Oil price shocks and domestic inflation in Thailand

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Abstract

Using monthly data during 1993 and 2017, this paper examines the impact of oil price shocks on the domestic inflation rate in Thailand. Both linear and nonlinear cointegration tests are used to examine the long-run relationship between price level, industrial production, and the real price of oil. Furthermore, the two-step approach is used to examine how an oil price shock and oil price volatility affect the inflation rate. In addition, the asymmetry of oil price shocks on inflation is also investigated. The results show that price level is positively affected by the real oil price and industrial production index in the long run. The short-run analysis reveals that there is a positive relationship between oil price shock and domestic inflation. The estimated results from the two-step approach show that an oil price shock causes inflation to increase while oil price uncertainty does not cause inflation. Furthermore, the short-run relationship between inflation and oil price shocks is statistically significant. However, the asymmetric impacts of oil price shocks on inflation are not apparent. The findings from this study will encourage the monetary authorities to formulate a more accommodative policy to respond to oil price shocks, which positively affect the inflation rate. In addition, oil subsidization by the government should not be abandoned.

Keywords: Oil shocks, inflation, cointegration, VAR, bivariate GARCH, causality
JEL Classification: E31, Q43

1. Introduction

One of the interesting topics related to the relationship between oil shocks and macroeconomic variables is the impact of oil price shocks on the inflation rate. The rise of oil prices can cause firms' production costs to increase. Therefore, an oil price hike is reflected in an increase in the general price level of an economy. In addition, changes in the oil price in the last five decades exhibit oil price volatility that can distort the decisions made by economic agents. Oil price shocks affect economic performances via both demand and supply channels (Lee and Ni, 2002; Valcarcel and Wohar, 2013). Inflation induced by oil price shocks can reduce real balances, a measure of purchasing power, in the economy, thus causing a recession (Mork and Hall, 1980 and Mork, 1989). The stagflation threat from the oil shocks in the 1970s should not be underestimated (Bernanke et al., 1997). The US Federal Reserve adopted too high an interest rate policy and thus did not respond to oil price shocks accurately. This results in either decreased output or recession in the US. Oil price changes matter because they disrupt spending by consumers and firms on key sectors, and thus reducing output growth (Hamilton, 2003). However, Barsky and Kilian (2004) argue that the perception that only oil price shocks can explain the U. S. stagflation of the 1970s is not necessarily true. As to the supply channel, oil price shocks can cause consumer prices to increase via the supply channel. This phenomenon depends on the share of the oil price in the price index.

Oil price shocks and oil price volatility can exert impacts on key macroeconomic variables, and thus create problems for economic agents and policymakers. The present paper focuses on a specific analysis of the oil price pass-through into inflation in Thailand, which is an oil-importing country in the Southeast Asian region.

Co-movement between price level and the real domestic oil price series in Thailand is plotted in Fig. 1. Even though the real oil price is linked to the trend of the price level, the real oil price variable fluctuates more. Starting from a low oil price with some fluctuations, the new oil shocks in 2000 caused real oil price to increase. The oil price reaches its peak in mid-2008 and falls to the trough at the end of 2008, then starts to rise and keeps on fluctuating. Oil price volatility plotted in Fig. 2 shows that high volatility occurs around 2000 and again around 2009 and 2015.[1]

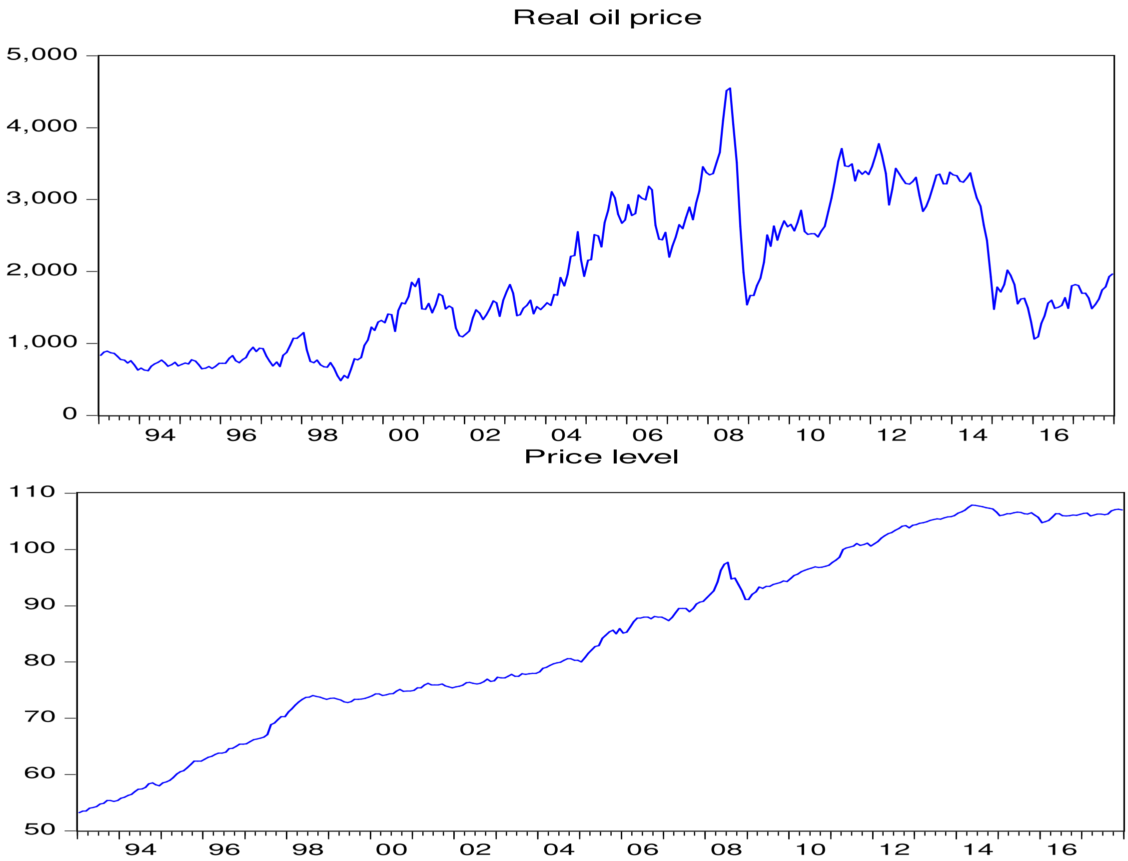

Fig. 1 Co-movement of price level with the real oil price.

[1] Real oil price volatility series are generated by a bivariate GARCH model reported in Section 3.

