The Asymmetric Effect of Oil Price on Growth across US States

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Abstract

This paper investigates the dynamic relationship between oil prices and growth across the U.S. States using a panel data framework. We use both annual and quarterly data spanning the periods 1973 to 2013 and 1948Q1 to 2013Q4, respectively. Following Hatemi-J (2012), we allow for the presence of asymmetry in the cointegration and causality testing by decomposing oil prices into cumulative sums of positive and negative oil prices. The null hypothesis of no cointegration is rejected. The long-run coefficients are found to be statistically significant across all empirical models, with positive oil prices reducing output, while negative oil prices increasing output. We also find evidence of both short- and long-run bidirectional causality between aggregate oil prices and output. However, there is evidence of unidirectional causality both from positive and negative oil prices to output based on annual data. The quarterly data generated slightly different result, indicating both long- and short-run bidirectional causality between positive and negative oil prices and output.

JEL Codes: C33, O00, Q41, R10

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Introduction

The relationship between oil price changes and the macroeconomic performance of national economies is an important subject that has been examined for the majority of industrialized countries, especially after the 1973 oil supply shock. The presence of a negative link between oil prices and macroeconomic activity has become generally accepted since Hamilton's (1983) pioneering work, demonstrating that oil price increases slowed down US output growth over the period 1948-1980.

Hamilton's results have been verified and extended both theoretically and empirically by numerous researchers. For instance, a distinction has been made between the impact of oil prices on both oil importing and oil exporting countries. Oil price increases are expected to have a positive effect on oil exporting countries, since this would have an increasing effect on their income. As income increases, both consumption and investment expenses are expected also to increase, leading to higher levels of productivity and lower unemployment (Bjornland, 2009; Jimenez-Rodriguez and Sanchez, 2005). In the case of an oil-importing country, oil price increases will have a negative effect (LeBlanc and Chinn, 2004; Hooker 2002), given that oil is considered to be a critical factor of production; hence, higher oil prices are expected to lead to higher cost of production (Filis et al., 2011; Arouri and Nguyen, 2010; Backus and Crucini, 2000; Kim and Loungani, 1992). These higher costs will eventually reach the consumers via higher consumer prices, resulting in lower levels of demand and consumer spending, lower production and higher unemployment (Bernanke, 2006; Lardic and Mignon, 2006; Brown and Yücel, 2002; Davis and Haltiwanger, 2001; Hamilton 1988).

A hot debate in the literature on the effect of oil prices on the macroeconomy is whether the structural responses of real economic variables triggered by both positive and negative oil price innovations are asymmetric. The relevant question is whether a positive oil price shock produces larger effects than an equal-sized negative oil price shock. Three explanations of asymmetries have been provided in the literature: counter inflationary monetary policy, sectoral shocks, and uncertainty (Sadorsdy, 1999; Kilian and Vigfusson, 2011a). However, only the latter two are strongly grounded by economic arguments, which are explained here. The first is that oil price shocks are relative price shocks that can be viewed as allocative disturbances causing sectoral shifts throughout the economy (Hamilton 1988). Using a multi-sector model, Hamilton

(1988) shows that it is costly to shift labour and capital inputs between sectors due to labour mobility and training costs. For instance, relative oil price shocks can cause consumers to reduce their demand for energy-intensive durables, such as automobiles. The dollar value of such purchases may be large relative to the value of the energy they use. If labour and capital can easily be moved to new areas, most of the output loss in the energy-intensive sectors will be regained from other, less adversely affected, sectors. However, if labour and capital are sector or product specific and, hence, costly to move, the losses of one sector may not necessarily be regained by the other sectors as capital and labour in these sectors will remain unemployed, leading to lower output levels. The same would happen if unemployed workers rationally chose not to relocate but rather wait for conditions in their sectors to improve.

The second explanation relates to the literature on irreversible investment under uncertainty. To the extent that unexpected changes in oil prices are associated with increased uncertainty about future oil prices, while cash flows from an irreversible investment project depend on oil prices, the real options theory implies that, all other things being equal, increased oil price uncertainty is expected to make firms delay their investments decisions, causing investment expenditures to decline (Bernanke 1983; Pindyck 1991). This uncertainty effect is expected to amplify the impact of unexpected oil price increases and offset the effects of unexpected oil price declines.

