

*Andrzej Geise, Mariola Pilatowska**

Synchronization of Crude Oil Prices Cycle and Business Cycle for the Central Eastern European Economies**

A b s t r a c t. The main purpose of the paper is to study the degree to which the Brent crude oil price cycle is correlated and synchronized with business cycle in a set of chosen Central Eastern European (CEE) economies. To identify the oil price cycle and business cycles for chosen individual countries the Markov-switching autoregressive model (MS-AR) is used. The identification of the smoothed probabilities of being in regime 1 and regime 2 enables the calculation of correlation coefficients between those probabilities and the concordance index to evaluate the synchronization of oil price cycle and business cycles for the CEE economies.

K e y w o r d s: Markov switching model, crude oil prices, business cycle, price cycle.

J E L Classification: C22, C32, E23, Q40.

Introduction

The energy prices have the significant impact on the economic and civilization development. Rising energy prices observed in recent years accelerated the realization of well-balanced energy consumption programs and the reduction of global energy intensity. Energy prices can be affected not only by economic factors but to high extent also by unpredictable and non-

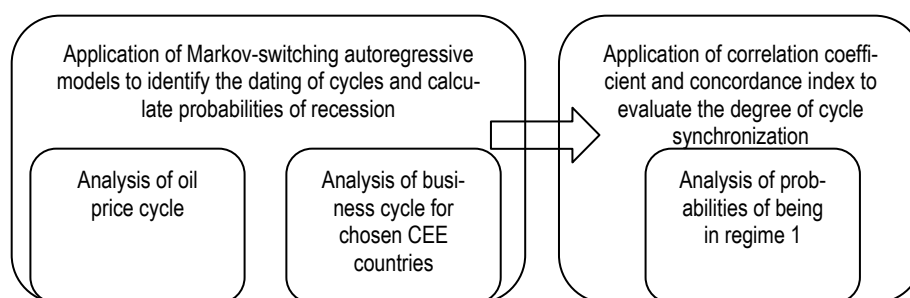
* Correspondence to: Andrzej Geise, Department of Econometrics and Statistics, 13A Gagarina Street, 87-100 Toruń, Poland, e-mail: a.geise@doktorant.umk.pl, Mariola Pilatowska, Department of Econometrics and Statistics, 13A Gagarina Street, 87-100 Toruń, Poland, e-mail: mariola.pilatowska@umk.pl.

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economic events, such as wars, political situation in regions of fossil fuels mining or devastating weather. Specifically, the ecology and the international environmental protection should be taken into account when having in mind the future development of fossil fuel market. Special significance in the energy consumption structure has the crude oil which is treated as a major determinant of world economic activity. The economic and social significance of crude oil did not diminish along with the development of alternative energy source, like wind, water and sun energy. Irrespective of the potentials of renewable energy the impact of crude oil on the economy is still strong. Therefore, it is expected that the relation between economic growth and crude oil prices volatility exists.

The main purpose of the paper is to study the degree to which the crude oil price cycle is correlated and synchronized with business cycle in Czech Republic, Hungary, Poland and Slovenia. First, the Markov-switching autoregressive model (MS-AR) is used to identify oil price cycle and the business cycle for chosen individual countries. Next, the correlation coefficients between smoothed probabilities of a recession for business cycle in given countries and crude oil price cycle are calculated in order to evaluate the extent to which turning points in the two series occur near each other. Finally, the concordance index is computed to reflect the degree to which two series are in the same state or in other words are synchronized. The concept of a research study is presented in Scheme 1.

Scheme 1. The concept of studying the synchronization of oil price cycle and business cycle



In the analysis the monthly data of Brent crude oil prices from January 1995 to April 2013 (220 observations) and industry production index for a set of Central Eastern European (CEE) economies from January 1995 to April

2013 were used¹. The data were taken from the U.S. Energy Information Administration database (www.eia.gov) and the OECD database (www.stats.oecd.org).

The analysis of relations between crude oil prices and economic activity is an important issue. The volatility of crude oil prices has become one of the most important indicator of economic growth due to the significant share of crude oil in the energy consumption structure. Moreover, the government policy is driven by the crude oil market as the tax revenues (e.g. oil excise, oil VAT, oil fee) from oil sector are significant. Table 1 is as an example of the significance of oil sector to the Polish budget.

Table 1. Tax revenues from oil sector in Poland in 2012

Tax categories	Tax category revenue as percentage of total oil tax revenue (PLN 52bn)	Oil tax as percentage of total tax category ²
Oil excise	51.93%	41.00%
Oil VAT	40.38%	15.85%
Oil fee	7.69%	1.39%

Source: elaborated on the base of POPiHN, *Przemysł i handel naftowy – raport roczny 2012 (Industry and Oil trade – annual report 2012)*, Warszawa, 2013.

Tax revenues from oil sector are an important part of state income in Poland (in 2012 tax revenue reached ca. PLN 52 billion. The structure of taxes coming from oil sector was dominated by an oil excise and oil VAT which constituted respectively: 41% of total excise and 16% of total VAT. This confirms that the oil sector is of importance to the budget of Poland. Consequently, the relation between crude oil prices volatility and business cycle is expected. Therefore, the analysis directed to evaluate the degree to which the crude oil price cycle is correlated and synchronized with business cycle is needed.