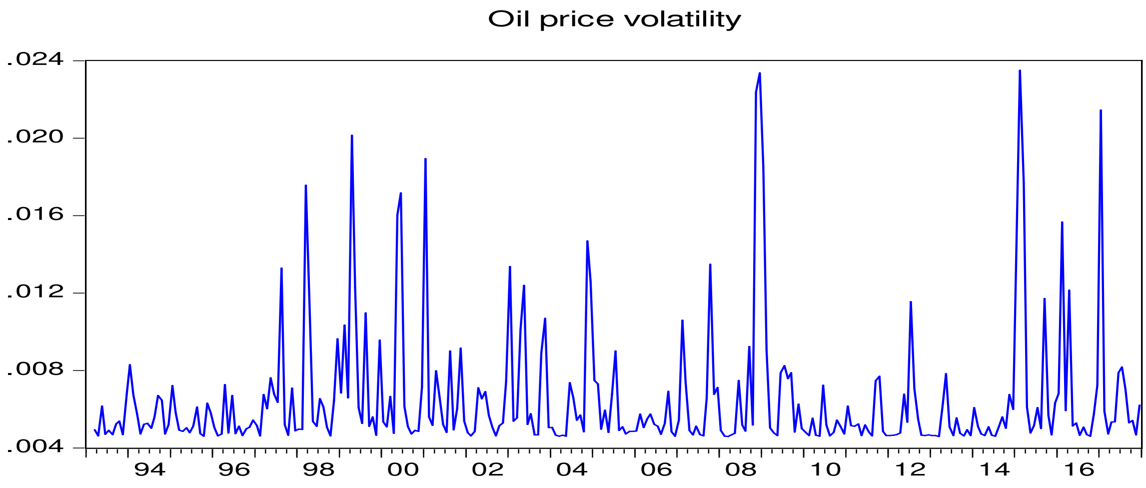

Fig. 2 Volatility of real oil price.

The monetary policymakers in Thailand have tried to maintain price stability by adopting inflation targeting in 2000. The main purpose of the present study is to examine the role of oil price and its volatility in exerting an impact on inflation besides the role of monetary policy. Cunado and Perez de Gacia (2005) use two different definitions of oil prices. This study uses their second definition, which is the real price of oil.[2] In addition, an oil price shock is the real domestic oil price in first differences or an oil price change. Furthermore, an increase in real oil price is defined as a positive shock while a decline in real oil price is a negative shock. This paper contributes to the existing literature by providing evidence showing that the long-run impact of oil price shocks on domestic inflation in a net oil-importing country is found from a nonlinear cointegration test. Even though this impact is not large, but long-lasting. In addition, the short-run impact of oil price shocks on inflation is not asymmetric as found in some previous studies. Furthermore, oil price volatility does not cause inflation, but inflation itself causes inflation uncertainty in the Thai economy. This paper is organized as follows. The next section reviews the literature. Section 3 presents the data and estimation methods that are used in the analysis. Section 4 presents the empirical results and discusses the results found in this study. The last section gives concluding remarks and some policy implications based on the results of this study.

2. Literature Review

Specific studies have focused on the impact of oil price changes on the inflation rate. This impact is often referred to as oil price pass-through to changes in the price index (Olifin and Salisu, 2017). The oil price pass-through has been examined for both advanced and developing economies. The oil price changes on inflation in the US seem to affect inflation through the direct share of the oil price in consumer prices. Furthermore, monetary policy has become less accommodative of oil price shocks, thus preventing oil price changes from

[2] Most studies concerning the impact of oil prices on macroeconomic variables in advanced countries use different definitions of the price of oil. For example, Hamiltion (1996) uses the world price of crude oil while Cologni and Manera (2009) use the real price of oil as one of the various definitions of oil shocks. Also, real oil prices can reflect both the true purchasing power and the cost of production.

passing directly into core inflation (Hooker, 2002). Oil prices also lead the cycle of consumer prices. The impact of oil price on inflation in industrialized countries can decline due to certain factors, i.e., the inflationary effect of oil price change is limited (Ewing and Thompson, 2007; De Gregorio et al., 2007; Alvarez et al., 2011). There is an argument that the effect of oil price changes is stronger due to temporarily accommodating monetary policy and structural changes in the US economy (Fukac, 2011). In addition, the oil price-inflation nexus may have shifted from a supply-side to a demand-side phenomenon in the US since the great moderation period (Valcarcel and Wohar, 2013). Therefore, it can affect the ability of monetary policymakers in dealing with the impacts of oil price shocks on output and inflation. Some studies find that the degree of positive impacts of oil price shocks on disaggregate US consumer prices is observed only in energy-intensive consumer price indices (De Gregorio et al., 2007 and Gao et al., 2014). In addition, the main causes of the impacts are the rises in the prices of energy-related commodities. For selected OPEC and EU nations, the relationship between oil price shocks and inflation seems to be stronger in oil-exporting than in oil-importing countries (Olofin and Salisu, 2017). A decline in the impacts of oil price shocks on inflation has been found in the US and European economies. Bachmeier and Cha (2011) use the US disaggregate inflation data to explain the breakdown of the relationship between oil price shocks and consumer price inflation. They find that a decline in the oil price pass-through mainly depends on reduced energy usage in two-thirds of the sectors. Only one-third of the sectors have a decline in the pass-through due to a change in the monetary policy response to oil price shocks. Conflitti and Luciani (2017) examine the oil price pass-through into consumer prices in the US and the Euro area. They find that the estimated oil price pass-through is small but long-lasting. Castro et al. (2017) find that the oil price pass-through into inflation in four main European economies is quantitatively diverse across economies and items of disaggregate inflation.

Besides the direct role of oil price shocks on inflation, oil price volatility also plays a role in inflationary pressure. Previous studies document that oil price shocks can have an adverse impact on the output because they raise the level of oil prices and oil price volatility. Similar to the impact on output, oil price shocks not only have a positive impact on inflation, but the impact can be asymmetric due to the response of the economy to oil price volatility (Federer, 1996). The asymmetric impacts imply that positive and negative oil price shocks have different impacts on inflation. However, Farzanegan and Markwardt (2009) find that both positive and negative oil price shocks exert positive impacts on inflation in Iran. Their results also show that negative oil price shocks have a stronger short- and long-run effect on inflation compared to positive oil price shocks. Therefore, the asymmetric impacts of oil price shocks on inflation are not found. Ajmi et al. (2015) find similar results for South Africa. They use an asymmetric causality test to examine the relationship between international oil prices and price level. They find no cointegration between oil prices and price level. However, they find a causal relationship running from oil prices to price level. Furthermore, both positive and negative oil price shocks have a positive impact on price level changes, but a negative oil price shock has a stronger effect. Olofin and Salisu (2017) find that oil price asymmetries seem to matter more in oil-exporting than oil-importing countries. Castro et al. (2016) find that the inflationary effect of oil prices remains in the Euro area because no deflationary effect of oil prices will result in a negative inflation rate. However, a recent study by Lahiani (2019) shows that the US inflation reacts differently to positive and negative oil price shocks.

