Oil price shocks and domestic inflation in Thailand

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Abstract

This paper employs monthly data to examine the impact of oil price shocks on the domestic inflation rate in Thailand from 1993 to 2017. 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 response of inflation to oil price shocks does not seem to be asymmetric. 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 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 price 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). Inflation induced by oil price shocks can reduce real balances, a measure 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). 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.

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 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; 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 Gregono and Lanerretche, 2007; 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 oil-importing countries (Olofin and Salisu, 2017). 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. Huang and Chao (2012) examine the effects of international and domestic oil prices on the price indices in Taiwan using monthly data from January 1999 to December 2011. They find that changes in international oil prices have more crucial impacts on the price indices than do changes in domestic oil prices. Chu and Lin (2013) find that oil price shocks have both long-term and short-term effects on Taiwan's producer price index.

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 Markwadt (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 even though a negative oil price shock has a stronger effect. 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 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. 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 remain in the Euro area because no deflationary effect of oil prices will result in a negative inflation rate.

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. The real price of oil is used as in the study by Cunado and Perez de Gacia (2005) who use two different definitions of oil prices. This study uses their second definition, which is the real price of oil.[1] 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. 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 presents the data and estimation methods that are used in the analysis. Section 3 presents the empirical results. Section 4 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. Data and Methodology

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

2.1 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 does 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 dollar per barrel is obtained from the US Energy Information Administration. The oil price series is international oil price. By multiplying the oil price series by the US dollar exchange rate and deflating by the consumer price index, the domestic real oil price series is obtained.[2] All series used in estimation are seasonally adjusted and transformed into logarithmic series. The sample size comprises 300 observations.

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[1] 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 various definitions of oil shocks. Also, real oil price can reflect both the true purchasing power and the cost of production.

[2] 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 of 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
Δ*p* (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
Δ*ip* (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
Δ*op* (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 bracket represent the optimal lag length determined by 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. It should be noted that the test for the level of the consumer price index with constant only and the test for the level of industrial production index with constant and linear trend seem to reject the null hypothesis, but the level of significance is only 10%. Therefore, it can be concluded that all series are I(1). This is suitable in performing cointegration tests. The stationary property of first differences of series is also suitable in the estimate of a bivariate gerneralized autoregressive conditional heteroskedastic (GARCH) model as well as the standard pairwise causality test described in the next sub-section.[3]

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

[3] 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 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 both 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.

Co-movement between price level and the real domestic oil price series is plotted in Fig. 1. Even though the real oil price is linked to the trend of price level, the real oil price variable fluctuates more. Starting from a low oil price with some fluctuations, the impact of a new oil shock in 2000 causes the price to increase. The oil price reaches its peak in mid-2008 and falls to the trough at the end of 2008, the 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.[4]

