How Persistent are Shocks to Energy Prices?

*Atanu Ghoshray**

ABSTRACT

Whether shocks to energy prices are permanent or transitory remains a contentious issue. This may result from mis-specification of the econometric tests, due for example to the uncertainty over the presence of a trend, or the possible presence of structural breaks and non-stationary volatility in the data. This paper makes a contribution by addressing the underlying characteristics of energy price data that influence such econometric tests. First, we detect whether the data are characterised by non-stationary volatility and possible trend breaks. The next step involves employing novel unit root tests that unify the underlying characteristics, such as trend break and/or nonstationary volatility, of the data. We conclude shocks to energy prices are not transitory. We further decompose a benchmark oil price and its demand and supply components into their permanent and transitory components and compute the cross correlations to find that they conform to standard theories of commodity storage models.

Keywords: Energy prices; volatility; persistence; structural breaks.

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1. INTRODUCTION

Prior to the early 1970s, crude oil prices were largely set by the major oil companies, and after the 1970s oil price shock, the Organisation of Petroleum Exporting Countries (OPEC) emerged as a price setter. However, since the mid-1980s the influence of OPEC has waned and market forces have played an increasingly important role in the determination of crude oil prices (Fattouh 2007). It has been argued that oil prices have become increasingly volatile (Charles and Darné 2014), exceeding the levels of volatility of other commodities, and are more susceptible to market disturbances. Dvir and Rogoff (2009) state that since 1973, oil supplies became restricted and prices as a result, have become volatile and persistent. The volatility of oil prices has been changing over time. For example, Regnier (2007) notes that during the late 1980s oil prices were relatively higher than the late 1990s, however, during the relatively lower price period, prices were more volatile than during the period of relatively higher oil prices. This underlying characteristic of high volatility in oil prices could be reflected in the prices of refined products of crude oil, such as the price of gasoline and heating oil, as well as the price of seemingly closely correlated energy prices such as coal and natural gas.

This paper examines whether exogenous shocks have a transitory or permanent effect on energy prices. Knowledge of how persistent the shocks are to energy prices is important for policy makers as energy commodities, especially oil, can have a large impact on the macroeconomy (Hamilton 2009). This is also important for countries that are heavily export dependent on energy commodities such as natural gas, oil and coal, to consider the efficacy of counter-cyclical stabilization policies. Further, permanent oil price shocks can have consequences in terms of creating inflationary

* Newcastle University Business School, Newcastle University, 5 Barrack Road, Newcastle-upon-Tyne, NE1 4SE, UK. E-mail: Atanu.Ghoshray@newcastle.ac.uk

pressures. A long-lived increase in energy prices can have a direct effect of reducing household income and also an indirect effect on lowering consumer confidence. Not surprisingly, there have been several studies that examine whether energy prices are persistent to shocks (see e.g., Barros and Gil-Alana (2016), Sun and Shi (2015), Narayan and Liu (2015), Noguera (2013) and references within). However, the fact that energy prices are likely to exhibit concurrent volatility (see Dvir and Rogoff 2009), can severely affect the methods used in recent studies to test for persistence, leading to erroneous conclusions that can have far reaching consequences. This paper carries out novel and robust tests that can examine persistence of such shocks allowing for changing volatility, paying particular attention to the presence of trends and possible structural breaks.

The most common method to study the dynamics of energy prices are unit root tests. Unit root tests can conclude whether the price series is a trend or a difference stationary process. If the price series is a trend-stationary process, or in other words prices are integrated of order zero, then shocks to prices would have transitory effects. Otherwise, if the price series is a difference-stationary process, or in other words, there is a unit root in the series, or prices are integrated of order one, then shocks to prices would have permanent effects. If prices were forecast to be stationary that would be consistent with a competitive market which ensures that prices mean revert to marginal cost (Pindyck 1999). Recent developments in the field of unit root tests allow for structural breaks (e.g., Zivot and Andrews (1992), Lumsdaine and Papell (1998), Lee and Strazicich (2003) among others), as it is well known that the standard unit root tests suffer from low power in the presence of structural breaks in the data. However, unit root tests that allow for structural breaks suffer from the problem by assuming breaks exist in the data when the true data generating process does not contain any breaks at all. At an intuitive level, it seems more natural to be first able to ascertain if breaks are at all present before proceeding to conduct unit root tests allowing for such breaks. In the absence of breaks, these tests may suffer from low power and size distortions due to the inclusion of extraneous break dummies thereby potentially leading the researcher to estimate a differenced specification when a level specification is in fact more appropriate (see Ghoshray et. al. 2014). We adopt an intuitive approach to first determine if structural breaks exist at all before conducting unit root tests that allow for such breaks, by applying the method due to Perron and Yabu (2009) and Harvey et. al. (2009a) that allows for consistent estimation of a structural break in the energy price series while being agnostic about the order of integration of data. If we find evidence of a structural break, then using the sequential procedure of Kejriwal and Perron (2010) we test for multiple breaks.