1. The Relationship between Crude Oil Prices and Economic Activity – the Literature Review

The economic theory suggests that assets prices should be determined by their expected, discounted cash flows. Then, every factor influencing the discounted cash flows should have a significant impact on assets prices (Fisher, 1930). Expanding this rule to the crude oil prices we have that every

¹ The Brent crude oil constitutes the global benchmark of crude oil coming from North Eastern Europe and reflects changes in crude oil prices in the European region.

² E.g. oil excise as percentage of total excise, oil VAT as percentage of total VAT, etc.

rise in crude oil prices entails the increase of production costs which causes the profits and capital value to drop. Hence, every rise of oil prices brings on the decrease of stock prices. This effect should have occurred both for crude oil exporters and importers. However, many economists discuss whether the influence of oil price dynamics on the stock price dynamics results from an indirect effect which is escalated by the impact of macroeconomic indicators. According to Bjornland (2008) and Jimenez-Rodriguez, Sanchez (2005) the rise of oil prices has a positive effect on domestic income in oil exporting countries. Also, the increase in expenditure and investment is expected which as a result will cause the increase of productivity and reduction of unemployment. By contrast, in oil importing countries every rise in oil prices will bring on the opposite effect, in the form of decrease in income (Bacon, 2005; Hooker, 2002). Rising oil prices lead to an increase in production costs in oil-importing countries because the crude oil is the important input in the production process (Arouri, Nguyen, 2010; Backus, Crucini, 2000; Kim, Loungani, 1992).

The increase in costs is transferred to consumers by increasing consumer prices, and this leads to the reduction of demand and consumer expenditure. The decrease in demand may result in a decrease in production, and hence in an increase in unemployment. Consequently, the reaction of stock market will be negative. However, it should be remembered that the influence of oil shocks on stock market depends on the type of shock (demand or supply shock). In the case of demand shock the stock market is supposed to react positively, and in the case of supply shock – negatively (Hamilton, 1988, 2006; Abel, Bernanke, 2001; Brown, Yucel, 1999, 2002). Summing up, rising oil prices may lead to positive economic effects (like an increase in domestic income) in oil exporting countries, and negative ones (like a decrease in domestic income) in oil importing countries.

Studying relationship between the oil price volatility and economic activity many economists have suggested that the indirect transmission mechanism may have a fundamental impact in identifying which oil price shocks are important from a macroeconomic point of view. For example, Loungani (1986) and Davis, Haltiwanger (2001) investigated the relationship between oil prices and employment market; Bernanke (1983) and Dixit, Pindyck (1994) analyzed connections of oil prices and uncertainty of investment; Hamilton (1988) and Lee, Ni (2002) studied the ability of oil price volatility to smooth the consumption in durable goods sector; Pierce, Enzler (1974) and Bruno, Sachs (1982) analyzed the impact of oil price volatility on inflation. It should be emphasized that these studies did not decide clearly neither about the impact of crude oil price on economic growth, nor the impact of

economic growth on oil price levels. Therefore, it is important to investigate how the oil shocks are transmitted into economic activity.

2. Univariate Markov Switching Model

In recent years many theoretical and empirical business cycle studies have analyzed the co-movements of macroeconomic time series and the regime-switching nature of macroeconomic activity. As the statistical measurement of macroeconomic fluctuations the Markov-switching autoregressive time series model has become increasingly popular since Hamilton's (1989) application of this technique to measure the US business cycle. In the Markov-switching models the recessions and expansions are modeled as switching regimes of the stochastic process generating the growth rate of economic activity indicator, e.g. GDP (Artis et al., 2004). The general form of switching model for growth rate of gross output Δy_t can be written (Krolzig, Toro, 2005):

$$\Delta y_t - \mu_{s_t} = \alpha_1(\Delta y_{t-1} - \mu_{s_{t-1}}) + \dots + \alpha_q(\Delta y_{t-q} - \mu_{s_{t-q}}) + \varepsilon_t, \quad (1)$$

where $\varepsilon_t \sim N(0, \sigma^2)$ is an independent and identically distributed series with zero mean and finite variance, $s_t \in \{1, \dots, M\}$ – regime variable. The regimes are associated with different conditional distributions of the growth rate of real output, where the mean μ_{s_t} depends on the value of a discrete state variable s_t . For instance, in two-regime model the mean may be negative in the first regime (recession), $\mu_1 < 0$, and positive in the second regime (expansion), $\mu_2 > 0$. The variance of the disturbance term ε_t is assumed to be the same in both regimes (Artis et al., 2004).

The general idea behind the class of regime-switching models is that the parameters of equation (1) depend upon a stochastic, non-observable regime variable $s_t \in \{1, \dots, M\}$. The stochastic process generating the unobservable regimes is an ergodic Markov chain defined by the transition probabilities (Artis et al., 2004; Kořko, 2006; Stawicki, 2004):

$$p_{ij} = \Pr(s_{t+1} = j | s_t = i), \quad \sum_{j=1}^M p_{ij} = 1 \quad \forall i, j \in \{1, \dots, M\}. \quad (2)$$

To highlight the stylized fact of growth output, i.e. greater volatility in recessions than expansions, the approach with time-dependence in the variances of the growth output depending on the regime can be used.

In this paper the Markov two regime-switching autoregressive model with regime-dependent intercepts, variances and autoregressive parameters will be applied.

3. Brent Crude Oil Price Cycle in 1995–2013 Period

The data correspond to monthly Brent crude oil prices from January 1995 to April 2013 (220 observations in US dollars per barrel) and were drawn from the U.S. Energy Information Administration database. Based on these data the two-regime Markov switching model was estimated where regime 1 stands for the drops in crude oil prices and regime 2 stands for the rises in crude oil prices.