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

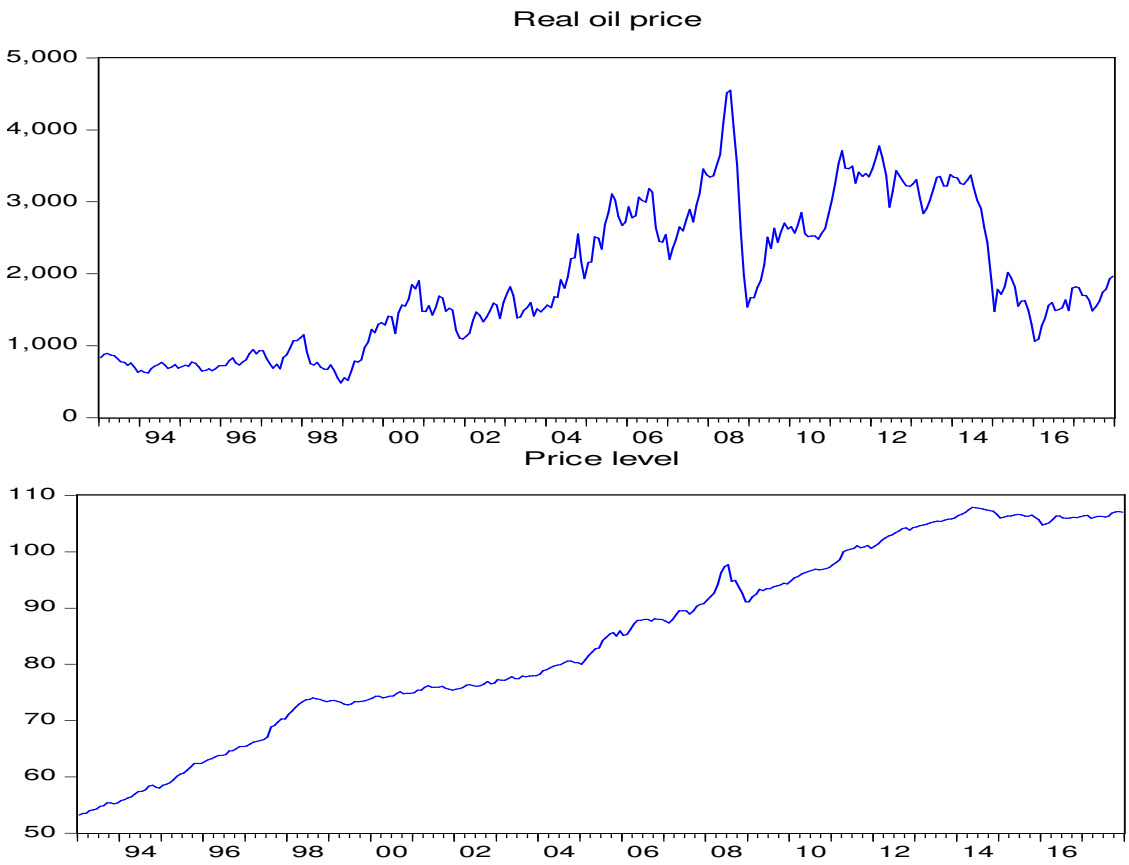

Fig. 1 Co-movement of price level with real oil price.

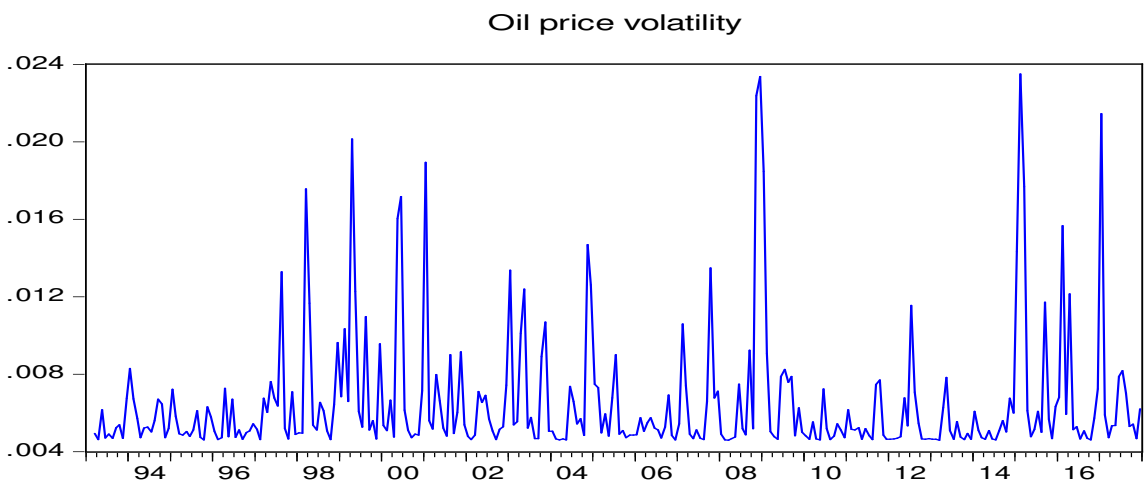

Fig. 2 Volatility of real oil price.

2.2 Estimation Methods

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

2.2.1 Cointegration tests

The existence of cointegration between price level and real oil price implies that there is a long-run relationship between the two variables in a bivarite framework. However, industrial production can interact with both price level measured by consumer price index and real oil price. Therefore, trivariate cointegration analysis can be used to test whether there is positive long-run relationship between price level and the real domestic oil price when industrial production is treated as a control variable.