As highlighted by Cavaliere et. al. (2011), another problem with unit root tests is the possible presence of time varying unconditional volatility in the data. This has been flagged up in the literature (see for example, Busetti and Taylor 2003, Sensier and van Dijk 2004, Cavaliere and Taylor 2008) as a considerable drawback for a data series that displays simultaneous breaks in the trend and unconditional volatility. Given the recent developments, the literature is far from a consensus on the unit root properties of energy prices, which may result from low power or size properties of unit root tests, due to the presence of structural breaks and possible non-stationary volatility which are believed to plague energy prices. We adopt a novel unit root test due to Cavaliere et. .al (2011), CHLT hereafter, that allows for a possible trend break and nonstationary volatility. The nonstationary volatility could take the form of a single or multiple abrupt variance breaks, smooth transition variance breaks or a trending variance. The simulation results shown by CHLT conclude a large impact on both power and size properties of the unit root tests for just a single break in volatility.

Campbell and Perron (1991) have stressed that proper specification of the deterministic components is essential in obtaining unit root tests with reliable finite sample properties. The upshot is that the unit root question in energy prices cannot be properly analysed until some characterisation of the underlying deterministic component is made. The uncertainty of whether or not to include a constant, or a constant and linear trend in a unit root test regression, is a problem that plagues unit root tests. The inclusion of an underlying time trend has been emphasised in recent studies of energy prices. Narayan and Liu (2015), and Westerlund et. al. (2015) note that almost all energy variables can be characterised to contain a linear time-trend and have argued for the need to include the trend when testing for a unit root. However, others may argue that volatility in energy prices may overshadow the need of accounting for a trend. Given the possibility that there may be no trend break and also that there may be no significant trend in the data, we apply the novel test due to Smeekes and Taylor (2012) which is a union unit root test that is robust to nonstationary volatility, trend uncertainty, and uncertainty about the initial condition[1].

Given that structural breaks and nonstationary volatility are highly plausible features of energy prices, we consider novel procedures due to CHLT and Smeekes and Taylor (2012) to determine whether shocks to energy prices are stationary, which is the primary contribution of this paper. We take a comprehensive approach, by choosing two sets of data measured at monthly and weekly frequencies, which cover the benchmark crude oils, being West Texas Intermediate (WTI), Brent and Dubai/Oman[2], as well as other energy commodities such as natural gas, coal, heating oil and gasoline prices. Our finding that shocks to energy prices are not transitory in nature allows us to extend the analysis by decomposing a benchmark crude oil price into its permanent and transitory components; and in the same manner decompose a demand side variable as well as a supply side variable to examine the cross correlations between the permanent and transitory components. The aim is to capture linkages that can explain models of commodity prices with storage that help validate our findings. This paper is structured as follows: the following section provides a brief but critical review of the literature; this is followed by a section which describes the novel econometric methods employed in this study; the next section describes the data and the empirical results; and finally the last section concludes.

2. LITERATURE REVIEW

In this section we briefly review the plethora of studies that test for unit roots in energy prices. We begin by those that employ unit root tests without structural breaks, followed by unit root tests with structural breaks. We finally review those studies that employ robust tests for structural breaks as a pre-test for unit roots. Silvapulle and Moosa (1999), Alizadeh and Nomikos (2002), Serletis and Rangel-Ruiz (2004) and Coimbra and Esteves (2004) are among the prominent studies that have applied conventional unit root tests without structural breaks such as the ADF (Dickey and Fuller, 1979), the PP (Phillips and Perron 1988) and the KPSS (Kwiatkowski et al., 1992) tests to benchmark crudes and the overall conclusion is that the null hypothesis of a unit root cannot be rejected. While most studies have employed data that are measured at monthly or weekly frequencies, there are some studies that have been applied to long spans of data measured annually. Pindyck (1999) employing annual data from 1870 to 1996, and Krichene (2002) examining data from 1918 to 1999, could not reject the null hypothesis of a unit root for energy prices using standard ADF tests. The broad conclusion is that using standard unit root tests, the shocks are not transitory in nature. However, it is well known that a problem with 'no break' unit root tests is their low power in the presence of a structural break in the data (Perron 1989).