For some Asian economies, Cunado and De Gracia (2005) use quarterly data from 1975 to 2000 to examine the impact of oil price shocks on economic activities and inflation in Japan, Singapore, South Korea, Malaysia, Thailand, and the Philippines. They find that the impact is more pronounced when oil prices are measured in domestic currencies. Chu and Lin (2013) find that oil price shocks have both long-term and short-term effects on Taiwan's producer price index. Mandal et al. (2012) estimate the impact of international oil prices on the Indian economy after the deregulation of prices of some petroleum products. They find that the oil price pass-through to domestic inflation and industrial output is higher than that of the global trend. Du et al. (2010) find that there are nonlinear impacts of world oil price shocks on inflation and growth in China. Also, Ju et al. (2014) find that oil price shocks cause China's consumer price index to rise. Razmi et al. (2016) find that oil prices exert stronger impacts on consumer price indices in the post-global financial crisis than the pre-global financial crisis in Indonesia, Malaysia, the Philippines, and Thailand. For the role of oil price volatility, Rafiq et al. (2009) examine the impact of oil price volatility, measured by realized volatility, on key macroeconomic indicators of Thailand using quarterly data from 1993 to 2006. They find that there is a unidirectional causality running from oil price volatility to economic growth, investment, unemployment, and inflation. However, the results from impulse response analysis show that the impact of oil price volatility on inflation lasts for only a short time horizon. Rafiq and Salim (2014) find that oil price volatility affects output growth, but does not affect inflation in Thailand. However, the impact on output growth disappears after the financial crisis because the Thai government implemented oil subsidization after the crisis.

In previous studies mentioned above, different methods are used to analyze the impact of oil price changes and oil price volatility on the inflation rate. The popular methods are symmetric and asymmetric cointegration techniques, structural vector autoregression (SVAR), and generalized autoregressive conditional heteroskedastic (GARCH) models.

3. Data and Methodology

In this section, the data and their properties are presented. The estimation methods used in the analysis are described.

3.1 Data

The dataset used in this study comprises monthly data during 1993 and 2017. The rationale for using this period is that the availability of industrial production index dated from 1993. In addition, monthly data give a larger sample size than using quarterly data. The consumer price index, industrial production index, and the US dollar exchange rate series are obtained from The Bank of Thailand's website. The series of Brent crude oil spot price expressed in US dollars per barrel is obtained from the US Energy Information Administration. The oil price series is the international oil price. By multiplying the oil price series by the US dollar exchange rate and deflating by the consumer price index, the real domestic oil price series is obtained.[3] All series used in estimation are seasonally adjusted and transformed into logarithmic series. The sample size comprises 300 observations.

[3] Cunado and De Gracia (2005) find that this measure of real domestic oil price is more important than that of real international oil price, which does not take into account the impact of the exchange rate that influences the domestic oil price.

The conventional unit root tests can have low power in the presence of structural breaks in the series. To overcome this problem, unit root tests with an unknown structural break date proposed by Zivot and Andrews (1992) are performed on both levels and first differences of the series. The results are shown in Table 1.

Table 1
Results of Zivot-Andrews tests for unit root: 1993M01-2017M12.

Variables	Test A	Break date	Test B	Break date
p (Level of consumer price index)	-3.905 [2] (0.192)	2005M01	-2.907[2] (0.936)	2014M05
Δp (Difference in consumer price index)	-9.487***[1] (0.000)	1998M06	-9.856***[0] (0.000)	1997M08
ip (Level of industrial production index)	-3.350 [7] (0.474)	2002M12	-3.375 [7] (0.765)	2002M12
Δip (Difference in industrial production index)	-8.554***[6] (0.000)	2011M10	-8.543***[6] (0.000)	2011M10
o*p* (Level of real oil price)	-4.055 [1] (0.139)	1999M02	-3.639 [1] (0.604)	1999M02
Δop (Difference in real oil price)	-15.100***[0] (0.000)	1999M13	-7.116***[12] (0.000)	2016M12

Note: Test A includes intercept only while Test B includes intercept and a linear trend. The numbers in the bracket represent the optimal lag length determined by the Schwarz information criterion (SIC),. ***, ** and ** denote significance at the 1%, 5% and 10% level, respectively. The numbers in parenthesis represent the probability of accepting the null hypothesis of unit root provided by Vogelsang (1993).

The results from unit root tests show that the degree of integration of all series is one, i.e., they are I(1) series. The null hypothesis of unit root cannot be rejected for the levels of series, but it is rejected at the 1% level of significance for the first difference of series. Therefore, it can be concluded that all series are I(1). This is suitable for performing cointegration tests. The stationary property of the first differences of series is also suitable in the estimate of a bivariate generalized autoregressive conditional heteroskedastic (GARCH) model as well as the standard pairwise causality test described in the next sub-section.[4]

The basic characteristics of the level and first difference of the time series data are described in Table 2.

[4] A bivariate GARCH model requires that all series be stationary.

Table 2
Descriptive statistics: 1993M01-2017M12.

A. Level of series			
Variable	*p*	*ip*	*op*
Mean	4.417	4.927	7.407
Median	4.421	5.041	7.423
Maximum	3.978	5.435	8.347
Minimum	4.678	4.205	6.191
Standard deviation	0.204	0.378	0.564
Skewness	-0.400	-0.280	-0.265
Kurtosis	2.104	1.488	1.844
JB	18.048	32.501	20.211
	(0.000)	(0.000)	(0.000)
Observations	300	300	300
B. First difference of series			
Variable	Δp	Δip	Δop
Mean	0.002	0.004	0.003
Median	0.002	0.004	0.010
Maximum	0.022	0.246	0.246
Minimum	-0.028	-0.354	-0.275
Standard deviation	0.004	0.041	0.082
Skewness	-0.723	-1.697	-0.355
Kurtosis	12.048	26.419	4.166
JB	1046.016	6976.407	23.211
	(0.000)	(0.000)	(0.000)
Observations	299	299	299

Note: JB is Jarque-Bera statistic with the p-value in parenthesis.

For the level of series, consumer price index, domestic real oil price, and industrial production are negatively skewed, but all series do not show excess kurtosis. The Jarque-Bera statistics reveal that all series are not normally distributed. The average monthly inflation rate is 0.2 percent, whereas the average monthly oil price shock is 0.3 percent and the average monthly industrial production is 0.4 percent. All series exhibit excess kurtosis and are negatively skewed. The Jarque-Bera normality test rejects the null hypothesis of normal distribution of all series, which indicates that there may be the presence of an autoregressive conditional heteroskedastic (ARCH) effect.

3.2 Estimation Methods

The methods used in the analysis comprise cointegration tests, vector autoregressive (VAR) model analysis, and bivariate generalized autoregressive conditional heteroskedastic (GARCH) model along with Granger causality tests.

3.2.1 Cointegration tests

Since industrial production can interact with both price levels measured by consumer price index and real oil price, trivariate cointegration analysis can be used to test whether there is a positive long-run relationship between the price level and the real domestic oil price when industrial production is treated as a control variable.