We consider a two-regime switching Markov model where the logs of crude oil prices are modeled in first differences:

$$\Delta c_t = \mu_{s_t} + \sum_{j=1}^q \gamma_{j,s_t} \Delta c_{t-j} + \xi_t, \quad \xi_t | s_t \sim iid(0, \sigma_{s_t}^2), \quad (3)$$

where Δc_t stands for the log first differences of Brent crude oil prices³, $\mu_{s_t}, \gamma_{j,s_t}, \sigma_{s_t}^2$ – denote regime-dependent intercept, autoregressive parameters and variance of disturbance term ξ_t respectively. The unobserved variable $s_t = 1$ in the first regime (drop in oil prices) and $s_t = 2$ in the second regime (rise in oil prices).

The Markov-switching autoregressive models allowing the switches between regimes in different parameters according to a hidden Markov chain were chosen based on the Akaike Information Criterion (AIC) and log-likelihood function (LL). We considered four types of Markov switching models introducing switches in the intercept (MSI), in intercept and variance (MSIH), in intercept and autoregressive parameters (MSIA) and in intercept, variance and autoregressive parameters (MSIHA). The value of AIC and LogLik for different MS-AR models⁴ are given in Table 2.

³ To test the non-stationarity of the logs of Brent crude oil prices the ADF test was used. The ADF statistic for log-levels amounts 0.85 ($p=0.89$) – without intercept, and -3.6 ($p=0.03$) – with intercept and trend, and for log-first differences: -6.76 ($p=0.00$) – without intercept, and -6.92 ($p=0.00$) – with intercept and trend. This indicates that the log-levels of crude oil prices are non-stationary and become stationary when they are first differenced. Therefore the log-differences of crude oil prices are used in further analysis.

⁴ In this paper we focus on the maximum likelihood estimation of parameters. All the calculations in the paper were carried out in OxMetrics6.

Table 2. The Akaike Information Criterion (AIC) and log-likelihood function (LL) for the univariate Markov switching model for the crude oil prices, January 1995 to April 2013

Markov switching model	AIC	LL
MSI(2)-AR(1)	-674.27	343.13
MSIH(2)-AR(1)	-672.48	343.24
MSIA(2)-AR(1)	-687.23	350.61
MSIHA(2)-AR(1)	-678.62	347.31
MSI(2)-AR(2)	-715.86	364.86
MSIH(2)-AR(2)	-722.10	369.05
MSIA(2)-AR(2)	-716.98	368.49
MSIHA(2)-AR(2)	-719.35	368.57

Table 3. Univariate MSIH(2)-AR(2) model for Brent crude oil prices

Parameters	Parameters estimates	p-value
Regime-dependent intercepts		
μ_1	-0.0063 (0.0149)	0.102
μ_2	0.0156 (0.0069)	0.004
Autoregressive parameters		
γ_1	0.8678 (0.0629)	0.000
γ_2	-0.4611 (0.0627)	0.000
Regime-dependent variances		
σ^2_1	0.0645 (0.0083)	0.000
σ^2_2	0.0359 (0.0033)	0.000
LL	369.05	
AIC	-722.1	
LR test	15.63	0.0036
Jarque-Bera test	0.8013	0.6699
RCM	37.99	
Transition probabilities across regimes		
	p_{1i}	p_{2i}
Regime 1	0.9321	0.0292
Regime 2	0.0679	0.9708

Note: in parentheses the standard errors of estimates are given.

Using the AIC and the LL function the Markov switching model with regime-dependent intercept and variance (MSIH(2)-AR(2)) was found to be the best. The detailed estimation results for the period January 1995 to April 2013 are displayed in Table 3.

The transition matrix allows us to observe the asymmetry of oil price cycle in terms of the duration of drops (regime 1) and rises (regime 2) in oil prices. Drops have the duration of approximately 15 months and rises have an average duration of two times the regime 1 (34 months). Both states are characterized by high probability of remaining in regime ($p_{11} = 0.93$, $p_{22} = 0.97$) – see Table 3 and 4. In the regime 1 the oil prices tend to decrease ($\mu_1 < 0$) while in regime 2 they tend to increase ($\mu_2 > 0$); the volatility of oil prices is higher in the regime 1 (drops in oil prices) than in regime 2 (rises in oil prices) ($\sigma_1^2 > \sigma_2^2$) – see Table 3. The RCM (Regime Classification Measure⁵) statistic indicates that the MSIH(2)-AR(2) model is able to confidently distinguish which regimes are occurring at each point in time ($RCM < 50$).

In the studied period the first serious oil prices fall was observed in 1997–1999 period (Table 4 and Figure 1) which may be connected with financial (currency) crisis in South Eastern Asia and then in Russia. The situation in Russia has made the market react by falling oil prices because Russia is the greatest crude oil producer (with average daily production about 10 millions of barrels; see International Energy Agency, 2013). The next two oil prices falls in 2001 and 2008–2009 period may be influenced by the economic crises in the U.S. (the third largest oil producing and the largest oil consuming country in the world).