On the empirical front, there is a large number of studies that investigate the link between oil prices and output with a resuscitated interest, especially when oil prices reached 147 USD per barrel in August 2008. Earlier studies include that by Burbidge and Harrison (1984) who investigate the impact of oil price shocks on a number of macroeconomic variables for five developed countries (i.e., the U.S., Japan, Germany, the U.K. and Canada) using the methodology of VAR modeling. They document that the 1973-74 oil embargo event can explain a considerable part of the behaviour of industrial production in each of the countries under investigation and they find little evidence that changes in oil prices can have an effect on industrial production. Gisser and Goodwin (1986) display that oil price effects on the U.S. economy do not change after 1973, when the OPEC contributed to the first significant world oil shock. Using the VAR modeling approach, Mork (1989) examines the asymmetric response of the U.S. real GDP growth to oil price changes by decomposing oil price changes in real price increases and decreases. His analysis provides evidence that positive changes in real oil prices have a far more pronounced effect on real output than a negative shock. Mork et al. (1994) confirm the presence of the asymmetric effect for other OECD countries, noting, however, that oil price increases slow down economic growth in the U.S. more vis-à-vis other countries, i.e. Germany, France, and Japan, which are countries more dependent on imported oil.

More recent studies include that by Zhang (2008) who employs a flexible nonlinear approach to examine the relationship between oil prices and economic growth in Japan; he finds evidence that the relationship between oil price changes and macroeconomic activities are nonlinear as well as asymmetric. He also documents the presence of unidirectional causality running from oil prices to output. Lardica and Mignon (2008) employ an asymmetric cointegration approach and investigate the long-run relationship between oil prices and GDP for the G7 economies plus the U.S. and Euro area economies and find the presence of an asymmetric cointegration, which could not be found in standard cointegration tests. Cologni and Manera (2009) investigate the effect of oil price shocks on the output growth rate in the case of the G7 countries by comparing alternative regime switching models. They find a reverse relationship between the two series, albeit the impact of oil prices has changed over time. Balke et al. (2010) examine the various sources of oil price shocks and economic fluctuations and assess their effects on the U.S. economic activity by employing Bayesian methods through a dynamic stochastic general equilibrium model. They conclude that changes in oil prices are endogenous, with the oil price shocks in the 1970s, early 1980s and the 2000s reflecting differing mixes of shifts in oil supply and demand and that different sources of oil price shocks have different effects on economic activity.

Rahman and Serletis (2010) examine the asymmetric effects of oil price shocks and monetary policy on output using monthly data for the U.S. Evidence from a logistic smooth transition vector autoregression model indicates that oil prices along with their volatility have a substantial effect on output, while the response of output to those oil price shocks is asymmetric. The asymmetric effect of oil price uncertainty on output is also confirmed by Elder and Serletis (2010) who make use of a bivariate GARCH-in-mean model. In contrast, Kilian and Vigfusson (2011b) investigate the same relationship for the U.S. through impulse responses coming from an unrestricted structural model that encompasses both linear symmetric and asymmetric models. They find that the response of output to positive oil price shock (0.47% decline) is roughly of the same magnitude (with that of a negative shock (0.39% increase) . Overall, their findings suggest no evidence against the null of symmetric response functions.

Although the macroeconomic research literature has focused mainly on the role of oil prices in the case of developed countries, some recent studies have examined the same role for the case of developing and emerging countries as well. Farzanegan and Markwardt (2009) investigate the dynamic relationship between oil price shocks and major macroeconomic variables in Iran by applying a VAR approach. They point out the presence of asymmetric effects of oil price shocks and find a strong positive relationship between positive oil price changes and industrial output growth. Berument et al. (2010) investigate how oil price shocks affect output growth in selected MENA countries; they conclude that oil price increases have a positive effect on production in the cases of Algeria, Iran, Iraq, Kuwait, Libya, Oman, Qatar, Syria, and the United Arab Emirates. Iwayemi and Fowowe (2011) explore the effect of oil price shocks on the economy of Nigeria which is a developing oil exporter country. They provide supportive evidence that different measures of linear and positive oil shocks cannot cause output growth. Other studies in this strand of the literature examine the impact of oil prices in Kenya (Semboja, 1994), in twelve developing countries (Abeysinghe, 2001), in Korea (Glasure, 2002; Huntington, 2004), in a number of Middle East countries (Narayan and Smyth, 2007), in South Africa (Ziramba, 2010; Aye et al., 2014), and in China (Tang et, al., 2010).