2.2.1.1 Residual-based cointegration tests with unknown breakpoints

Similar to conventional residual-based test for cointegration, this test proposed 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 consumer price index, domestic oil price and industrial production. However, Gregory-Hansen procedure takes into account the impact of unknown level shift as well as regime shift. 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 coefficient b_2 and b_3 should be statistically significant. The dummy variable, D_t, captures an unknown break point. The residual series, e_t, obtained from the estimations of Eq. (1) and can be used to test for 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 models.

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 the

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.

2.2.1.2 Nonlinear cointegration tests

It is important to confirm that the relationship between variables is not nonlinear. In case of the absence of linear cointegration between variables, it is possible that the long-run relationship is nonlinear and asymmetric. 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 endogenously 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. 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 pertaining to the short-run relationship between price level, industrial production and real oil price. The coefficients of the lagged differences for industrial production and for 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 price level and industrial production, and price level and real oil price.

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

The two-step approach is employed to explain the relationship between nominal 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 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 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), comprises 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 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 causality tests.

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[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.

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

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

One of the important issues concerning short-run relationship between inflation and oil price shock is asymmetric effects of oil price increases and decreases on inflation rate. Following Mork (1989)'s procedure, 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:

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

and

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

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

and

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

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.

3. Empirical Results

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

3.1 Long-run relationship and short-run dynamics

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

The results in Table 3 show that there is no cointegration 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.

The results of Gregory-Hansen cointegration tests suggest that there is no linear cointegration between variables in the model. The relationship between variables in Eq. (1) is reported in Table 3.

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% level, 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 manufacturing firms, which are energy intensive, the reduction of oil use 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 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 break point at 1997M4, which is near the occurrence of 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 (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 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 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 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 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 (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 long-run equilibrium is 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% level, respectively.

Since SIC indicates that the lag should be zero, a parsimonious ECM with lag of 1 is estimated. The estimated short-run equation passes important diagnostic tests, i.e., the no serial correlation and no further ARCH effect 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

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

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).

3.2 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.

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) to is estimated to obtain volatility or uncertainty series. The lagged variables added to conditional mean equations in Eqs. (9) and (10) can remove autocorrelation problem 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 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***(0.001)	0.007 (0.013)
Δp_{t-1}	0.236**(0.097)	0.088 (2.853)
Δop_{t-1}	0.009**(0.003)	0.107*(0.067)
B. Variance equation		
Constant	0.001 (0.001)	0.004***(0.001)
ε_t^2	0.249** (0.123)	0.272**(0.114)
h_t	0.723***(0.061)	0.070 (0.171)
Log likelihood = 1592.193		
AIC = -10.592, SIC = -10.418		
Q(4) = 14.465 (p-value = 0.564), Q(8) = 40.222 (p-value = 0.151)		

Note: Standard errors are given in parenthesis. ***, ** 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 not affected by its one-period lag and one-period lagged inflation. Therefore, domestic oil price

[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 real domestic oil price. Therefore, it is possible that inflation and oil price shocks will be interdependent.

shock and inflation are not interdependent because only lagged oil price shock positively affects inflation. As for the estimates of ARCH (ε_t^2) and GARCH (h_t) coefficients in Panel B, the coefficients in the two conditional variance equations are non-negative. The conditional variance equation for inflation gives significant ARCH and GARCH terms. The sum of the two coefficients for the conditional variance of 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 residual portmanteau test for autocorrelation accepts the null of no autocorrelation as indicated by the Q(4) and Q(8) statistics. Therefore, the system equations are free of serial correlation. The volatility series are generated to examine their impacts on inflation and volatility in the standard Granger causality test. The results of a pairwise Granger causality tests are reported in Table 7.