3.2.1.1 Residual-based cointegration tests with unknown breakpoints

Similar to the conventional residual-based test for cointegration, this test proposed by Gregory and Hansen (1996) is similar to Engle and Granger (1987) procedure in that it can be used by estimating the relationship between non-stationary series: price level proxied by the consumer price index, domestic oil price and industrial production. However, the Gregory-Hansen procedure takes into account the impact of an unknown level shift or a structural break. The relationship in a trivariate framework can be expressed as:

$$p_t = a + b_1 D_t + b_2 ip_t + b_3 op_t + e_t \qquad (1)$$

In Eq. (1), if real oil price (op_t) and industrial production (ip_t) have impacts on price level (p_t), the coefficients b_2 and b_3 should be statistically significant. The dummy variable, D_t, captures an unknown breakpoint. The residual series, e_t, obtained from the estimate of Eq. (1) can be used to test for a unit root using the Augmented Dickey-Fuller (ADF) test, which is expressed as:

$$\Delta e_t = \rho e_{t-1} + \phi \Delta e_{t-1} \qquad (2)$$

The t-statistic obtained from the estimation of Eq. (2) is the ADF* statistic. This statistic is used for comparison with the critical value statistic provided by Vogelsang (1993). If the ADF* statistic is larger than the critical value, the null hypothesis of unit root in the residual series will be rejected. Therefore, there is cointegration or long-run relationship expressed in Eq. (1). On the contrary, the smaller value of the ADF* statistic than that of the critical value leads to an acceptance of the null hypothesis of unit root, and thus there is no cointegration between variables in the model.

The existence of cointegration from Eq. (1) indicates that the relationship between price level, industrial production, and real domestic oil price can be represented by the symmetric error correction model (ECM) that can be expressed as:

$$\Delta p_t = \varphi_0 + \lambda e_{t-1} + \sum_{i=1}^{k} \varphi_{21} \Delta p_{t-i} + \sum_{i=1}^{k} \varphi_{3i} \Delta ip_{t-i} + \sum_{i=1}^{k} \varphi_{4i} \Delta op_{t-1} + u_t \qquad (3)$$

where EC_{t-1} is the lagged value of the corresponding error term, which is called the error correction term (ECT), and λ, φ_{2i}, φ_{3i}, and φ_{4i} are the regression coefficients while u_t is a random variable. The sign of the coefficient of the ECT should be negative and has an absolute value of less than one. If this coefficient is statistically significant, any deviation from the long-run equilibrium will be corrected and thus the long-run relationship is stable.

3.2.1.2 Nonlinear cointegration tests

It is important to confirm that the relationship between variables is not nonlinear. In the case of the absence of linear cointegration between variables, the long-run relationship might be nonlinear with asymmetric adjustment towards the long-run equilibrium. Therefore, the threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) models can be utilized. The two models are residual-based tests developed by Enders and Granger (1998) and Enders and Siklos (2001). The residuals from the estimate of Eq. (1) are decomposed and the test equation is expressed as:

$$\Delta e_t = I_t \rho_1 e_{t-1} + (1 - I_t)\rho_2 e_{t-1} + \sum_{i=1}^{k} \beta_i \Delta e_{t-i} + u_t \qquad (4)$$

where $u_t \sim$ iid.$(0,\sigma^2)$ and the lagged augmented term ($\Delta\hat{e}_{t-i}$) can be added to yield uncorrelated residuals of the estimates of equation (4). The Heaviside indicator function for TAR is specified in Eq. (5) while this function for MTAR is specified in Eq. (6), which are:

$$I_t = \begin{cases} 1\_if\_e_{t-1} \geq \tau \\ 0\_if\_e_{t-1} < \tau \end{cases} \qquad (5)$$

and

$$I_t = \begin{cases} 1\_if\_\Delta e_{t-1} \geq \tau \\ 0\_if\_\Delta e_{t-1} < \tau \end{cases} \qquad (6)$$

where the threshold value τ can be determined endogenously. According to Pertrucelli and Woolford (1984), the necessary and sufficient conditions for the stationarity of $\{e_{t-1}\}$ are $\rho_1 < 0$, $\rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$. The long-run equilibrium value of the error term should be less than zero when these conditions are met. Ender and Siklos (2001) propose two test statistics for the null hypothesis of no cointegration, i.e., t-Max and the F statistic called Φ. If cointegration exists, the t-Max and Φ statistics should be larger than the 5% critical values in absolute terms. However, the Φ statistic has substantially more power than the t-Max statistic for testing the null hypothesis of $\rho_1 = \rho_2 = 0$ or no cointegration. The main drawback of the Φ statistic is that it can lead to the rejection of the null hypothesis when only one of the rho coefficients is negative. Therefore, Enders and Siklos (2001) suggest that the Φ statistic should be used when rho coefficients are both negative and have the absolute values of less than one.

If the tests indicate the existence of linear cointegration between price level, industrial production, and real oil price, the time-series dynamics of the relationship between the two variables can be explored by threshold error correction mechanisms (TECMs). The TECMs can be expressed as:

$$\Delta p_t = \varphi_0 + \lambda_1 e_{t-1} + \sum_{i=1}^{k} \varphi_{21} \Delta p_{t-i} + \sum_{i=1}^{k} \varphi_{3i} \Delta ip_{t-i} + \sum_{i=1}^{k} \varphi_{4i} \Delta op_{t-1} + u_{1t} \qquad (7)$$

and

$$\Delta p_t = \tilde{\varphi}_0 + \lambda_2 e_{t-1} + \sum_{i=1}^{k} \tilde{\varphi}_{21} \Delta p_{t-i} + \sum_{i=1}^{k} \tilde{\varphi}_{3i} \Delta ip_{t-i} + \sum_{i=1}^{k} \tilde{\varphi}_{4i} \Delta op_{t-1} + u_{1t} \qquad (8)$$

where k is the lag order, λ_1 and λ_2 are the coefficients showing the speeds of adjustment.[5] The short-run dynamics allow for testing the alternative hypothesis on the short-run relationship between price level, industrial production, and real oil price. The coefficients of the lagged

[5] The speed of adjustment is $\lambda_1 = I_t \rho_1$ in the first regime and $\lambda_2 = (1 - I_t)\rho_2$ in the second regime while I_t in equation (5) is used for the TAR model, and I_t in Eq. (6) is used for the MTAR model.

differences for industrial production and real oil price show the short-run impacts of the two variables on the first difference of price level while the coefficients of the asymmetric errors correction terms are the speeds of adjustment toward the long-run equilibrium. Eqs. (7) and (8) can also be used to test for short-run causality between the price level and industrial production, and the price level and real oil price when asymmetric adjustment is found.