The decrease in oil prices in 2001 corresponds to the recession in the U.S. that occurred in March 2000 when the NASDAQ crashed following the collapse of the Dot-com bubble (Figure 1). Consequently, the decline of stock prices occurred in many countries (Galbraith, Hale, 2004). The second decline in oil prices, i.e. in 2008–2009 period, is connected with the sub-prime mortgage crisis that broke out in the latter half of 2007. This financial crisis spread to the oil market (and other commodity markets) which is driven by global changes in supply and demand along with a number of other geopolitical factors.

⁵ The RCM statistic was proposed by Ang, Bekaert (2002). For two regimes it takes the form: $RCM = 400 \frac{1}{T} \sum_{t=1}^T p_t(1-p_t)$, where p_t is a smoothed probability of being in a certain regime at time t . The constant serves to normalize the statistic to be between 0 and 100. The RCM is a summary point statistic which describes the quality of regime classifications. An ideal model is that classifying regimes sharply and having smoothed probabilities which are either close to zero or one. If the RCM is close to zero, the regime classification is perfect whereas a value of 100 means that no information about regimes is revealed. The cut-off value of RCM statistic is 50 which is often used as a benchmark (Chan, et al., 2011).

Table 4. Dating of crude oil price cycle

Regime 1	Regime 2
	1995(04)–1997(10)
1997(11)–2001(11)	2001(12)–2006(07)
2006(08)–2006(12)	2007(01)–2008(06)
2008(07)–2009(02)	2009(03)–2013(04)
62 months (28.57%) out of 217 are in regime 1	155 months (71.43%) out of 217 are in regime 2
Average duration $d(i)$ 14.73 months	2 Average duration $d(i)$ 34.24 months

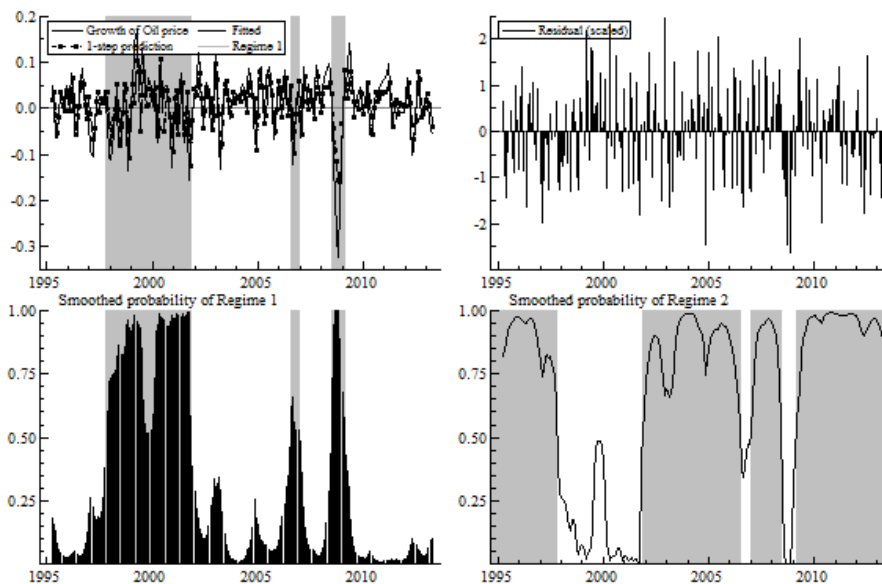


Figure 1. Smoothed probabilities of being in regime 1 and regime 2 for the univariate Markov switching model for the crude oil prices, January 1995 to April 2013

The oil price decline in 2006 is difficult to explain. In the second half of 2006 the oil price decline indeed occurred but it was rather small, however the MS-AR model captured it as a recession. Hence, this behavior of oil prices may be rather combined with general economic situation of oil exporters and importers and uncertainty on financial markets (although there were some inflammatory events in 2006, like the attack on Shell pipeline in Nigeria, worries over Iranian nuclear plans or Israel's War against Hezbollah).

In the next section the business cycle of chosen Central Eastern European economies (Poland, Czech Republic, Hungary, Slovenia) is identified.

4. Identification of Business Cycle in Central Eastern European (CEE) Economies in 1995–2013 Period

The data consists of seasonally adjusted⁶ monthly industry production index for a set of CEE economies from January 1995 to April 2013. These monthly series were taken from the OECD database. Our analysis is restricted to a subset of four CEE countries: Czech Republic, Hungary, Poland, Slovenia⁷. Log-transformed data along with seasonally adjusted data are displayed in Figure 2. It is seen that the economies experienced the trend break in 2008 referring to the raging financial crisis (subprime mortgage crisis). Specifically, the serious decline in the industry production index is apparent in the case of Czech Republic, Hungary and Slovenia. In the case of Poland the impact of financial crisis on the industry production index was not so evident.

Table 5. The ADF test results for industry production index (IPI) in a set of CEE countries

Time series	Lag	ADF test for IPI		ADF test for Δ IPI	
		without intercept	with intercept and trend	without intercept	with intercept and trend
Czech Republic	5	1.3983	-2.7621	-4.4468	-4.7682
		(0.96)	(0.2116)	(0.0000)	(0.0005)
Hungary	5	2.3805	-0.8783	-4.5967	-9.7907
		(0.9962)	(0.9569)	(0.0000)	(0.0000)
Poland	5	3.9040	-2.7320	-3.2319	-8.7368
		(0.9999)	(0.2235)	(0.0012)	(0.0000)
Slovenia	5	1.0068	-1.9891	-6.3661	-6.4719
		(0.9178)	(0.6067)	(0.0000)	(0.0000)

Note: in parentheses the p -values are given.