Table 7 Results of pairwise Granger causality test

Hypothesis	F-statistic	p-value
Δop does not cause Δp	9.175***(+)	0.003
Δp does not cause Δop	0.886** (+)	0.347
Δop does not cause $h^{\Delta op}$	11.889***(-)	0.001
$h^{\Delta op}$ does not cause Δop	1.817 (+)	0.178
$h^{\Delta op}$ does not cause $h^{\Delta p}$	0.232 (+)	0.631
$h^{\Delta p}$ does not cause $h^{\Delta op}$	8.908*** (-)	0.003
$h^{\Delta p}$ does not cause Δop	10.451***(+)	0.000
$h^{\Delta p}$ does not cause Δp	5.576**(-)	0.019
Δp does not cause $h^{\Delta op}$	27.723***(-)	0.000
Δop does not cause $h^{\Delta p}$	0.934 (-)	0.335
$h^{\Delta p}$ does not cause Δop	14.862***(-)	0.000
$h^{\Delta op}$ does not cause Δp	2.169 (-)	0.142

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 (1977) hypothesis, which postulates that 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) hypothesis. This negative relationship between inflation uncertainty and inflation might results from the action of an independent central bank (Holland, 1995). If the central bank is independent, it will decrease inflation rate when inflation uncertainty increases.

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

An oil price shock causes higher oil price volatility. In addition, oil price volatility tends to cause oil price shock to decrease. However, oil price volatility does not cause 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.

The short-run impact of lagged real oil price change on inflation is found from the symmetric adjustment process reported in Table 5. However, VAR specifications can also be analyzed to examine how one variable interact with other variables in the short run (Sims, 1980). An unrestricted VAR system with four variables is estimated to examine how one variable interacts with other variables. In addition, reporting IRFs without standard error bands is the same as reporting regression without t-statistics (Runkle, 1987; Sims and Zha, 1999). The IRFs are shown in Figure 3. The optimal lag determined by SIC is 1. The IRFs are based on Monte Carlo confidence bands with 500 simulations. The period is within 10 months.