3.2.2 Short-run analysis of oil price shock, oil price volatility, and inflation

3.2.2.1 VAR-GARCH(1,1) model and Granger causality

The two-step approach is employed to explain the relationship between real oil price and its uncertainty (or volatility) as well as inflation and its uncertainty. In the first step, a bivariate VECH-GARCH(1,1) model proposed by Bollerslev et al. (1988) is employed to generate inflation uncertainty and oil price volatility. This model allows the conditional covariance matrix of the dependent variables to follow a dynamic structure. Each conditional variance (or volatility) depends on past shocks and its own past conditional variance. In the second step, these generated series along with the inflation rate and the series of real oil price changes are employed in the standard Granger (1969) causality test. Pagan (1984) criticizes this procedure because it produces the generated series of volatility or uncertainty. When these generated series are used as regressors in the Granger causality test, the model might be misspecified. However, it can be argued that the main advantage of the two-step procedure is that it provides room for the ability to establish causality between variables.[6] The system equations in a diagonal VECH-GARCH(1,1) model proposed by Bollerslev et al. (1988), comprise the following five equations.

$$\Delta p_t = a_{1,0} + \alpha_{1,1}\Delta p_{t-1} + \alpha_{1,2}\Delta op_{t-1} + e_{1,t} \tag{9}$$

$$\Delta op_t = a_{2,0} + \alpha_{2,1}\Delta op_{t-1} + \alpha_{2,2}\Delta p_{t-1} + e_{2,t} \tag{10}$$

$$h_t^{\Delta p} = \mu_1 + \alpha_{1,1}(\varepsilon_{t-1}^{\Delta p})^2 + \beta_{1,1}h_{t-1}^{\Delta p} \tag{11}$$

$$h_t^{\Delta op} = \mu_2 + \alpha_{2,1}(\varepsilon_{t-1}^{\Delta op})^2 + \beta_{2,1}h_{t-1}^{\Delta op} \tag{12}$$

$$h_t^{\Delta p,\Delta op} = \mu_3 + \alpha_{31}(\varepsilon_t^{\Delta p})(\varepsilon_t^{\Delta op}) + \beta_{3,1}h_{t-1}^{\Delta p,\Delta op} \tag{13}$$

where Δp is the change in the price level or inflation, and Δop is the change in real oil price or oil price shock, $h^{\Delta p}$ is the conditional variance of inflation, $h^{\Delta op}$ is the conditional variance of real oil price change, and $h^{\Delta p,\Delta op}$ is the conditional covariance of the two variables. The system equations can be estimated simultaneously. The estimated results can be used in Granger (1969)'s causality tests.

3.2.2.2 Testing for asymmetric impacts of real oil price shock on inflation

One of the important issues concerning the short-run relationship between inflation and oil price shock is the asymmetric effects of oil price increases and decreases on the inflation rate. Following Mork (1989)'s procedure, the oil price shock series is separated into positive and negative shocks. By applying unrestricted VAR models and VAR Granger causality/block exogeneity Wald tests, the test equations can be expressed as:

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[6] The current value of one variable might not affect the current value of another variable, but some of its lags might do.

$$\Delta p_t = a_{01}^{+} + \sum_{i=1}^{k} a_{1,i}^{+} \Delta op_{t-i}^{+} + \eta_{1,t}^{+} \quad (14a)$$

and

$$\Delta op_t^{+} = a_{02}^{+} + \sum_{i=1}^{k} a_{2,i}^{+} \Delta p_{t-i} + \eta_{2,t}^{+} \quad (14b)$$

$$\Delta p_t = b_{01}^{-} + \sum_{i=1}^{k} b_{1,i}^{-} \Delta op_{t-i}^{-} + \nu_{1,t}^{-} \quad (15a)$$

and

$$\Delta op_t^{-} = b_{02}^{-} + \sum_{i=1}^{k} b_{2,i}^{-} \Delta p_{t-i} + \nu_{2,t}^{-} \quad (15b)$$

where Δop^{+} is a series of oil price increases and Δop^{-} is a series of oil price decreases. The lag order, k, can be determined by SIC. With this specification, the conventional Chi-square tests can detect the existence of causality. Furthermore, the estimates of unrestricted VAR models can determine the sizes of the impacts of positive and negative oil price shocks on the inflation rate.

4. Results and Discussion

This section reports the results from cointegration tests and short-run dynamics, impulse response functions (IRFs) and variance decomposition (VDCs), and Granger causality tests.

4.1 Long-run relationship and short-run dynamics

The models expressed in Eqs. (1) and (2) are used for testing the existence of a long-run relationship between variables. The results from Gregory-Hansen testing for cointegration are shown in Table 3.

The results reported in Table 3 show the level relationship between price level, industrial production, and the price of oil, but there is no cointegration between the three variables because the estimated model with level shift shows that the ADF* statistic of -3.99 is smaller than the 5% critical value of -5.29 for the three-variable model. The break date is 1997M04 for the level shift model, which is three months prior to the occurrence of the 1997 Asian financial crisis.

Table 3
Coefficients of level relationship obtained from the estimated model.
Dependent variable is p_t

Independent variable	Coefficient	t-statistic	p-value
D_t	0.096***	7.151	0.000
ip_t	0.427***	17.222	0.000
op_t	0.036**	2.440	0.015
intercept	2.207***	42.942	0.000
Adj. R^2 = 0.922			

Note: ***, ** and * indicate significance at the 1%, 5%, and 10% levels, respectively.

There is a positive level relationship between price level, industrial production, and real oil price. A one percent increase in industrial production causes the price level to rise by 0.43 percent and vice versa. Since the industrial sector comprises many energy-intensive manufacturing firms, the reduction of oil use by firms cannot be avoidable. Therefore, firms in the manufacturing sector can adjust themselves to oil price shocks in the long run. For the real price of oil, the positive relationship between the price level and real oil price is not surprising. A one percent increase in the real oil price causes the price level to rise by 0.04 percent. The breakpoint at 1997M4, which is near the occurrence of the 1997 Asian financial crisis, strengthens the level relationship of these three variables.

The claimed long-run positive impact of the real oil price on the general price level in a trivariate framework will be valid if threshold cointegration is found. In testing for threshold cointegration, the residuals are obtained from the estimation of level relationship by Gregory and Hansen's (1996) cointegration test, which includes the breakpoint dummy variable as suggested by the results from Table 3. The heavy side indicator (I) is specified in Eq. (5) for the TAR model while this indicator is specified in Eq. (6) for the MTAR model. The results from the estimated TAR and MTAR models are reported in Table 4.

Table 4

Results of threshold cointegration between the price level and real oil price.

	TAR	MTAR
ρ_1	-0.157 (0.041)	-0.104 (0.033)
ρ_2	-0.084 (0.041)	-0.812 (0.062)
τ	0.026	-0.015
κ	1	1
t-Max	-2.063 [-1.836]	-2.908 [-1.931]
Φ	9.049 [6.938]	8.861 [8.622]
F-Equal	1.618 [4.652]	1.262 [6.787]

Note: Standard error is in parenthesis, the number in the bracket represents the 5% critical value, τ is the threshold value, κ is the number of lagged augmented term determined by SIC, Φ is the F $\rho_1 = \rho_2 = 0$, and F-Equal is F $\rho_1 = \rho_2$.