Several prominent studies have applied unit root tests in the presence of structural breaks. Gulen (1997) and Gulen (1999) conduct unit root tests with a single exogenously chosen structural

1. By the initial condition we mean the deviation of the initial observation of price from its modelled deterministic trend.
2. This benchmark began to incorporate Oman following a significant decline in the production of Dubai since the 1990s.

break for selected crudes. The results broadly conclude the presence of a unit root in the data. However, it is now well known, that the choice of an exogenous structural break in unit root tests has come under severe criticism (see Christiano 1992) as this invalidates the distribution theory underlying conventional testing. Favouring the endogenously determined break instead, Serletis (1992) and Sadorsky (1999) employ endogenously determined structural break unit root tests due to Zivot and Andrews (1992) to selected energy prices. Both the studies reject the unit root null in favour of trend stationarity. A problem however, with these studies is that the breaks only appear in the alternative hypothesis and this leads to serious power and size distortions in unit root tests (see Ghoshray et. al. 2014). Addressing this asymmetric treatment of breaks, Postali and Picchetti (2006) employ the Lee and Strazicich (2004) Lagrange Multiplier (LM) type unit root test on crude oil prices and reject the null of a unit root. Using the same LM test, Lee et. al. (2006) reject the null of a unit root for non-renewable natural resource prices. However, Maslyuk and Smyth (2008), employ the same test on a different time period for higher frequency data and are unable to reject the presence of a unit root. Ghoshray and Johnson (2010) show that using the LM type test with breaks, all the energy prices except coal are found to be stationary. However, a drawback of all these studies is that in the absence of breaks, such LM type tests are likely to suffer from low power due to the inclusion of extraneous dummy variables.

More recently, addressing this drawback, robust tests have been developed that allow to determine whether structural breaks are at all present in the data. Following this line of reasoning, Noguera (2013) employs a robust method due to Kejriwal and Perron (2010) for estimating the structural breaks, followed by a unit root test taking into account the number of identified breaks. The results show that shocks to oil prices are transitory in nature. While the data set chosen by Noguera (2013) spans a long period of time, allowing for up to twelve structural breaks is a problem, as the unit root process is observationally equivalent to a stationary process with multiple breaks in the limit. Sun and Shi (2015) employ the same robust test to higher frequency data over a smaller sample period to detect structural breaks followed by the unit root test due to Carrion-i-Silvestre et. al. (2009). Their finding is that the null hypothesis of a unit root is rejected. To summarize the studies that use unit roots with structural breaks, the evidence is on balance, in favour of energy price shocks being transitory in nature. Numerous studies have concluded that oil prices are characterised by high volatility (see Narayan and Narayan (2007) and references within). In a recent study, Narayan and Liu (2015) employ a trend-GARCH unit root tests which is an extension of the Narayan and Liu (2011) method and find that for most energy prices the unit root can be rejected.

In general, most recent studies that have employed a unit root test allowing for structural breaks broadly conclude that energy prices are stationary in nature. However, it is clear that several features that are likely to exist in the data, that can influence these unit root tests, have not been addressed. This paper therefore makes an important contribution by dealing with the features of energy prices; the uncertainty of a trend and the initial condition (or the deviation of the initial observation of price from its modelled deterministic component), the possibility of a structural break in the trend, as well as nonstationary volatility which could take the form of a single or multiple abrupt variance breaks, smooth transition variance breaks or a trending variance.

3. ECONOMETRIC METHODS

The general model that we employ to identify structural breaks in the trend of energy prices is given by the equation below:

$$P_t = \mu_0 + \beta_0 t + \sum_{i=1}^{K} \mu_i DU_{it} + \sum_{i=1}^{K} \beta_i DT_{it} + u_t \tag{1}$$

where P_t is the energy price, $DU_{it} = I(t > T_i)$, $DT_{it} = (t - T_i)I(t > T_i)$, for $i = 1,2,...,K$ are the intercept and slope dummy variables. A break in the trend occurs at time $T_i = [T\lambda_i]$ when $\beta_i \neq 0$ and λ_i is the break fraction. The date of the breaks, T_i and the number of breaks K is assumed to be unknown. The error term u_t can be either I(0) or I(1). A test for a structural break is made using the novel procedure of Perron and Yabu (2009) that is robust to whether the underlying data is I(0) or I(1) based on a Feasible Quasi Generalized Least Squares method. They recommend the following exponential Wald ($ExpW$) statistic to determine the structural break:

$$ExpW = log\left[T^{-1}\sum_{\lambda_1 \in \Lambda_1} exp\left(\frac{1}{2}W_{QF}\left(\lambda_1\right)\right)\right] \tag{2}$$

A further robust test for structural break(s) in the trend is carried out based on the procedure by Harvey et. al. (2009a). The method involves the use of weighted statistics which are shown to have standard normal limiting null distributions and attain the Gaussian asymptotic local power envelope, in each case regardless of whether the shocks are I(0) or I(1) Details of the test are found in Harvey et. al. (2009a).