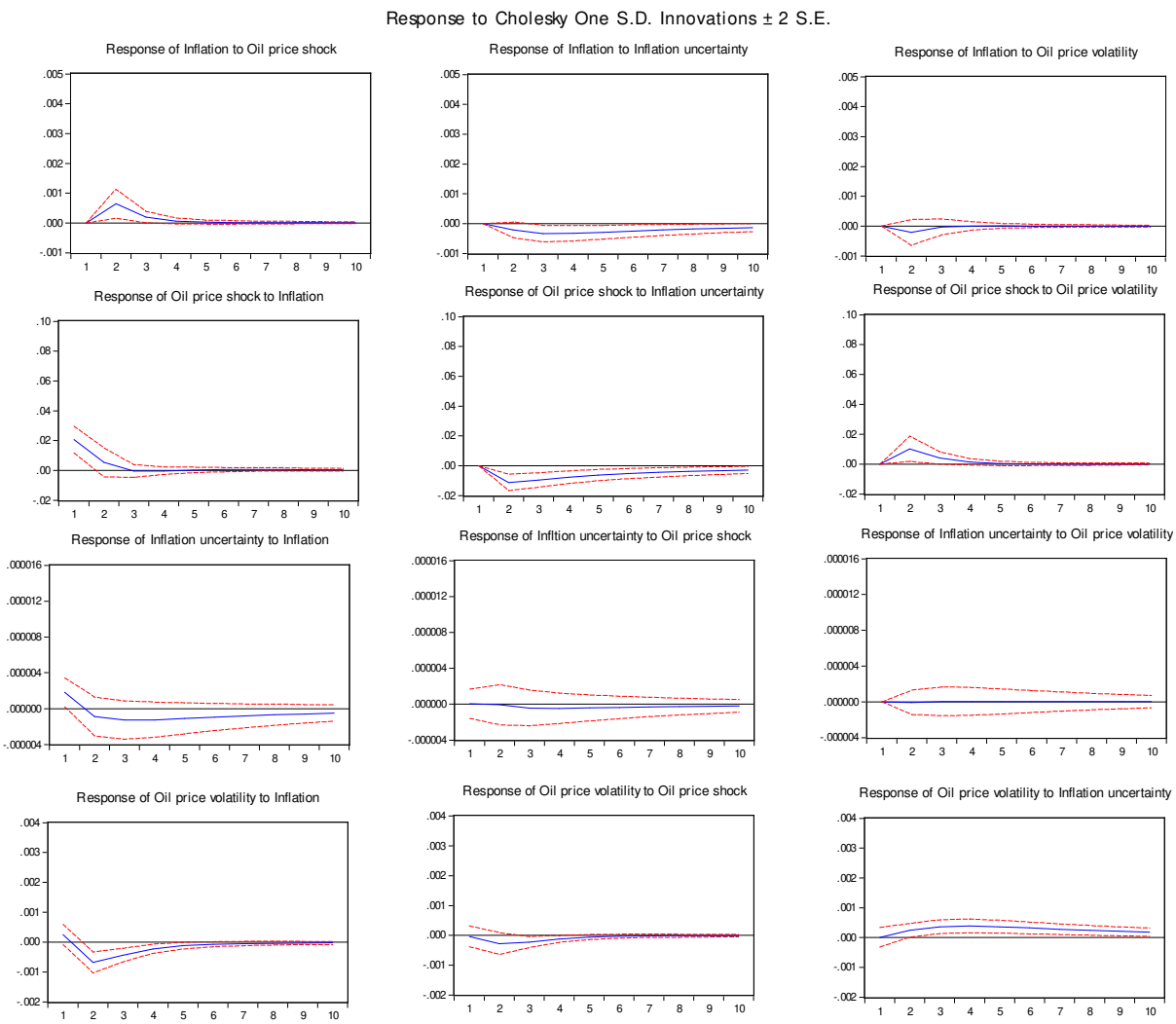


Fig. 3 Impulse response functions including volatility series.

The estimate of VAR(1) model allows for performing an analysis of IRFs and VDCs. The results of impulse response analysis are shown in Fig. 3. The figure shows the IRFs from the Monte Carlo simulated at 95 percent intervals. The response of inflation rate (Δp) to a shock in oil price (Δop) shows that inflation significantly increases in the next month and starts to subside after the 2nd month. The effect of this shock dissipates in the 4th month. The response of inflation to a shock in oil price volatility ($h^{\Delta OP}$) shows that inflation decreases until the 4th month and starts to recover and is incorporated in the 10th month. However, this impact is not significant because the upper band touches the horizontal line. The response of inflation to a shock in oil price volatility starts in the 2nd month, but the impact of oil price volatility is not significant because the horizontal line is between the two bands. For the oil price shock variable, the response of a shock in oil price to inflation starts in the next month, i. e., inflation has a significantly positive impact on the real price of oil, but decays and is incorporated in three months. Oil price shocks respond negatively to inflation uncertainty in the 2nd month and the impact subsides and is incorporated in the 6th month. On the contrary, oil price shocks respond positively to oil price volatility. The positive impact occurs in the 1st month and starts to decay later on. This impact is not significant because the lower band touches the horizontal line. The response of inflation uncertainty to inflation is positive but becomes negligible after the 2nd month while its slight impact on oil price shocks is negative and not significant. The response of inflation uncertainty to oil price volatility is negative, but the impact is very slight and not significant. As for oil price volatility, this variable responds negatively to inflation, and this significant response lasts for six months. The response of oil price volatility to its own shocks is insignificantly negative. Finally, the positive response of oil price volatility to inflation uncertainty is significant after the 2nd month and subsides within peak in six months. Even though the impact subsides later on, it never dissipates.

VDCs shown in Table 8 can be used to ascertain how important the innovations of other variables are in explaining the fraction of each variable at different step ahead forecast variances. The results of this analysis provide evidence for the independency of oil price shock and other variables. An oil price shock has a significantly positive impact on inflation and inflation uncertainty. Furthermore, oil price volatility has a slight impact on inflation, but no impact on inflation uncertainty.

Inflation explains only its own variances in the first month. It explains approximately 2% of the variances of oil price shocks in the 2nd month and only 2% and less than 1% of inflation uncertainty and oil price volatility, respectively. The oil price shock variable explains about 7% of the variances of inflation, but it explains 4% the variances of inflation uncertainty in the 4th month and 2% of oil price volatility in the 4th month. Inflation uncertainty variable explains less than 2% of the variances of inflation for the period of 10 months, and it explains less than 1% of the variances of oil price shock, and does not explain the variances of oil price volatility for the whole period. Finally, oil price volatility almost explains its own variances in the first month. It explains 7% of the variances of inflation in the 10th month and only 1.4% of the variance of oil price shock while it explains 7% of the variances of inflation uncertainty.