To test the non-stationarity of series the ADF test was applied. Table 5 reports the results of the ADF unit root test for the logs of industry production index (IPI) for levels and the first differences. For all IPI series in levels the null hypothesis of a unit root cannot be rejected (p -value $>$ $\alpha = 0.05$). The ADF results for first differences, Δ IPI, allow to reject the null hypothe-

⁶ For seasonal adjustment the TRAMO/SEATS procedure was used.

⁷ The intention of a research was to include all CEE economies. However, the final choice of countries was determined by the availability of data, i.e. sufficient length of time series.

sis $\Delta IPI \sim I(1)$ (p -value $< \alpha = 0.05$). The results of the ADF test indicate that all industry production indexes under consideration are non-stationary in their levels and become stationary when they are first differenced. Hence, in further analysis the first differences of IPI for given CEE economies were used.

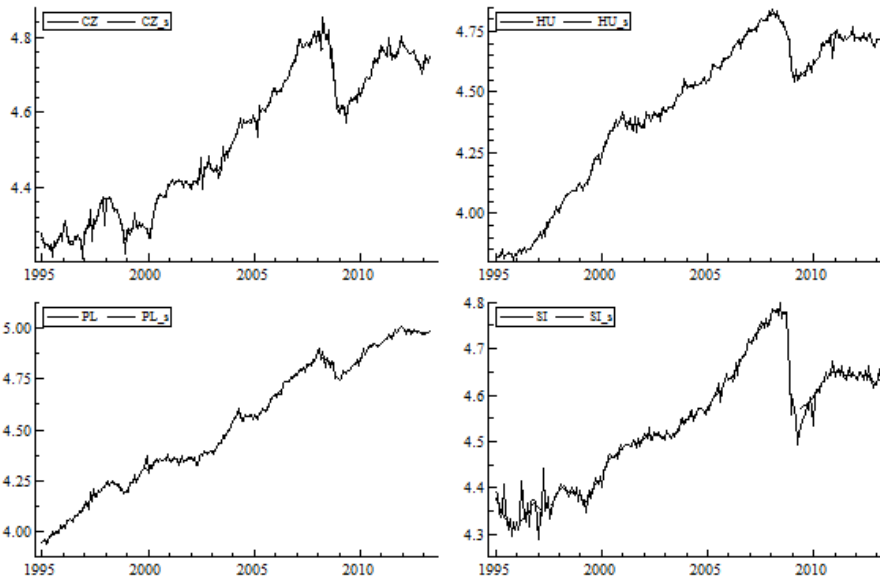


Figure 2. Industry production index in the CEE countries

To identify the business cycle of different CEE countries we applied the Markov switching model. The two-regime Markov switching model for industry production index modeled in first differences takes the form:

$$\Delta y_t = \mu_{s_t} + \sum_{j=1}^q \phi_{j,s_t} \Delta y_{t-j} + \varepsilon_t, \quad \varepsilon_t | s_t \sim iid(0, \sigma_{s_t}), \quad (4)$$

where Δy_t stands for first differences of the natural log of industry production index for a given country (economy), μ_{s_t} , ϕ_{j,s_t} , $\sigma_{s_t}^2$ – denote regime-dependent intercept, autoregressive parameters and variance of disturbance term ε_t respectively. The unobserved variable $s_t = 1$ in the first regime (recession) and $s_t = 2$ in the second regime (expansion).

Like in the case of the MS-AR models for crude oil price, four type of Markov switching models were considered, i.e. MSI, MSIH, MSIA, MSIHA. The value of AIC and LL for different MS-AR models are reported in Table 6.

Table 6. The Akaike Information Criterion (AIC) and log-likelihood function (LL) for the univariate Markov switching model for industry production index

Markov switching model	Czech Republic		Poland		Hungary		Slovenia	
	LL	AIC	LL	AIC	LL	AIC	LL	AIC
MSI(2)-AR(1)	774.99	-1537.9	849.35	-1686.7	798.03	-1584.1	1055.4	-2098.9
MSIH(2)-AR(1)	778.55	-1543.1	850.91	-1687.8	801.15	-1588.2	1056.0	-2088.0
MSIA(2)-AR(1)	780.04	-1546.1	852.42	-1690.9	802.51	-1591.0	1055.6	-2097.2
MSIHA(2)-AR(1)	784.92	1553.85	852.53	-1689.1	808.31	-1602.6	1061.1	-2106.2
MSI(2)-AR(2)	804.30	-1594.6	876.61	-1739.2	807.19	-1600.4	1078.2	-2142.3
MSIH(2)-AR(2)	811.06	-1606.1	879.28	-1742.6	822.95	-1629.9	1088.6	-2161.2
MSIA(2)-AR(2)	814.48	-1611.6	893.47	-1768.9	828.67	-1639.3	1086.0	-2153.9
MSIHA(2)-AR(2)	815.77	-1611.6	886.97	-1753.9	824.23	-1628.5	1097.5	-2174.9

The AIC criterion and LL function found the MSIHA(2)-AR(2) model to be the best for Czech Republic and Slovenia, and the MSIA(2)-AR(2) model – for Poland and Hungary (Table 6). The detailed estimation results for these models for the period January 1995 to April 2013 are displayed in Table 7.