In order to test the null hypothesis of constant unconditional variance, we make use of the well-known procedure due to Inclan and Tiao (1994). However, the Inclan and Tiao (1994) method has two drawbacks being: (a) it does not allow for conditional heteroscedasticity and (b) it ignores the fourth moments. Sanso et. al. (2004) takes into account these features and produces a correctly sized test at the expense of a slight loss of power in comparison to the Inclan and Tiao (1994) method. Sanso et. al. (2004) recommend to use both the tests to check for nonstationary volatility which we carry out in this study.

To test for unit roots in the presence of a possible break in trend and nonstationary volatility we make use of the procedure developed by CHLT. To implement the procedure we adopt the following model:

$$P_t = \alpha + \beta t + \gamma DT(\lambda) + \varepsilon_t;\ \ t = 1,2,.....,T. \tag{3}$$

$$\varepsilon_t = \rho_T \varepsilon_{t-1} + \eta_t;\ \ t = 2,3,....,T. \tag{4}$$

where, $\eta_t = C(L)e_t$, and $e_t = \sigma_t z_t$. Equation (3) allows for the possibility of a structural break in the trend; where γ denotes the break magnitude, λ denotes the unknown break fraction, so that the break point is given by λT. Using the above model given by (3) and (4), CHLT compute a set of M-statistics to test for the presence of a unit root in the presence of a possible break in trend and nonstationary volatility.

It is possible that there may be no break in the trend in which case a further test proposed by Smeekes and Taylor (2012) is employed. This is a bootstrap union test for unit roots in the presence of non-stationary volatility which builds on the test proposed by Cavaliere and Taylor (2008) who show that standard ADF tests are asymptotically not correctly sized for such time series data that is characterised by non-stationary volatility. Besides, both the presence or absence of a linear trend in the data series can lead to complications with unit root testing. For example, if the data does contain a linear trend but this is not included in the unit root test, the power of the test will decrease. Alternatively, if the data does not contain a trend but one is included in the unit root test, then there is also a loss of power (Marsh 2007). Further, if the initial condition is small, Muller and Elliott (2003) show that the Dickey Fuller (DF) test with ordinary least squares detrending, denoted DF – OLS, suffers from low power relative to the DF test with quasi-differenced (QD) or generalized least squares (*GLS*) detrending, denoted as $DF - QD$. Alternatively, if the initial condition is large, then the opposite happens. The upshot is that both the deterministic trend and the initial condition is not observed, leading to uncertainty with regards to which test to apply. To deal with these issues, Harvey, et. al. (2009b) construct a new test formed as a union of rejections of unit root tests with and without a deterministic linear trend and show that this union test can maintain high power and size irrespective of the true value of the trend. Harvey et. al. (2012) extend the analysis of Harvey et. al. (2009b) by considering the impact of both trend and initial condition uncertanty simultaneously. They propose a four-way union of rejections of $DF - QD$ and $DF - OLS$ tests, both with and without trend. The modified union test statistic, is given by:

$$UR_4 = min\left[DF - QD^{\mu}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{QD}^{\tau*}(\pi)}\right) DF - QD^{\tau}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{OLS}^{\mu*}(\pi)}\right) DF - OLS^{\mu}, \left(\frac{cv_{QD}^{\mu*}(\pi)}{cv_{OLS}^{\tau*}(\pi)}\right) DF - OLS^{\tau} \right]$$

where $cv_j^{\delta}(\pi)$ denotes the asymptotic critical value of the Dickey Fuller test which could contain an intercept (μ), or intercept and trend (τ), at nominal level π. This test thereby deals with the uncertainty about the trend and the initial condition.

However, Smeekes and Taylor (2012) note that the uncertainty about the presence of a trend and the initial condition needs to be dealt with the possible presence of nonstationary volatility. To this end, Smeekes and Taylor (2012) consider union tests that are robust to nonstationary volatility, trend uncertainty, and uncertainty about the initial condition. Their test is based on the wild bootstrap approach, combined with the sieve principle to account for stationary serial correlation, designed to be robust over uncertainty about the presence of a deterministic trend and uncertainty about the initial condition. This test is an improvement over the union tests of Harvey et. al. (2012) which are found to be incorrectly sized in the presence of nonstationary volatility. Smeekes and Taylor (2012) consider two bootstrap union tests, unit root A type test (UR_{4A}) and unit root B type test (UR_{4B}); the UR_{4A} test does not include a deterministic trend in the test, while the UR_{4B} test does include a trend in the bootstrap data generating process. These tests would appear to constitute a valuable option if one needs to deal simultaneously with uncertainty regarding the trend and the initial condition and to provide results that are simultaneously robust to the possible presence of nonstationary volatility.

