now given by

$$(C.10)\, E_j = tr\left[\sum_{i=0}^{j} F_i\right] = \sum_{i=0}^{j} tr(F_i),$$

where tr denotes the trace-operator and where

$$(C.11)\; F_i = H^i K, i = 0,1,2,...$$

When j goes to infinity we obtain the long-run elasticity, which also may be obtained directly from (C.1) by utilizing the Bårdsen (1989) formula

$$(C.12)\; E_\infty = \lim_{j \to \infty} E_i = -\theta / \lambda.$$

The derivation when considering models in which the log oil-price is smoothed over a different number of periods and where the lag order differs from those used in conjunction with (C.1) are analogous to the one considered in this appendix.

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