While the parameter estimation is successfully made by the Markov switching models, it is worthwhile to make a more formal assessment of the performance of these models. Therefore, the formal testing for nonlinearity in the context of Markov switching models is carried out. The results of the Likelihood Ratio (LR) test indicate that the significant evidence of nonlinear behavior in series under consideration has been found (see Table 7). Therefore the Markov switching model are considered to be attractive in addressing the nonlinear nature of studied series (especially their asymmetric behavior).

The state of low economic activity (regime 1) is highly persistent in case of Poland ($p_{11} = 0.93$) and relatively persistent in case of Czech Republic and Hungary ($p_{11} = 0.73$ and 0.70 respectively) – see Table 7, with exception of Slovenia for which the probability of remaining in regime 1 is equal $p_{11} = 0.5$ what along with $p_{12} = 0.49$ may indicate a poor model specifica-

Table 7. The univariate Markov switching model for industry production index for chosen CEE countries

Parameters	Czech Rep.	Poland	Hungary	Slovenia
	MSIHA(2)–AR(2)	Type of Markov switching model MSIA(2)–AR(2)		MSIHA(2)–AR(2)
Regime-dependent intercept				
μ_1	–0.0036 (0.0012)	–0.0014 (0.0005)	–0.0011 (0.0011)	–0.0017 (0.001)
μ_2	0.0037 (0.001)	0.0051 (0.0006)	0.0045 (0.0009)	0.0025 (0.0005)
Regime-dependent autoregressive parameters				
ϕ_{1-1}	1.2943 (0.0827)	0.7356 (0.0974)	1.0516 (0.1199)	1.0107 (0.1772)
ϕ_{1-2}	0.8437 (0.0894)	0.9876 (0.0754)	0.8048 (0.0799)	1.3040 (0.0588)
ϕ_{2-1}	–0.2759 (0.0951)	–0.4848 (0.0829)	–0.0706 (0.0527)	–0.8902 (0.1948)
ϕ_{2-2}	–0.6406 (0.0698)	–0.6155 (0.0739)	–0.6188 (0.0928)	–0.4957 (0.0567)
Regime-dependent variances				
σ^2_1	0.0056 (0.0005)	0.0034 (0.0002)	0.0049 (0.0003)	0.005 (0.0015)
σ^2_2	0.0042 (0.0005)	–	–	0.0014 (0.0001)
LL	815.77	893.47	828.67	1097.5
AIC	–1611.6	–1768.9	–1639.3	–2174.9
LR test	22.95 ($p=0.0008$)	33.72 ($p=0.000$)	31.28 ($p=0.000$)	34.47 ($p=0.000$)
Jarque–Bera test	5.549 ($p=0.0624$)	3.360 ($p=0.186$)	14.84 ($p=0.0006$)	3.516 ($p=0.172$)
RCM	62.28	10.16	58.37	6.66
Transition probabilities				
p_{11}	0.7309	0.9205	0.7026	0.5005
p_{22}	0.8094	0.9619	0.8648	0.9782
p_{12}	0.2691	0.0795	0.2974	0.4995
p_{21}	0.1906	0.0381	0.1352	0.0218
Average duration of regime				
$d(i)_1$	3.72	12.58	3.36	2.00
$d(i)_2$	5.25	26.25	7.40	45.87

Note: in parentheses the standard errors of parameter estimates are given.

tion⁸. What is interesting the RCM statistic (see Table 7, RCM=6.66 is substantially below 50) seems to indicate something opposite, i.e. that MSIHA(2)-AR(2) model for Slovenia performs well in distinguishing between two regimes. However, it should be noted that a low value of the RCM statistic does not necessarily imply that the switches are correctly predicted, but only that they are sharp, with a small frequency of periods of uncertainty on the nature of regimes (Guidolin, 2011). The regime 2 (expansion) is highly persistent for all countries. In the recessionary periods the industry production index tends to decrease ($\mu_1 < 0$) whereas in the expansionary periods – tends to increase ($\mu_2 > 0$) for all considered countries. The volatility of industry production index is higher in the regime 1 (recession) than in regime 2 (expansion) for Czech Republic and Slovenia ($\sigma_1^2 > \sigma_2^2$). The duration of recessions and expansions is asymmetric for all countries, e.g. the duration of expansions is about twice as long as the recession state (except Slovenia) – see Table 7.

Smoothed probabilities of being in a given regime are calculated (see Figure 3) and every observation is assigned to a given regime according to the highest smoothed probability. For the case of two regimes, the rule reduces to assigning the observation to the first regime if $\Pr(s_t = 1 | Y_T) > 0.5$ and assigning it to the second regime if $\Pr(s_t = 1 | Y_T) < 0.5$. Corresponding dating of business cycle in studied countries is given in Table 8.

According to the classical methodology of NBER (National Bureau of Economic Research) the U.S. economic recession occurred in 2001 (collapse of the Dot-com bubble, March 2001 to December 2001) and in 2007–2009 period (subprime mortgage crisis, December 2007 to July 2009)⁹. The smoothed probabilities of being in a recession (Figure 3) obtained with the Markov switching models indicate that in case of Hungary and Poland the financial crisis affected these economies. The impact of the dot-com crisis has been demonstrated in the case of Hungary and Poland (see Figure 3 and Table 8). Generally, dating of business cycle obtained with the MS-AR models is partly consistent with the dating of world economic recession ob-

⁸ In order to find a better specification with more plausible values of transition probabilities we tried out different model specifications (see Table 6). Only for the MSIA(2)-AR(1) model the reasonable values of p_{11} and p_{12} were obtained, i.e. 0.74 and 0.26 respectively. However, simultaneously the RCM statistic took the higher value, RCM=36.06, and what it is more important the switches between regimes occurred very often what indicates the difficulties in distinguishing between two regimes. Therefore the results for Slovenia seem rather implausible and should be taken with caution.