This empirical study contributes to the large volume of the literature on the relationship between oil prices and economic growth by conducting a regional analysis of the relationship for the case of the U.S. states. Specifically, it examines the cointegration relationship as well as the long- and short-run causal relationship between oil prices and output across U.S. states. Furthermore, an additional novelty of the paper is the investigation of potential asymmetric effects of oil prices on regional growth rates. The only study, to the best of our knowledge, that is close to our analysis is that by Kang et al. (2011), which investigates the impact of oil price changes on U.S. States using a regime-switching model. Their empirical findings highlight that the tolerance and speed of response to oil price changes vary across States. However, their methodology restricts the regional analysis to only 13 States. By contrast, in our study all States are included in the panel analysis, while analytical tests of cross-sectional dependence reveal that the panel-based approach is more appropriate than a state-by-state estimation. In addition, our study decomposes oil price shocks into cumulative sums of positive and negative oil prices that allow us to test for asymmetric effects both in the cointegration and in the causality model. To the best of our knowledge, this is the first attempt that analyses both short- and long-run effects relative to the impact of oil prices on the U.S. economic growth across all States, which also decomposes oil price shocks into their positive and negative components.

Data and methodology

We use real personal income as a proxy for state output. These nominal personal income data are obtained from the regional account database of the US Census Bureau, while real values are derived by deflating the nominal values by the US consumer price index, given that consistent data for state-level consumer price indexes (CPIs) are not available for the entire sample period; data were obtained from the FRED database of the St. Louis Federal Reserve Bank. Our oil price measure is that of the West Texas Intermediate (WTI), obtained from the Global Financial database. We used annual data spanning the period 1929 to 2013. As a robustness check we also conducted the analysis using quarterly data spanning the period 1948Q1 to 2013Q4. Understandably, start and end-points of the sample are purely data-driven. We only considered the 48 contiguous US States, by leaving out of the analysis the Alaska and Hawaii, as data for these two states only starts from 1950.

We decomposed oil prices into their cumulative sums of positive and negative components using the method developed by Hatemi-J (2012). The empirical model takes the following alternative forms:

$$ry_{it} = \alpha_{i1} + \beta_1 npo_t + \varepsilon_{1it} \tag{1}$$

or

$$ry_{it} = \alpha_{i2} + \beta_2 rpo_t + \varepsilon_{2it} \tag{2}$$

or

$$ry_{it} = \alpha_{i3} + \beta_3 npo_t^{+} + \beta_4 npo_t^{-} + \varepsilon_{3it} \tag{3}$$

or

$$ry_{it} = \alpha_{i4} + \beta_5 rpo_t^{+} + \beta_6 rpo_t^{-} + \varepsilon_{4it} \tag{4}$$

where $i = 1,..., \mathrm{N}$ for each U.S. State in the panel and $t = 1,..., \mathrm{T}$ refers to the time period. ry denotes real income, npo is nominal oil prices, rpo denotes real oil prices, $npo^{+}(npo^{-})$ is the cumulative sum of positive (negative) nominal oil prices, and $rpo^{+}(rpo^{-})$ denotes the cumulative sum of positive (negative) real oil prices. More specifically, we use *po* as a generic term to indicate nominal (*npo*) or real oil prices (*rpo*), defined as a random walk process provided in (5):

$$po_t = po_{t-1} + u_{1t} = po_0 + \sum_{i=1}^{t} u_{1i} = po_0 + \sum_{i=1}^{t} u_{1i}^{+} + \sum_{i=1}^{t} u_{1i}^{-} \quad (5)$$

where $t = 1,2,\ldots,T$, po_0 is a constant representing the initial value of po_t, and u_{1i} indicates a white noise error term defined as the sum of positive and negative shocks, i.e., $u_{1i} = u_{1i}^{+} + u_{1i}^{-}$, where $u_{1i}^{+} = \max(u_{1i}, 0)$ and $u_{1i}^{-} = \min(u_{1i}, 0)$. Hatemi-J (2012) defines positive and negative shocks in a cumulative form as: $po^{+} = \sum_{i=1}^{t} u_{1i}^{+}$ and $po^{-} = \sum_{i=1}^{t} u_{1i}^{-}$. The parameters α_{is} allow for the presence of state-specific fixed effects. Finally, ε_{sit} denotes the estimated residuals, representing deviations from the long-run relationship. According to theoretical expectations, the signs of the estimated coefficients are expected to be: $b_1<0$, $b_2<0$, $b_3<0$, $b_4>0$, $b_5>0$ and $b_6>0$.