Table 8
Variance decompositions of Δp, Δop, $h^{\Delta p}$, and $h^{\Delta op}$.

Variance decomposition of Δp				
Month	Δp	Δop	$h^{\Delta p}$	$h^{\Delta op}$
1	100.00	0.00	0.00	0.00
2	97.25	2.25	2.43	0.25
4	95.92	2.44	1.38	0.25
10	94.59	2.41	2.74	0.25
Variance decompositions of Δop				
1	6.80	93.20	0.00	0.00
2	6.94	89.54	1.91	1.61
4	6.77	87.31	4.07	2.25
10	6.67	85.79	5.72	1.82
Variance decompositions of $h^{\Delta p}$				
1	1.61	0.00	98.39	0.00
2	1.18	0.00	98.82	0.00
4	1.40	0.08	98.52	0.00
10	1.60	0.15	98.25	0.00
Variance decompositions of $h^{\Delta op}$				
1	0.81	0.15	0.00	99.17
2	5.50	0.82	0.63	93.05
4	7.47	1.41	3.33	84.78
10	7.36	1.40	7.31	83.92

Note: Δp is inflation rate, Δop is oil price shocks, $h^{\Delta p}$ is inflation uncertainty, and $h^{\Delta op}$ is oil price volatility.

3.3 Asymmetric or symmetric impacts of oil price shock on 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 an unrestricted VAR models and VAR Granger/block exogeneity Wald tests and the results are reported in Tables 9 and 10.

The null hypothesis to be tested is that the coefficients of lagged positive oil price shocks (op^+) and lagged negative oil price shocks (op^-) are the same. The lag length determined by SIC is 1. The results in Table 9 show that both positive and negative oil price shocks positively causes inflation. The coefficient of lagged positive oil price shock is about 0.02 and significant at the 1% level (Panel A of Table 9). However, the coefficient of negative oil price shock has a borderline significance with the value of about 0.01 (Panel B of Table 9). Therefore, the impacts of positive and negative oil price shocks are not 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. The estimates of unrestricted VAR models also indicate that lagged inflation rate does not affect both positive and negative oil price shocks.

Table 9
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 10.

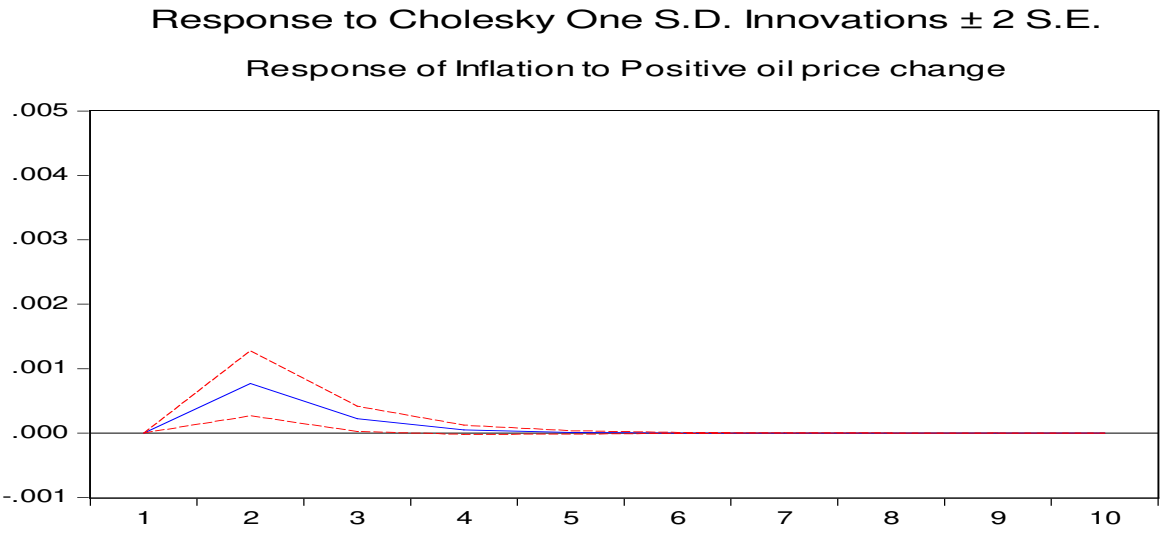
Table 10
VAR Granger causality/block exogeneity Wald tests.