4. DATA AND EMPIRICAL RESULTS

Ambiguous results are likely when the data considered is of higher frequency as it is likely to exhibit fat tails and volatility clustering (Boswijk and Klaassen, 2005). Further, one could argue that in order to deal with the low power of the unit root test, researchers increase the power by obtaining more observations. This can be possible by either considering a longer time span, or employing data recorded at a higher frequency. However, Campbell and Perron (1991) have argued that increasing the frequency does not necessarily lead to higher power. They state that the power of the unit root test depends on the span of the data chosen rather than the frequency of the observations. The value of long sample periods of 50–140 years, as employed by studies considering annual data may be questioned, given the typical time horizon in oil and gas field investment of three decades (Maslyuk and Smyth, 2008). Conversely, short sample periods of high frequency data are unlikely to capture the long lags in energy investment cycles. Keeping in perspective these considerations, we have relied on monthly data and weekly data that covers the period of interest which we believe to be sufficient to track the nature of shocks in benchmark crudes and other energy variables.

The data used in this analysis are drawn from two sources using two different frequencies. We choose the three benchmark crude oils being West Texas Intermediate (WTI), Brent and Dubai/ Oman along with two other energy prices being natural gas and coal obtained from the *International Financial Statistics*. These are free market price indices with monthly frequency spanning from January 1986 to March 2016. The second set of data include two benchmark crudes, WTI and Brent along with Gasoline and heating oil prices from the New York Harbour which were obtained from the *U.S. Energy Information Administration* (EIA). These are spot prices measured at a weekly frequency. In this study, the starting point of the data is January 1986 for monthly data and slightly different starting time periods for the weekly prices depending on availability. Specifically, WTI starts from January 1986, Brent from May 1987, and Heating oil along with gasoline prices start from June 1986. All data periods for weekly prices end at May 2016. A plot of the data sets are shown in Figures 1 and 2. The monthly data is shown in Panels A – E in Figure 1 and the weekly data in Panels A – D in Figure 2.

Eyeballing the data it would not seem unusual to expect the possibility of a break in trend and signs of changing volatility. The time span is chosen so that we can include most energy prices.

We first determine whether the price series examined in this study show any evidence of nonstationary volatility. To determine whether there is any evidence of a break in variance we employ the test due to Inclan and Tiao (1994) and also conduct the Sanso et. al. (2004) test that takes into account heteroscedasticity and the fourth moments of the data. Both tests show evidence of nonstationary volatility with several breaks in variance. Interestingly we find the preponderance of breaks in variance to be centred around 1999 and 2007 for monthly prices of energy commodities; while for weekly data the breaks in variance are more frequent and tend to be more concentrated over the last 15 years in comparison to the 1980s and 1990s. The full set of results are shown in Table 1. While the number and locations of the break in variance differ for the two different methods, the evidence that breaks in variance are present in energy price data cannot be rejected.

For detecting structural breaks we employ the Perron and Yabu (2009) and Harvey et. al. (2009a) tests. The advantage of employing this procedure is that we can be agnostic about the underlying order of integration of the price series. The results are shown in Table 2. The null hypothesis is that of no structural break. For monthly data we only find evidence of a slope break in natural gas. The estimated statistic for the Harvey et. al. (2009a) test (i.e. 3.11) is greater than the critical value, leading us to reject the null hypothesis of no break in favour of the alternative hypothesis of a structural break. For weekly data the results show a break in trend for WTI and heating oil. This is determined by the Perron and Yabu (2009) test where the estimated statistic is greater than the critical concluding the presence of a structural break.[3] The break date in 1998 for natural gas could be linked to the Asian financial crisis, the break date in 1994 for weekly WTI crude oil prices is close to the time of OPEC over production coupled with weak demand; and for heating oil, the break point of 2011 may have occurred following the European crisis.

The upshot from these results is that while there is clear evidence of nonstationary volatility present in all monthly and weekly energy prices, the presence of a trend break is not that evident for most of the energy prices considered in this study. Accordingly we divide the data into two separate parts to test for unit roots. For those energy prices that are only found to exhibit non-stationary volatility we apply the method of Smeekes and Taylor (2012). For the remaining energy prices that are characterised by non-stationary volatility as well as a trend break we apply the unit root test due to CHLT.