Empirical results

Panel and time series unit root tests

There are a variety of panel unit root tests, which include Levin and Lin (1993), Harris and Tzavalis (1999), Maddala and Wu (1999), Breitung (2000) and Hadri (2000). The results in Table 1 point out that all seven panel variables in both on an annual and quarterly basis are characterized as I(1) variables.

Table 1. Panel and time series unit root tests.

Variables	LL	Han (hom)	Han (het)	F-ADF	F-PP	HT	Breit
Annual							
ry	-1.62	23.25*	22.45*	16.61	17.62	-1.16	-1.23
Δry	-5.28*	1.39	1.26	114.25*	110.51*	-6.52*	-4.66*
npo	-1.15	39.45*	49.18*	19.68	19.08	-1.37	-1.58
Δnpo	-6.47*	1.24	1.55	105.57*	104.24*	-6.07*	-5.67*
rpo	-1.27	48.62*	44.57*	18.71	17.51	-1.26	-1.09

Δrpo	-6.41*	1.16	1.36	107.61*	147.11*	-6.28*	-5.95*
npo^+	-1.29	47.73*	35.57*	17.74	18.50	-1.28	-1.26
Δnpo^+	-6.15*	1.24	1.62	102.62*	137.73*	-6.21*	-5.84*
npo^-	-1.36	48.44*	43.59*	17.38	19.09	-1.47	-1.37
Δnpo^-	-5.17*	1.32	1.27	110.94*	112.54*	-5.35*	-5.85*
rpo^+	-1.49	43.61*	42.49*	16.71	19.41	-1.09	-1.38
Δrpo^+	-6.08*	1.26	1.30	106.52*	105.07*	-6.37*	-5.71*
rpo^-	-1.33	41.45*	43.58*	15.05	15.52	-1.49	-1.08
Δrpo^-	-5.82*	1.28	1.26	109.21*	109.33*	-6.51*	-5.35*
Quarterly							
ry	-1.25	28.72*	28.49*	18.24	21.24	-1.24	-1.36
Δry	-5.71*	1.33	1.28	125.61*	131.25*	-6.26*	-4.52*
npo	-1.24	34.62*	43.64*	17.77	20.68	-1.46	-1.46
Δnpo	-6.16*	1.22	1.27	116.52*	134.47*	-6.38*	-5.49*
rpo	-1.38	45.37*	40.76*	19.16	25.36	-1.51	-1.14
Δrpo	-6.29*	1.24	1.52	114.24*	132.52*	-6.85*	-5.57*
npo^+	-1.41	49.09*	39.79*	18.43	23.45	-1.23	-1.42
Δnpo^+	-6.48*	1.29	1.36	114.23*	139.68*	-6.36*	-5.48*
npo^-	-1.55	46.38*	42.94*	18.72	21.12	-1.30	-1.17
Δnpo^-	-5.72*	1.46	1.83	116.53*	118.69*	-5.58*	-5.90*
rpo^+	-1.29	48.72*	44.83*	19.27	19.36	-1.13	-1.29
Δrpo^+	-6.51*	1.37	1.39	116.39*	114.937*	-6.46*	-5.46*
rpo^-	-1.52	43.62*	44.92*	24.48	18.26	-1.35	-1.23
Δrpo^-	-5.97*	1.37	1.42	121.84*	128.61*	-6.18*	-5.58*

Notes: Δ denotes first differences. LL denotes the Levin and Lin test, Han denotes the Hadri test, F-ADF and F-PP denotes the Maddala and Wu test, HT denotes the Harris and Tzavalis test, Breit denotes the Breitung test.
* denotes statistical significance at 1%

Unit roots and structural breaks

Zivot and Andrews (1992) propose a testing procedure where the time of the break is estimated, rather than assumed as an exogenous phenomenon. The null hypothesis in their method is that the variable under investigation contains a unit-root with a drift that excludes any structural break, while the alternative hypothesis is that the series is a trend stationary process with a one-time break occurring at an unknown point in time. Table 2 summarizes the result of the Zivot and Andrews (1992) test in the presence of a structural break, allowing for a change in both the intercept and trend. In this model, the break point is endogenously determined by running the model sequentially allowing for this break point to be any day within a 15 percent trimming region. The optimal lag length is determined on the basis of the Bayesian Information Criterion. Using the Zivot and Andrews (1992) procedure, the time of the structural changes (impacting on both the intercept and the slope of each series) for each of the variables is detected and the results are presented in Table 2. As shown, the most significant structural breaks occur around 1973 (for the annual version of the data) and around the third quarter of 1973 (for the quarterly version of the data). This date corresponds broadly to the pronounced structural changes associated with the 1973 global oil crisis. Nevertheless, the previous unit root testing findings remain robust.