The results reported in Table 4 show that the estimated ρ_1 and ρ_2 are negative with the absolute value of less than one and $(1+\rho_1)(1+\rho_2)$ is equal to 0.772 for the TAR model. Similarly, for the MTAR model, ρ_1 and ρ_2 are also negative with the absolute value of less than one and $(1+\rho_1)(1+\rho_2)$ is equal to 0.168. Therefore, the convergence condition is met for both the TAR and MTAR models. The Φ statistic is 9.049 and 8.861 for the TAR and MTAR models while the simulated critical values at the 5% level are 6.938 and 8.662 for the TAR and MTAR models, respectively. Even though the t-Max statistic has low power of test than the Φ statistic, the absolute value of this statistic is larger than the 5% critical value. Therefore, the null hypothesis of no threshold cointegration is rejected at the 5% level of significance for both models. In other words, there is nonlinear cointegration between price level, industrial production, and the real price of oil. Furthermore, the null hypothesis of no asymmetric adjustment toward the long-run equilibrium cannot be rejected because the F-Equal statistic is smaller than the critical value at the 5% level of significance for both the TAR and MTAR models. Thus the test results indicate that there is nonlinear cointegration between the three variables without asymmetric adjustment towards long-run equilibrium. Therefore, the results reported in Table 3 can be the long-run relationship between price level, industrial production index, and real oil price is nonlinear without asymmetric adjustment. This result is different from the result of Cundo and De Gracia (2005) who find no long-run relationship between the price level and domestic oil price in Thailand using a linear cointegration test under a bivariate framework.

Even though the results from Gregory and Hansen's (1996) test reveals that there is no long-run relationship between price level, industrial production, and domestic real oil price, the estimated TAR and MTAR models show the presence of threshold cointegration without asymmetric adjustment in the nonlinear long-run relationship shown in Table 3 at the 5% level of significance. Therefore, the asymmetric threshold ECMs specified in Eqs. (7) and (8) cannot be utilized. The short-run dynamics of symmetric adjustment towards the long-run equilibrium specified in Eq. (3) are estimated. The results are reported in Table 5.

Table 5
Results of short-run dynamics.
Dependent variable: Δp_t

Variable	Coefficient	Standard Error	t-statistic	p-value
$\hat{e}_{t-1}$	-0.014***	0.004	-3.179	0.002
Δp_{t-1}	0.183***	0.058	3.171	0.002
Δip_{t-1}	-0.006	0.006	-0.983	0.326
Δop_{t-1}	0.010***	0.003	3.147	0.002
Intercept	0.002***	0.001	6.944	0.000
Adjusted R^2 = 0.115, F = 10.665				
Serial correlation test: $\chi^2_{(2)}$ = 5.631 (p-value = 0.059)				
ARCH test: $\chi^2_{(1)}$ = 1.947 (p-value = 0.163)				

Note: ***, ** and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Since SIC indicates that the lag should be zero, a parsimonious ECM with a lag of 1 is estimated. The estimated short-run equation passes important diagnostic tests, i.e., there are no serial correlation and no further ARCH effects in the residuals. The coefficient of the error correction term has a correct sign with the absolute value of 0.014, which is less than one and significant at the 1% level. Therefore, any deviation from long-run equilibrium will be corrected.[7] However, the coefficients of lagged changes in industrial production are insignificant. Using the Wald F test, $F_{1,293}$ = 0.967 with p-value = 0.326, and thus short-run causality running from changes in industrial production to inflation is not found. On the contrary, positive short-run causality running from oil price shock to inflation is found at the 1% level of significance because the Wald $F_{1, 293}$ = 9.902 with p-value = 0.002, which leads to a rejection of the null hypothesis of no causality. The Wald $F_{1,293}$ for the coefficient of the error correction term is 10.107 with p-value = 0.002, which indicates that there is long-run causality running from industrial production and real oil price to inflation (Granger, 1988).

4.2 Short-run relationship and the role of oil price volatility

In analyzing the short-run relationship, the two-step approach explained in the previous section is utilized. First, a bivariate GRACH(1,1) model is estimated to obtain two volatility series. The next step is to employ the standard Granger causality test and an unrestricted VAR model to examine short-run causality.

[7] In other words, a deviation from the long-run equilibrium will be corrected by 1.4% per month.

In performing a bivariate GARCH estimate, the unit root statistics for the full sample period reported in Table 1 show that the first differences of the two series are stationary and thus suitable for the estimation. The bivariate GARCH model for the system Eqs (9) to (13) is estimated to obtain volatility or uncertainty series. The lagged variables added to conditional mean equations in Eqs. (9) and (10) can remove autocorrelation problems in the system. The two series, Δp and Δop, are stationary as required. The model performs quite well in the dataset. The mean equation for domestic inflation rate is assumed to be dependent on the lag of domestic oil price change while the mean equation for domestic oil price change is assumed to be dependent on the inflation rate.[8] The results are reported in Table 6.

Table 6 Results from the bivariate diagonal VECH-GARCH(1,1) estimation.

Variable	Inflation	Oil price shock
A. Mean equation		
Constant	0.001***(0.000)	0.001 (0.941)
Δp_{t-1}	0.245***(0.000)	1.973** (0.036)
Δop_{t-1}	0.008***(0.009)	0.117(0.110)
B. Variance equation		
Constant	0.001** (0.013)	0.003***(0.000)
ε_{t-1}^2	0.285*** (0.000)	0.391***(0.000)
h_{t-1}	0.687***(0.000)	0.150 (0.263)
Log likelihood = 1604.666		
AIC = -10.669, SIC = -10.483		
Q(4) = 21.069 (p-value = 0.176), Adj. Q(4) = 21.251 (p-value = 0.169)		

Note: The number in parenthesis is p-value of the test statistic. ***, ** and * denotes significance at the 1%, 5% and 10%, respectively. Q(k) is the statistical test for the residuals obtained from system residual Portmanteau tests for autocorrelations, where k is the lag length.

The lags are chosen so that the system equations are free of serial correlation. Panels A and B contain the results of the conditional means and variances for inflation rate and oil price changes, respectively. Referring to Panel A, the inflation rate is positively affected by its own one-period lag and one-period lagged oil price change. In addition, oil price change is positively affected by one-period lagged inflation, but not affected by its own lag. Therefore, domestic oil price shock and inflation are interdependent. For the estimates of ARCH (ε_{t-1}^2) and GARCH (h_{t-1}) coefficients in Panel B, the coefficients in the two conditional variance equations are non-negative. The conditional variance equation for inflation has significant ARCH and GARCH terms. The sum of the two coefficients for the conditional variance of the inflation rate is 0.972.[9] The coefficients of the ARCH and GARCH terms in the oil price change conditional variance are also positive, but the coefficient of the GARCH term is insignificant. The sum of the coefficients for the conditional variance of oil price change is 0.342. These results show that the GARCH variance series as measures of volatility or uncertainty is stationary. The system diagnostic test using the residual portmanteau test for autocorrelation accepts the null of no autocorrelation as indicated by the Q(4) and adjusted Q(4) statistics. Therefore, the system equations are free of serial correlation because the Box-

[8] Even though the country is a small oil-importing country, its inflation rate should not affect the world oil price. However, the oil price series is converted to the real domestic oil price. Therefore, it is possible that inflation and oil price shocks will be interdependent.

[9] It should be noted that the conditional variance of inflation exhibits volatility persistence because the coefficient of the GARCH term is large.

Pierce statistics show that there is no serial correlation up to 4 lags, which is larger than the system lag of 1. The volatility series are generated to examine their impacts on inflation, the oil price shock and their volatility series in the standard Granger causality test. The results of Granger causality tests are reported in Table 7.

Table 7 Results of Granger causality test.