3. No evidence of further breaks are found when applying the sequential procedure due to Kejriwal and Perron (2010).

Figure 1: Plot of Energy Prices: Monthly Data

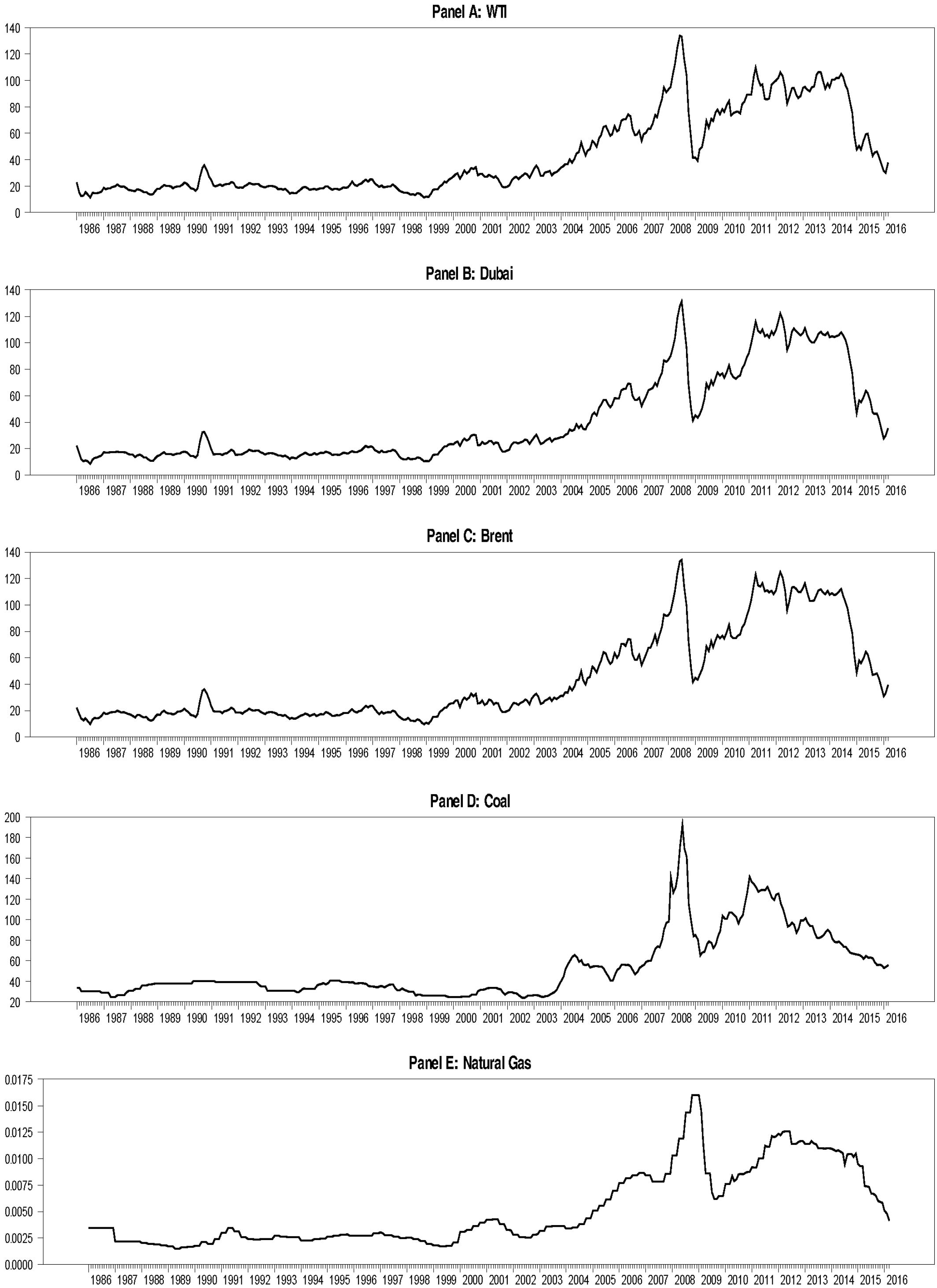

Figure 2: Plot of Energy Prices: Weekly Data

Table 1: Structural Breaks in Variance

	IT	SAC
Monthly data		
WTI	Nov 2002, Jul 2005	Aug 1999, Jul 2005
Brent	Aug 1999, Sep 2007	Aug 1999, Sep 2007
Dubai	July 1999, Oct 2007	July 1999, Oct 2007
Coal	Nov 1997, Aug 2004, Jun 2007, Apr 2012, Sep 2014	Oct 2007
Gas	Dec 1999, Dec 2005;	Dec 2007
Weekly data		
WTI	24 Oct 1986, 11 Sep 2015	24 Oct 1986, 20 Mar 2015
Brent	19 Dec 1997, 1 Mar 2002, 12 Sep 2003, 9 Dec 2005, 23 Feb 2007, 15 Jun 2007, 3 Apr 2009, 29 Dec 2012, 6 Sep 2013, 11 Jul 2014, 8 Aug 2014	24 Apr 1987, 19 Jul 1991, 19 Dec 1997, 9 Dec 2005, 11 Jul 2014
Gasoline	9 May 1986, 19 Jun 2015	9 May 1986, 27 Mar 2015
Heating	21 Aug, 1998, 25 Jul 2003, 12 Mar 2004, 8 Apr 2005, 27 Oct 2006, 30 Nov 2007, 18 Apr 2008, 13 Mar 2009, 5 Mar 2010, 13 Dec 2013, 8 Aug 2014.	21 Aug, 1998, 7 Mar 2003, 25 Jun 2004, 15 Jan 2010, 20 Sep 2013