Table 2. Zivot-Andrews unit root tests with break in the intercept and trend

Variable	k	t_a	Break
Annual			
ry	3	-3.12	1973
Δry	2	-6.08	
npo	2	-3.25	1973
Δnpo	1	-5.91	
rpo	3	-4.11	1973
Δrpo	1	-7.32	
npo^+	3	-3.38	1974
Δnpo^+	2	-6.61	
npo^-	3	-3.35	1973
Δnpo^-	1	-6.24	
rpo^+	4	-4.13	1973
Δrpo^+	2	-7.24	
rpo^-	3	-3.65	1973
Δrpo^-	2	-6.39	
Quarterly			
ry	5	-3.37	1973Q4
Δry	3	-6.84	
npo	6	-3.62	1973Q3

Δnpo	5	-6.48	
rpo	5	-3.83	1973Q3
Δrpo	4	-6.57	
npo^{+}	4	-3.61	1974Q1
Δnpo^{+}	3	-6.82	
npo^{-}	5	-3.62	1973Q3
Δnpo^{-}	3	-6.73	
rpo^{+}	6	-4.34	1973:Q4
Δrpo^{+}	5	-8.51	
rpo^{-}	5	-3.78	1973Q3
Δrpo^{-}	4	-6.82	

Notes: t_a is the estimated t-statistic related to the null hypothesis of the presence of a unit root under a break and k is the number of lags in the test. Critical values at 1, 5 and 10 percent levels are -5.57, -5.08 and -4.82, respectively.

Panel cointegration

We test for the presence of a long-run relationship using the Pedroni (1999, 2001) panel cointegration tests. The panel cointegration test results are presented in Table 3 with the lag length chosen on the basis of the Akaike information criterion (AIC) with individual intercepts and trends. The test results strongly reject the null of no cointegration in favor of the presence of a long-run relationship.

Table 3. Panel cointegration tests

Test	Annual	Quarterly
ry-npo		
Panel v-statistic	30.525*	37.919*
Panel rho-statistic	-34.541*	-36.008*
Panel PP-statistic	-35.524*	-36.236*
Panel ADF-statistic	-5.167*	-7.2387*
Group rho-statistic	-33.355*	-36.635*

Group PP-statistic	-32.119*	-36.872*
Group ADF-statistic	-7.058*	-7.237*
ry-rpo		
Panel v-statistic	32.248*	34.242*
Panel rho-statistic	-33.894*	-34.327*
Panel PP-statistic	-33.238*	-34.609*
Panel ADF-statistic	-4.983*	-6.246*
Group rho-statistic	-32.609*	-36.352*
Group PP-statistic	-32.442*	-36.854*
Group ADF-statistic	-5.265*	-6.568*
ry-npo^{+}-npo^{-}		
Panel v-statistic	37.812*	38.491*
Panel rho-statistic	-36.355*	-37.241*
Panel PP-statistic	-36.625*	-37.409*
Panel ADF-statistic	-6.298*	-6.815*
Group rho-statistic	-35.328*	-36.096*
Group PP-statistic	-35.809*	-36.215*
Group ADF-statistic	-6.133*	-6.656*
ry-rpo^{+}-rpo^{-}		
Panel v-statistic	36.562*	38.476*
Panel rho-statistic	-34.358*	-37.236*
Panel PP-statistic	-34.247*	-37.805*
Panel ADF-statistic	-5.8409*	-7.488*
Group rho-statistic	-34.135*	-36.562*
Group PP-statistic	-33.214*	-36.657*
Group ADF-statistic	-5.246*	-7.894*

* indicates 1% rejection level (The rejection of the null hypothesis of no-cointegration).

Having established cointegration, we estimate the long-run model by using the FMOLS methodology. The asymmetric long-run estimations are reported in Table 4. Model 1, Model 2, Model 3 and Model 4, respectively corresponds to equations 1 to 4. In all the four models and based on both annual and quarterly data, the coefficient of oil price is significant at 1% level. As expected the coefficients of both real and nominal oil prices in the real personal income equation is negative as shown in Models 1 and 2. The coefficient of the cumulative sums of positive oil prices in Models 3 and 4 is negative, while that of the negative oil price is positive. This implies that an increase in oil prices reduces output in the U.S. States, while lower oil prices increase output, in line with the theoretical exposition. Further, we can infer asymmetric effects by comparing the size of the coefficients of the cumulative sums of the positive and negative oil prices in Models 2 and 4. The impact of negative oil prices is slightly larger than that of positive oil prices. This appears to contradict the conventional thinking.