Null hypothesis	F-statistic	p-value
Δop does not cause Δp	9.175***(+)	0.003
Δp does not cause Δop	0.886 (+)	0.347
$h^{\Delta p}$ does not cause Δp	5.838**(-)	0.016
Δp does not cause $h^{\Delta p}$	7.769***(-)	0.006
$h^{\Delta op}$ does not cause Δp	2.487 (-)	0.116
Δp does not cause $h^{\Delta op}$	22.741***(-)	0.000
$h^{\Delta p}$ does not cause Δop	16.164***(-)	0.000
Δop does not cause $h^{\Delta p}$	0.788 (-)	0.375
$h^{\Delta op}$ does not cause Δop	2.198 (+)	0.139
Δop does not cause $h^{\Delta op}$	4.261**(-)	0.040
$h^{\Delta op}$ does not cause $h^{\Delta p}$	0.002 (+)	0.882
$h^{\Delta p}$ does not cause $h^{\Delta op}$	7.961***(+)	0.005

Note: Δp and Δop stand for inflation and oil price shocks, respectively. The conditional variances, $h^{\Delta p}$ for inflation rate and $h^{\Delta op}$ for oil price shocks. ***, ** and * denotes significance at the 1%, 5% and 10% level, respectively. The + sign indicates positive causation while the – sign indicates negative causation. The lag length in the pairwise causality test is 1 determined by SIC.

The important results of causality tests suggest that inflation is positively affected by oil price shocks, but it is not affected by oil price volatility in the short run. On the contrary, inflation does not cause oil price shock, but it negatively causes oil price volatility to decrease. Furthermore, inflation also causes inflation uncertainty to decrease, and thus this evidence does not support Friedman's (1977) hypothesis, which postulates that a higher inflation rate should raise inflation uncertainty and thus reduce output growth. The results also show that inflation uncertainty negatively causes inflation, which is contradictory to Cukierman and Meltzer (1986)'s hypothesis. This negative relationship between inflation uncertainty and inflation might result from the action of an independent central bank (Holland, 1995). If the central bank is independent, it will decrease the inflation rate when inflation uncertainty increases.

An oil price shock causes oil price volatility to decrease. In addition, oil price volatility tends to cause oil price shock to increase. However, oil price volatility does not cause inflation and inflation uncertainty, but inflation uncertainty causes oil price volatility to increase. In addition, inflation uncertainty causes oil price shock to decrease. Therefore, this effect can partly reduce the size of oil price shock when oil price volatility rises. Furthermore, inflation causes oil price volatility to decrease, but oil price shock does not cause inflation uncertainty. Even though inflation uncertainty does not cause oil price volatility, inflation uncertainty causes oil price shocks to decrease. Finally, oil price volatility does not cause inflation. The results seem to be complicated. The net positive impact of oil price shock on inflation suggests that the size of the negative impact of inflation uncertainty on oil price shock is relatively small.

4.3 Asymmetric or symmetric impacts of oil price shock on the inflation rate

As mentioned above, one of the important aspects of the relationship between inflation and oil price shocks is whether the short-run relationship is either symmetric or asymmetric. The asymmetric causality is tested using unrestricted VAR models and VAR Granger/block exogeneity Wald tests and the results are reported in Tables 8 and 9.

The null hypothesis to be tested is that the coefficients of lagged positive oil price shocks (op^+) and lagged negative oil price shocks (op^-) have the same positive impact on inflation. The lag length determined by SIC is 1. The results in Table 8 show that both positive and negative oil price shocks positively cause inflation. The coefficient of lagged positive oil price shock is about 0.02 and significant at the 1% level (Panel A of Table 8). However, the coefficient of negative oil price shock has a borderline significance with a value of about 0.01 (Panel B of Table 8). Therefore, the impacts of positive and negative oil price shocks do not seem to be asymmetric because the coefficient of lagged negative oil price shock is at least non-negative. This result is consistent with the evidence found by Ajmi et al. (2015) for South Africa and Farzanegan and Markwardt (2009) for Iran. The estimates of unrestricted VAR models also indicate that lagged inflation rate does not affect both positive and negative oil price shocks.

Table 8
VAR estimates

A: Positive oil price shock and inflation		
Variable	Δp_t	Δop_t^+
Δp_{t-1}	0.242***	-0.432
	(0.056)	(0.629)
Δop_{t-1}^+	0.017***	0.043
	(0.005)	(0.059)
Constant	0.001***	0.032***
	(0.000)	(0.004)
Log likelihood	1,213.641	490.587
F statistic	16.250	0.460
B: Negative oil price shock and inflation		
Variable	Δp_t	Δop_t^-
Δp_{t-1}	0.022***	0.963
	(0.060)	(0.710)
Δop_{t-1}^-	0.009*	0.262***
	(0.005)	(0.060)
Constant	0.002***	-0.024***
	(0.000)	(0.004)
Log likelihood	1,210.241	474.767
F statistic	12,654	14.767

Note: Standard error is in parenthesis. ***, ** and * indicate significance at the 1%, 5%, and 10%, respectively.

The estimates of VAR Granger causality/block exogeneity Wald tests are reported in Table 9.

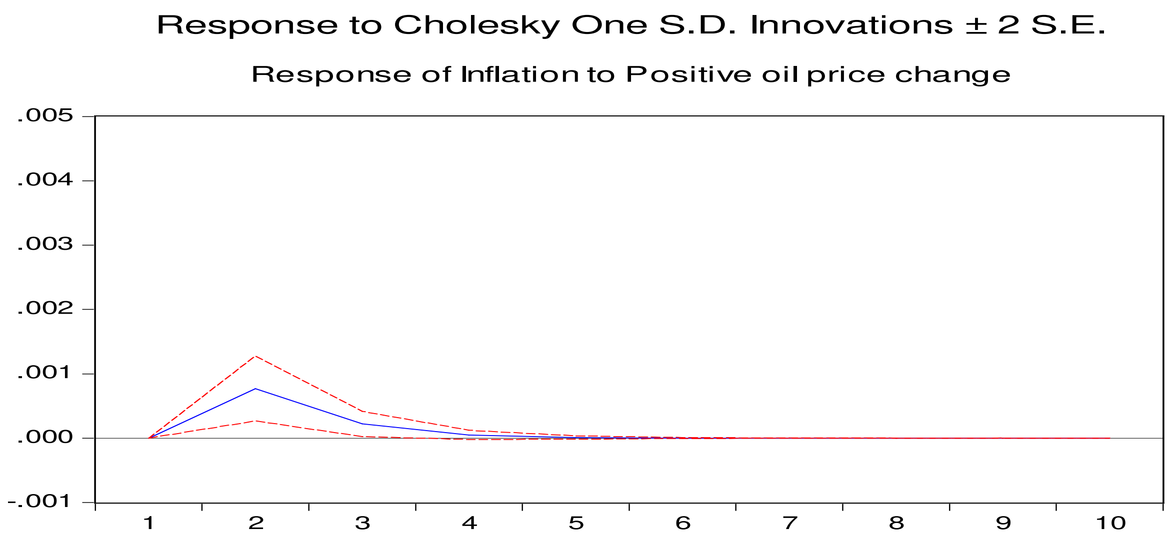
Table 9
VAR Granger causality/block exogeneity Wald tests.