Notes: IT denotes the Inclan and Tiao (1994) test while the SAC denotes the Sanso et. al. (2004) test for nonstationary volatility in the data. The dates recorded in the cells denote the periods when a break in variance is found.

Table 2: Robust Tests for Structural Breaks in Trend

	PY	*HLT*
Monthly data		
WTI	–0.18	2.30
Brent	–0.16	2.21
Dubai	–0.17	2.14
Coal	0.03	2.57
Gas	–0.15	3.11* (Jul 1998)
Weekly data		
WTI	4.17* (14 Jan 1994)	2.45
Brent	0.21	2.35
Gasoline	0.04	1.82
Heating	1.18* (24 Jun 2011)	2.56

Notes: The * sign denotes significance at the 10% level; implying rejection of the null of no break. The numbers in parentheses denote the break date. The column heading PY denotes the Perron and Yabu test for the robust detection of structural breaks, while the column heading HLT denotes the Harvey et. al. (2009a) test for the detection of structural breaks.

The CHLT test is based on a battery of M-type unit root tests being *MZa*, *MSB* and *MZt* tests. When we run the tests due to CHLT that allow not only for the possibility of a structural break but also for non-stationary volatility, we find that the null hypothesis of a unit root at the conventional significance levels cannot be rejected. These results are contained in Table 3.

In the first column of results in Table 3 we report the estimated test statistics for the CHLT test for unit roots. Comparing the tests statistics with the bootstrapped critical values at the 5% and 10% level, we find that the null hypothesis of a unit root cannot be rejected in all three of the energy prices considered. For example, considering natural gas prices, the estimated statistic of the *MZa* statistic is estimated as –10.7 which is greater than the bootstrapped critical values –15.27 and –12.66 at the 5% and 10% significance levels respectively. We therefore cannot reject the null of a

Table 3: Unit root tests with possible trend break and non-stationary volatility.

	MZa statistic			MSB statistic			MZt statistic		
		Bootstrapped Critical Value			Bootstrapped Critical Value			Bootstrapped Critical Value	
	statistic	5%	10%	statistic	5%	10%	statistic	5%	10%
Monthly data									
Gas	–10.7	–15.27	–12.66	0.21	0.17	0.19	–2.26	–2.72	–2.47
Weekly data									
WTI	–0.57	–63.81	–35.39	0.93	0.08	0.11	–0.53	–5.41	–3.98
Heating oil	–8.56	–79.34	–43.62	0.16	0.07	0.10	–1.42	–6.29	–4.44

Notes: The results are based on the unit root test procedure by Cavaliere et. al. (2011). The unit root tests labelled MZa, MSB and MZt statistics are the M-type tests. For all the tests, the null hypotheses is that of a unit root. The boot strapped critical values at conventional levels (that is 5% and 10%) are reported for each M-type test.

Table 4: Bootstrap union tests for unit roots in the presence of non-stationary volatility

	Test Statistic	UR_{4A} Bootstrap C.V. [p – value]	UR_{4B} Bootstrap C.V. [p – value]
Monthly data			
WTI	–1.154	–2.061 [0.778]	–1.965 [0.747]
Brent	–1.176	–2.015 [0.739]	–1.935 [0.703]
Dubai	–1.484	–2.071 [0.467]	–1.998 [0.429]
Coal	–1.225	–2.118 [0.704]	–2.060 [0.705]
Weekly data			
Brent	–1.866	–1.965 [0.134]	–1.866 [0.106]
Gasoline	–1.087	–1.978 [0.544]	–2.074 [0.554]

Notes: The results are based on the unit root test procedure by Smeekes and Taylor (2012). The column headings UR_{4A} and UR_{4B} denote the unit root tests that include and exclude a trend respectively. The bootstrapped critical values are calculated at the 10% significance level. The numbers in square brackets denote probability values.

unit root. In the case of the *MSB* as well as the *MZt* statistics we obtain the same result, the estimated statistics (0.21 and –2.26 respectively) is greater than the bootstrapped critical values leading to non-rejection of the null hypothesis of a unit root. We conclude that for natural gas prices (measured at monthly frequencies) and WTI along with Heating oil (measured at weekly frequencies) we cannot rule out the possibility that the effect of a shock is likely to be permanent.