⁹ See: <http://www.nber.org/-cycles/cyclesmain.html>.

tained with the NBER methodology, however there are other recessionary periods which are country specific, e.g. Poland.

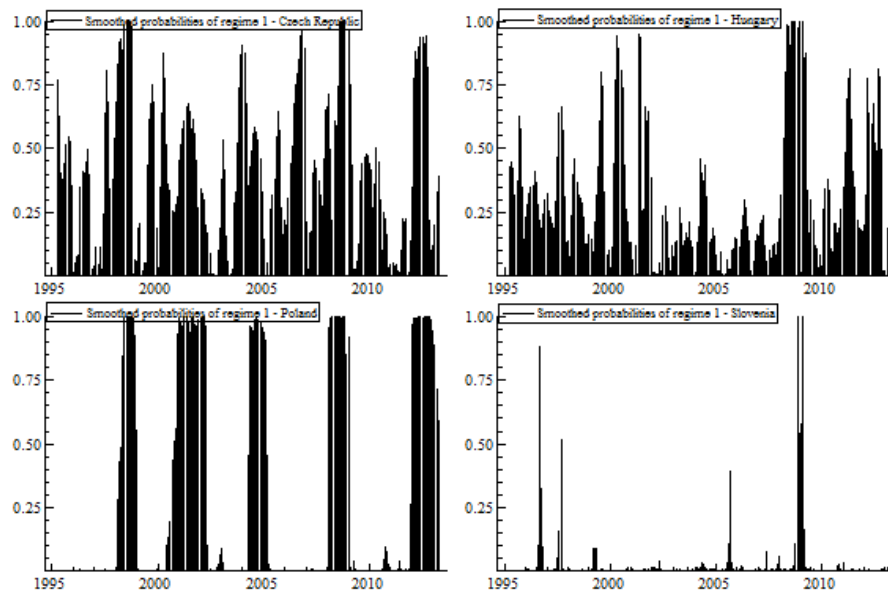


Figure 3. Smoothed probabilities of being in a recession for industry production index in chosen CEE countries

Whereas for Poland the univariate MS-AR model seems to capture relatively well the different recessionary periods, in the case of Czech Republic and Hungary the MS-AR model delivers the worst fit, with difficulties distinguishing clearly the recessionary periods (switches between regimes occur very often). Besides, the RCM statistic (higher than 50, see Table 7) in the case of Czech Republic and Hungary suggests rather that the chosen Markov-switching models are not able to distinguish between two regimes and as a result the existence of two regimes is doubtful (additionally, the model evaluation is weakened by the fact that the standardized residuals are not standard normally distributed – see Table 7). Only in the case of Poland the Markov-switching model can clearly differentiate one regime in each period considered (RCM statistic substantially below 50 and normal distribution of residuals – see Table 7).

Table 8. Dating of business cycle (recessions) for considered Central Eastern European economies and crude oil price cycle (drops in oil prices), 1995–2013

Oil price cycle (drops in oil prices)	Business cycle (regime 1 – recession)			
	Czech Rep.	Poland	Hungary	Slovenia
1997(11)–2001(11)	1995(4)–1995(5)	1998(5)–1999(1)	1995(9)–1995(10)	1996(9)–1996(9)
2006(08)–2006(12)	1995(9)–1995(11)	2000(10)–	1997(8)–1997(10)	1997(9)–1997(9)
2008(07)–2009(02)	1997(7)–1997(9)	2002(4)	1999(7)–1999(9)	2008(11)–2009(2)
	1997(12)–1998(10)	2004(5)–2005(2)	2000(4)–2000(8)	
	1999(8)–1999(11)	2008(2)–2009(1)	2001(5)–2001(6)	
	2000(3)–2000(6)	2012(1)–2013(4)	2001(9)–2001(11)	
	2001(2)–2001(10)		2008(3)–2008(12)	
	2003(2)–2003(2)		2009(2)–2009(4)	
	2003(10)–2004(3)		2011(3)–2011(6)	
	2004(7)–2004(10)		2012(3)–2012(7)	
	2005(8)–2005(10)		2012(9)–2012(10)	
	2006(5)–2006(12)			
	2007(12)–2008(2)			
	2008(5)–2008(11)			
	2009(1)–2009(2)			
	2010(5)–2010(5)			
	2012(2)–2012(10)			

In order to evaluate the extent to which turning points in the two series occur near each other and to evaluate synchronization of cycles the correlation of business cycle in given countries and crude oil price cycle and concordance index respectively have been calculated¹⁰. For the comparison purposes these calculations are carried out for all chosen countries, although only in the case of Poland they are sensible.

The correlation coefficients between probabilities of being in regime 1 (recession and drops in oil prices for business cycle and crude oil price cycle respectively) are displayed in Figure 4.

Generally, the correlations between probabilities of being in regime 1 (Figure 4) are weak what means that turning points of business cycle in the recessionary periods does not occur close to turning points of crude oil price cycle. The highest correlation of business cycle and oil price cycle is observed in the case of Poland (32%), and the lowest – in the case of Slovenia (10%).