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

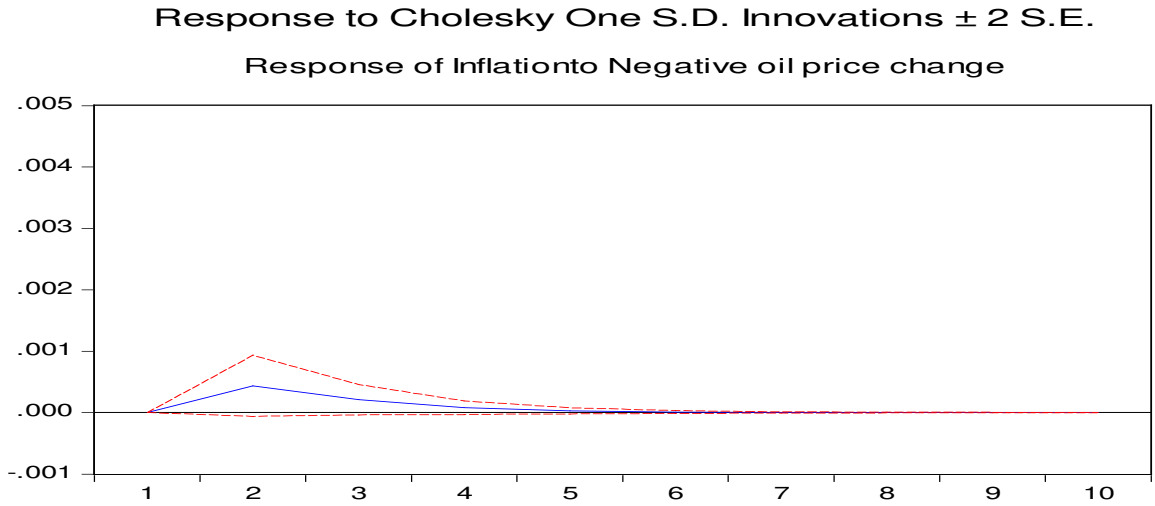
Note: ***, ** and * indicate significance at the 1%, 5% and 10%, respectively. The + sign indicates positive causation while the – sign indicates negative causation. The lag length in the pairwise causality test is 1 determined by SIC.

The results in Table 10 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 inflation rate to positive and negative oil price shocks obtained from the estimates of unrestricted VAR models are illustrated in Fig. 4.



a. Positive oil price shock and inflation



b. Negative oil price shock and inflation

Fig. 4 Positive and negative oil price shocks and inflation

In Fig. 4a, the response of inflation to 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.

4. Discussion

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 a threshold cointegration test that includes the 1997 Asian financial crisis dummy variable. The short-run dynamics reveal that the adjustment toward long-run equilibrium is observed only in the regime below the threshold value even though this adjustment is likely to be symmetric at the 5% level of significance. The results of short-run analysis reveal that domestic oil price shocks Granger cause domestic inflation and this result is contradictory to Huang and Chao (2012) who find that international oil price plays a more important role than does domestic oil price on price indices. 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 positively causes inflation uncertainty, which supports Friedman (1977)'s hypothesis. On the contrary, 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. Concluding Remarks 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, one threshold cointegration between price level, industrial production and real domestic oil price is found in the threshold autoregressive model. Both industrial production and real oil price have positive impacts on price level. In addition, asymmetric adjustments toward long-run equilibrium are found at the low level of significance. Secondly, oil price shocks positively cause inflation, but oil price volatility does not significantly cause inflation. Furthermore, inflation itself positively causes inflation uncertainty. This finding is also confirmed by impulse response analysis and variance decompositions. 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. The oil fund as subsidization should not be discarded. Furthermore, energy policy should focus more on energy efficiency such that 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 in the Asia and the Pacific region, which wish to assess the impact of oil price shocks on inflation rate.

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