Hypothesis	χ_1^2	p-value
A positive oil price shock does not cause inflation	10.267***(+)	0.001
Inflation does not cause a positive oil price shock	0.476 (-)	0.492
A negative oil price shock does not cause inflation	3.381*(+)	0.066
Inflation does not cause a negative oil price shock	1.841(+)	0.175

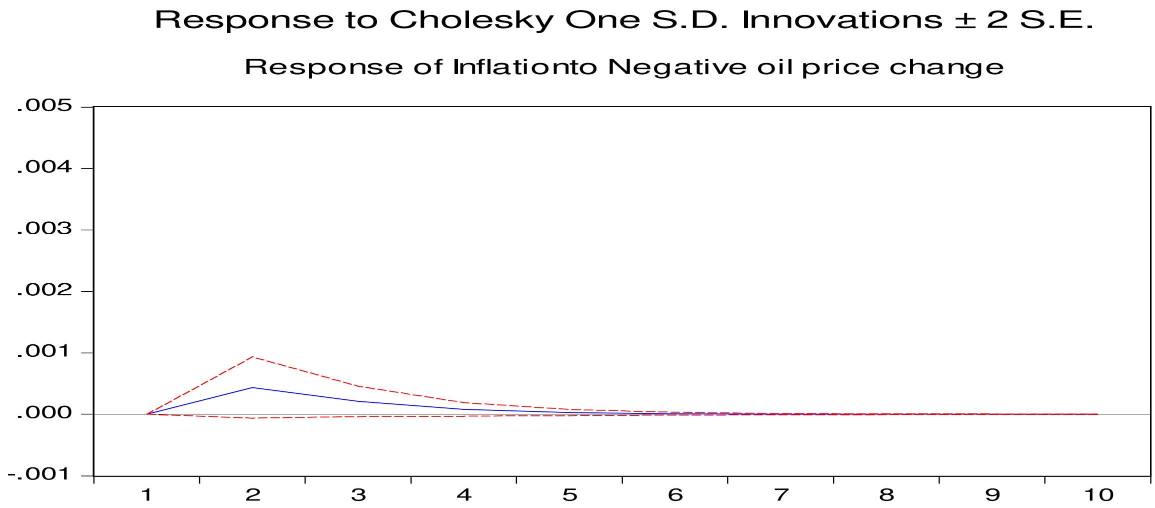
Note: ***, ** and * indicate significance at the 1%, 5% and 10%, respectively. The + sign indicates positive causation while the – sign indicates negative causation. The lag length in the pairwise causality test is 1 determined by SIC.

The results in Table 9 show that positive oil price shock causes inflation to increase because the Chi-square statistic significantly rejects the null hypothesis. However, inflation does not affect positive oil price shock. In addition, negative oil price shock marginally causes inflation to increase while inflation does not cause negative oil price shock. Therefore, asymmetric impacts of oil price shock are not evidence. This finding is in line with the finding by Cunando and de Gracia (2005), which indicates that there is no evidence of an asymmetric relationship between oil price shock and inflation rate in Thailand.

The responses of the inflation rate to positive and negative oil price shocks obtained from the estimates of unrestricted VAR models are illustrated in Fig. 3.



a. Positive oil price shock and inflation



b. Negative oil price shock and inflation

Fig. 3 Positive and negative oil price shocks and inflation

In Fig. 3a, the response of inflation to the initial positive oil price shock is significantly positive after the 1st month of the shock. The peak is reached during the 2nd month. The impact of this shock dissipates in the 3rd month. The response of inflation to negative oil price shock is similar, but the size of the impact is smaller. The results from impulse response analysis suggest that there do not seem to be asymmetric impacts of oil price shock on inflation because the coefficients of positive and negative oil price shock variables are both positive. The inflation rate seems to respond to the lagged positive and negative oil price shocks in a similar manner even though the level of significance is not the same.

4.4 Discussion

The results from the analysis are worth discussing. Previous studies find that oil price shocks affect domestic inflation. Furthermore, there is a non-linear adjustment between oil price changes and price indices. The present study uses two techniques of cointegration analysis to examine the long-run relationship between price level, industrial production, and real oil price. The presence of cointegration is not found in linear cointegration tests with unknown structural breaks. However, cointegration is found when using threshold cointegration tests that include the dummy variable to capture the structural break, which occurred few months before the 1997 Asian financial crisis. The short-run dynamics reveal that the adjustment toward long-run equilibrium is observed by the estimated parsimonious ECM. The results of the short-run analysis reveal that domestic oil price shocks Granger cause domestic inflation. In addition, oil price volatility does not cause inflation as found by Rafiq and Salim (2014). Even though oil price uncertainty does not affect inflation, inflation itself negatively causes inflation uncertainty, which does not support Friedman (1977)'s hypothesis. Furthermore, inflation uncertainty lowers the inflation rate, which is contradictory to Cukierman and Meltzer (1986)'s hypothesis. However, the impact of oil price shocks on inflation might surpass the negative impact of inflation uncertainty on inflation. Therefore, the inflation induced by oil price shocks should not be ignored by the monetary authorities. The main finding in the short run that oil price shocks cause inflation is in line with one of the main findings of Cunado and De Gracia (2005) who use quarterly data in their analyses. However, the evidence that the impacts of oil price shocks are not asymmetric is consistent with the findings of Fazanegan and Markadt (2009) and Ajmi et al. (2015).

5. Conclusion and Policy Implications

This study investigates the impact of oil price shocks on domestic inflation in Thailand. Monthly data from January 1993 to December 2016 are used. Various techniques to capture the impact of oil price shocks on inflation are employed. Both linear and nonlinear cointegration tests with structural breaks are adopted to detect the long-run relationship between price level, industrial production, and the real price of oil. In the short run, the two-step approach is also adopted to examine the impact of oil price shocks and oil price volatility on inflation. In addition, an asymmetric causality test is also used to test for asymmetric impacts of oil price shocks on inflation. The main findings are threefold. Firstly, threshold cointegration between price level, industrial production, and real domestic oil price is found in both the TAR and MTAR models. Both industrial production and real oil price have positive impacts on the price level. In addition, asymmetric adjustments toward long-run equilibrium are not found. Secondly, oil price shocks positively cause inflation, but oil price volatility does not significantly cause inflation. Furthermore, inflation itself negatively causes inflation uncertainty. Finally, the presence of asymmetric impacts of oil price shock on inflation is not found in the Thai economy. The implications based upon the results of this study are that, besides the inflation targeting that has been implemented by the monetary authorities, monetary measures should also be designed to accommodate inflation induced by oil price shocks in both the short and long run. The oil fund as a subsidization should not be discarded. Furthermore, energy policy should focus more on energy efficiency such that the inflationary pressure from oil price shocks can be alleviated.

Even though this study is limited to Thailand, it can provide insights for other oil-importing emerging market economies, especially in Asia and the Pacific region, which wish to assess the impact of oil price shocks on the inflation rate.

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