¹⁰ Harding, Pagan (2006) and Konopczak (2009) indicate that the correlation coefficient and concordance index are the most frequently used measures of the synchronization of cycles identified by the Markov-switching model.

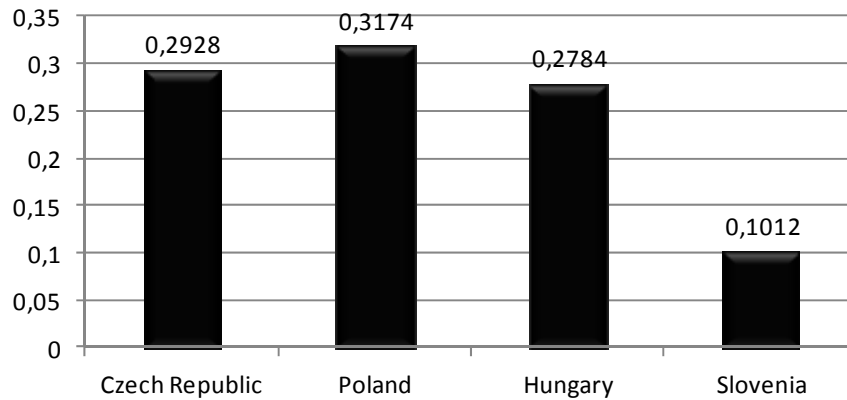


Figure 4. Correlation coefficients for probabilities of being in regime 1 for business cycle in given countries and crude oil price market in 1995–2013 period

Table 9. Concordance index for oil price cycle and business cycle for given economies

	Crude oil price cycle vs. business cycle			
	Czech Republic	Poland	Hungary	Slovenia
Concordance index	0.7143	0.6866	0.7051	0.7235

To evaluate synchronization of business cycle for different economies and oil price cycle the concordance index was applied. This index (for two series x_t , y_t and a sample size of T) takes the form (Harding, Pagan, 2006):

$$\hat{I} = \frac{1}{T} \left[\sum_{t=1}^T S_{xt} S_{yt} + \sum_{t=1}^T (1 - S_{xt})(1 - S_{yt}) \right], \quad (5)$$

where S_{xt} and S_{yt} denote binary variables that takes the value unity in case of recession regime and zero – in case of expansion regime at time t .

Values of concordance index measuring the synchronization of cycles (in terms of being in the same state) indicate a high degree of concordance of the business cycles and oil price cycle (Table 9). However, having in mind earlier remarks on the quality of regime classification, the above conclusion refers only to the case of Poland. This means that business cycle in Poland in 1995–2013 period has the same state (recession or expansion) as the Brent oil price cycle (drops or rises in oil prices). Even in that case the caution

should be taking when interpreting the results because a low value of correlation coefficients corresponds to a high degree of concordance of business cycles and oil price cycle. This may suggest the overestimated degree of concordance.

Conclusions

In this paper we used the approach innovated by Hamilton in his analysis of the US business cycle to identify the Brent crude oil price cycle and business cycles in a set of CEE economies in 1995–2013 period. The obtained results support to a varying degree our modelling approach based upon Markov-switching. The MSIH(2)-AR(2) model produced the sharpest classification of regimes (drops or rises in oil prices) was obtained in the case of crude oil price cycle. While the MSIA(2)-AR(2) model produced the sharpest classification of regimes (recession on expansion) for Poland. In the case of Czech Republic and Hungary the MS-AR models were not able to distinguish between two regimes (switches between regimes occurred very often) and as a result the existence of two regimes is doubtful. The MS-AR models obtained for Slovenia indicated either a poor model specification or difficulties in distinguishing between two regimes. Therefore the results for Slovenia should be taken with caution.

Our results indicated that the MS-AR models captured the different regimes of Brent crude oil price cycle and business cycle for the case of Poland in a satisfactory way. The correlation of oil price cycle and business cycle in Poland turned out to be rather weak, however the synchronization of cycles occurred (a high degree of concordance) what indicates that there are co-movements in crude oil cycle and business cycles. Week correlation of oil price cycle and business cycle may come from the specificity of Poland, i.e. the size of crude oil imported by these countries from the North Sea region is small (in Poland the import of crude oil from Norway amounts to 3.5% of total crude oil import; POPiHN, 2013). Poland but also the other CEE countries are still dependent on the crude oil imported from Russia which is cheaper but worse in quality (it contains a high amount of the impurity sulfur and has high density).

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Analiza zbieżności cykli cenowych rynku ropy naftowej z cyklami koniunkturalnymi gospodarek Europy Środkowo-Wschodniej

Z a r y s t r e ś c i. Głównym celem artykułu jest zbadanie, w jakim stopniu cenowe cykle ropy naftowej (Brent) są skorelowane i zsynchronizowane z cyklem koniunkturalnym dla wybranych gospodarek państw Europy Środkowo-Wschodniej (EŚW). W celu identyfikacji cyklu cenowego ropy naftowej i cyklu koniunkturalnego został zastosowany przełącznikowy model Markowa (MS-AR). Określenie wygładzonych prawdopodobieństw w zależności od reżimu umożliwiło wyznaczenie korelacji tych prawdopodobieństw oraz obliczenie indeksu konkordancji dla oceny stopnia synchronizacji cyklu cenowego ropy naftowej i cyklu koniunkturalnego w badanych państwach EŚW.

S ł o w a k l u c z o w e: przełącznikowe modele Markowa, ceny ropy naftowej, cykle koniunkturalne, cykle cenowe.