Table 4. Long-run FMOLS estimates

Coefficients	Model 1	Model 2	Model 3	Model 4
Annual				
b_1	-0.696*			
	(-4.57)			
b_2		-0.806*		
		(-6.88)		
b_3			-0.439*	
			(-6.15)	
b_4			0.474*	
			(6.25)	
b_5				-0.322*
				(-4.48)
b_6				0.250*

				(5.73)
$\bar{R}^2$	0.46	0.50	0.55	0.58
Quarterly				
b_1	-0.175*			
	(-5.16)			
b_2		-0.494*		
		(-6.37)		
b_3			-0.272*	
			(-6.48)	
b_4			0.342*	
			(5.61)	
b_5				-0.252*
				(-5.39)
b_6				0.371*
				(5.61)
$\bar{R}^2$	0.49	0.53	0.56	0.56

Notes: Figures in parentheses denote t-statistics and * indicates significance at 1%.

Causality results

Having found the presence of a long-run equilibrium, we are also interested in examining the direction of causality between the variables. We perform Wald *F*-tests on the significance of the coefficients, evaluating two different causality relationships: a short-run causality, testing the significance of the coefficients related to the lagged factors and a long-run causality related to the coefficient for the error correction term (EC). Table 5 reports the causality results. The short-run causality analysis provides evidence of bidirectional causality between aggregate oil prices and output. However, there is evidence of unidirectional causality running from both positive and negative oil prices to output based on the annual data. The quarterly data produced slightly different result with respect to positive and negative oil prices as it documents evidence of bidirectional causality. As expected, the positive and negative oil prices do not Granger cause each other. Moreover, we find evidence of long-run bidirectional causality between aggregate oil prices and output, but the presence of unidirectional causality running from positive and negative oil prices to output. For the case of quarterly data, there is supportive evidence for bidirectional long-run causality between positive and negative oil prices and output.

Table 5. Causality test results

		Sources of Causation		
		Short-Run		**Long-Run**
Annual				
	Δry	Δnpo		EC
Δry	---	13.27		-0.106
		[0.00]		[0.00]
Δnpo	13.72	---		-0.165
	[0.00]			[0.00]
	Δry	Δrpo		EC
Δry	---	22.01		-0.137
		[0.00]		[0.00]
Δrpo	24.32	---		-0.172
	[0.00]			[0.00]
	Δry	Δnpo^{+}	Δnpo^{-}	EC
Δry	---	26.21	29.01	-0.128
		[0.00]	[0.00]	[0.00]
Δnpo^{+}	0.53	---	0.63	-0.016
	[0.75]		[0.70]	[0.83]

Δnpo^-	1.20	0.26	---	-0.031
	[0.30]	[0.65]		[0.49]
	Δry	Δrpo^+	Δrpo^-	EC
Δry	---	27.68	25.34	-0.217
		[0.00]	[0.00]	[0.00]
Δrpo^+	0.76	---	0.34	-0.009
	[0.56]		[0.79]	[0.89]
Δrpo^-	0.54	0.18	---	-0.012
	[0.76]	[0.77]		[0.62]
Quarterly				
	Δry	Δnpo		EC
Δry	---	22.46		-0.185
		[0.00]		[0.00]
Δnpo	18.40	---		-0.179
	[0.00]			[0.00]
	Δry	Δrpo		EC
Δry	---	26.45		-0.197
		[0.00]		[0.00]
Δrpo	24.37	---		-0.214
	[0.00]			[0.00]
	Δry	Δnpo^+	Δnpo^-	EC
Δry	---	26.90	24.57	-0.175
		[0.00]	[0.00]	[0.00]
Δnpo^+	7.95	---	0.49	-0.074
	[0.00]		[0.78]	[0.10]
Δnpo^-	11.48	0.22	---	-0.053
	[0.00]	[0.69]		[0.09]
	Δry	Δrpo^+	Δrpo^-	EC

Δry	---	23.48	25.83	-0.236
		[0.00]	[0.00]	[0.00]
Δrpo^{+}	23.48	---	0.17	-0.096
	[0.00]		[0.88]	[0.06]
Δrpo^{-}	20.11	0.27	---	-0.084
	[0.00]	[0.71]		[0.05]

Notes: Figures in brackets denote p-values.

Another important issue that can adversely affect the above findings, especially those related to the panel unit root tests, is the presence of cross sectional dependence. Under the null hypothesis of no cross-sectional dependence, the test statistic is asymptotically distributed as standard normal. Pesaran's (2004) approach has remarkable positive qualities in samples of practically all relevant sizes and remains robust in a variety of settings (Pesaran, 2004). Table 6 reports the Pesaran's (2004) CD test for cross-sectional dependence. This test is based on the residuals from the cross correlation of the ADF(p) regressions. The results indicate that the null hypothesis of independence is strongly rejected across all panels, revealing that the panel methodological approach is more appropriate than a state-by-state estimation. Table 6 also reports the results from the slope homogeneity test of Pesaran and Yamagata (2008). The test rejects the null of the slope homogeneity hypothesis, supporting the state-specific heterogeneity.

Table 6. CD cross-section dependence and heterogeneity tests

Panels	p-value (CD test)	p-value (Heterogeneity test)
Annual		
ry-npo	[0.00]	[0.00]
ry-rpo	[0.00]	[0.00]
ry-npo^{+}-npo^{-}	[0.00]	[0.00]
ry-rpo^{+}-rpo^{-}	[0.00]	[0.00]
Quarterly		
ry-npo	[0.00]	[0.00]

ry-rpo	[0.00]	[0.00]
ry-npo^{+}-npo^{-}	[0.00]	[0.00]
ry-rpo^{+}-rpo^{-}	[0.00]	[0.00]

As earlier shown, we observe significant structural breaks in 1973. Therefore, we also perform sub-sample analysis over the pre- and post-1973 period, to check out whether the estimates vary between sub-samples. The results are presented in the Appendix. The null hypothesis of no cointegration is rejected in the two sub-samples for both the annual and the quarterly version of the data. The long-run estimates from the FMOLS are statistically insignificant in both the 1929-1972 and 1948-1972 sub-samples for annual and quarterly data, respectively. However, the estimates turn out to be statistically significant at the 1% level in the 1973-2013 sub-sample period.

Conclusions

This study examined the dynamic relationship between oil prices and growth across the 48 U.S. States, using a panel data framework. Both annual and quarterly data spanning the periods 1973 to 2013 and 1948 to 2013, respectively, were used. To allow for asymmetry in both the cointegration and causality testing, we decomposed oil prices into cumulative sums of positive and negative oil prices, consistent with Hatemi-J (2012). Overall, we estimated four alternative empirical panel models. The suitability of the panel methodology over state-by-state equation was documented by using the Pesaran's (2004) CD test for cross-sectional dependence and the slope homogeneity test by Pesaran and Yamagata (2008). The results provided evidence in favour of a long-run relationship between the different components of oil prices and real output. Long-run estimates in the case of aggregate, positive and negative oil prices were shown to be statistically significant. Furthermore, the study tested for the presence of both short- and long-run causality. We found evidence of bidirectional causality between aggregate oil prices and output in both the short- and long-run, irrespective of the data frequency used. However, there was ample evidence of unidirectional causality running from both positive and negative oil prices to output, based on annual data, while the case of quarterly data documented evidence of bidirectional causality in both the long- and short-run. We also implemented a sub-sample analysis for the pre- and post-1973 period, given that a preliminary check of unit root tests with breaks in both the intercept and trend indicated the presence of a significant break occurring in 1973. The panel unit root tests along with the long-run estimates were similar to those concerning the full-sample.

The policy implications of the empirical findings point out that given that oil prices and output growth exhibit a complementarity relationship to each other, higher oil prices would act as an inflation tax on consumers, producers and investors across the U.S. States, thus, reducing the income available for spending on other goods. In contrast, if the economy is booming, consumers may have more to spend and this would possibly lead to higher demand for vehicles and other energy-intensive products, consequently, pushing oil prices up. In that case, both monetary and fiscal policies are needed to counterbalance inflationary pressures and a weak aggregate demand. The optimal response of monetary policy to movements in oil prices will be to stabilize inflation, while fiscal policy will be required to contain aggregate demand. However, caution should be exercised in the use of monetary policy, since the inflation stabilization objective could lead to increases in the federal funds rate, causing further declines in output, hence, exacerbating recessionary conditions. Furthermore, output declines, arising from oil shocks, may be mitigated by keeping the interest rate constant, though this is expected to have a short-run effect. Over the long-run, however, once the public's expectations relative to monetary policy implemented are formed, this mitigation effect is expected to dissipate.

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Appendices

Sub-samples Analysis

Table A1. Panel cointegration tests across sub-samples

	Annual		Quarterly	
	1929-1972	1973-2013	1948-1972	1973-2013
ry-npo				
Panel v-statistic	24.127*	39.329*	26.239*	41.384*
Panel rho-statistic	-23.329*	-36.139*	-24.109*	-38.288*
Panel PP-statistic	-23.138*	-36.238*	-24.236*	-38.189*
Panel ADF-statistic	-4.783*	-7.298*	-4.658*	-8.091*
Group rho-statistic	-23.458*	-36.138*	-24.222*	-40.256*
Group PP-statistic	-22.093*	-35.014*	-23.427*	-40.133*
Group ADF-statistic	-4.154*	-7.244*	-4.287*	-8.219*
ry-rpo				
Panel v-statistic	22.381*	38.330*	23.647*	44.215*
Panel rho-statistic	-20.167*	-35.315*	-22.753*	-42.387*
Panel PP-statistic	-20.541*	-35.352*	-22.312*	-42.121*
Panel ADF-statistic	-4.086*	-7.984*	-4.339*	-7.342*
Group rho-statistic	-22.149*	-36.083*	-22.309*	-42.926*
Group PP-statistic	-22.683*	-36.476*	-22.432*	-42.009*
Group ADF-statistic	-4.151*	-7.3177*	-4.109*	-7.925*
ry-npo^{+}-npo^{-}				
Panel v-statistic	26.872*	37.573*	24.573*	44.215*
Panel rho-statistic	-25.139*	-35.267*	-22.327*	-43.873*
Panel PP-statistic	-25.309*	-35.462*	-22.452*	-43.199*
Panel ADF-statistic	-5.218*	-6.806*	-4.533*	-7.093*
Group rho-statistic	-24.109*	-34.244*	-24.031*	-44.167*

Group PP-statistic	-24.244*	-34.429*	-22.167*	-44.343*
Group ADF-statistic	-5.450*	-6.548*	-4.131*	-7.757*
ry-rpo^{+}-rpo^{-}				
Panel v-statistic	24.236*	35.265*	25.130*	45.034*
Panel rho-statistic	-22.259*	-34.327*	-23.208*	-43.145*
Panel PP-statistic	-22.190*	-34.684*	-23.765*	-43.084*
Panel ADF-statistic	-5.133*	-6.908*	-4.543*	-7.686*
Group rho-statistic	-23.764*	-34.133*	-24.116*	-44.490*
Group PP-statistic	-22.746*	-34.276*	-24.348*	-64.339*
Group ADF-statistic	-5.491*	-6.644*	-4.327*	-7.927*

* indicates 1% rejection level (The rejection of the null hypothesis of no-cointegration).

Table A2. Long-run FMOLS estimates using annual data

Coefficients	Model 1	Model 2	Model 3	Model 4
1929-1972				
b_1	-0.042			
	(-1.08)			
b_2		-0.036		
		(-1.14)		
b_3			-0.041	
			(-1.17)	
b_4			0.035	
			(0.51)	
b_5				-0.021
				(-0.73)
b_6				0.034
				(1.15)

$\bar{R}^2$	0.09	0.11	0.13	0.10
1973-2013				
b_1	-0.514*			
	(-4.82)			
b_2		-0.648*		
		(-6.26)		
b_3			-0.329*	
			(-5.49)	
b_4			0.380*	
			(5.61)	
b_5				-0.268*
				(-4.71)
b_6				0.219*
				(5.71)
$\bar{R}^2$	0.41	0.43	0.46	0.42

Notes: Figures in parentheses denote t-statistics and * indicates significance at 1%

Table A3. Long-run FMOLS estimates using quarterly data

Coefficients	Model 1	Model 2	Model 3	Model 4
1948-1972				
b_1	-0.055			
	(-0.94)			
b_2		-0.068		

		(-1.29)		
b_3			0.028	
			(0.72)	
b_4			0.029	
			(0.44)	
b_5				0.035
				(1.24)
b_6				0.049
				(1.30)
$\bar{R}^2$	0.06	0.08	0.10	0.12
1973-2013				
b_1	-0.579*			
	(-5.46)			
b_2		-0.594*		
		(-6.73)		
b_3			-0.375*	
			(-5.90)	
b_4			0.394*	
			(5.82)	
b_5				-0.291*
				(-5.38)
b_6				0.249*
				(5.81)
$\bar{R}^2$	0.44	0.48	0.47	0.51

Notes: Figures in parentheses denote t-statistics and * indicates significance at 1%