Table 4 shows the results of the bootstrapped union tests for unit roots on the remaining six energy prices. In the first column of results in Table 4 we report the estimated unit root test statistics. These are compared with the bootstrapped critical values that are computed at the 10% level of significance and reported in the second and third column of results in Table 4. The second and third columns of results report the UR_{4A} and UR_{4B} union tests. The associated probability values are recorded in square brackets alongside the critical values.

Given the estimated statistics that we obtain we cannot reject the unit root null hypothesis. For example, for WTI, the probability value is 0.778 for the UR_{4A} test and 0.747 for the the UR_{4B} test indicating that we cannot reject the null hypothesis of a unit root in the data. This is true for all the other energy prices using both the UR_{4A} and UR_{4B} tests. We conclude that for all the six energy prices, that is, WTI, Brent Dubai/Oman and coal (measured at monthly frequencies) and Brent and Gasoline (measured at weekly frequencies) we reject that the effect of a shock to these prices is transitory. Overall, we find from the results in Tables 3 and 4, that for all the energy prices considered, shocks to these prices are not transitory in nature.

Our results are in sharp contrast with some of the very recent studies that have been conducted into inquiring whether shocks to energy prices are transitory or permanent. Most of the recent studies such as Postali and Pichetti (2006), Ghoshray and Johnson (2010), Noguera (2013),

Table 5: Correlation measures on decomposed transitory and permanent components.

Correlation between permanent and transitory components	Demand	Supply
Correlation between permanent industrial prod. and permanent oil price	0.51 (0.046)	
Correlation between permanent industrial prod. and transitory oil price	–0.56 (0.052)	
Correlation between transitory industrial prod. and permanent oil price	–0.66 (0.039)	
Correlation between transitory industrial prod. and transitory oil price	0.69 (0.043)	
Correlation between permanent oil supply and permanent oil price		–0.74 (0.076)
Correlation between permanent oil supply and transitory oil price		0.68 (0.083)
Correlation between transitory oil supply and permanent oil price		0.37 (0.062)
Correlation between transitory oil supply and transitory oil price		–0.29 (0.055)

Notes: The decomposition method and correlations are calculated based on the procedure due to Sinclair (2009). Numbers in parentheses denote standard errors.

Narayan and Liu (2015) and Sun and Shi (2015) have broadly concluded in favour of energy prices being stationary, or in other words shocks to energy prices being transitory in nature. Our results are in line with that of Maslyuk and Smyth (2008) that conclude shocks to energy prices are not transitory in nature. One reason for this departure from past studies is that the focus in those studies has been on the incorporation of structural breaks. However, it needs to be noted that there is limited evidence of a trend break in energy prices. We find a single trend break for only three out of the nine possible cases with no evidence of further breaks. Also, energy prices have experienced prominent movements in recent months and weeks which may lead to the overall detection of a slope break to be non-existent. Our results are robust to the start date of the chosen time period as we take into account the initial condition; which to our knowledge has not been considered in recent studies.

Finally, given the finding that shocks to oil prices are not transitory in nature we decompose oil prices into permanent and transitory components. As discussed earlier, relationships can exist between oil prices and its demand, and supply, in the spirit of Dvir and Rogoff (2014) for example, who formulate a model to argue that permanent demand increases will lead to permanent price increases. The decomposition of these variables into their permanent and temporary components allows us to compute cross correlations which can be helpful for economic modelling of oil prices. The method of decomposition and calculation of correlations follows that of Sinclair (2009) where the economic variables (say y_t, where y_t can be oil prices, industrial production, or oil stocks and production) be represented as the sum of a permanent component ζ_t and a temporary component μ_t. The permanent component is assumed to be a random walk, given by $\zeta_t = \zeta_{t-1} + \nu_t$ and the transitory component is modelled as an AR(2) process, $\mu_t = \lambda_1 \mu_{t-1} + \lambda_2 \mu_{t-2} + \eta_t$. The error terms from the components, that is, ν_t and η_t are assumed to be jointly normally distributed with mean zero and a general covariance matrix, and the maximum likelihood estimation is carried out using a Kalman filter (see Sinclair 2009 for details). The results of the cross correlations are given in Table 5.

We choose the benchmark oil price WTI, to determine the nature of its correlation with demand shocks (measured by industrial production) and supply shocks (measured by the combined level of production and stocks). We have already shown that the for oil prices we cannot reject the null hypothesis that any shocks to these variables are not transitory in nature. We find that we cannot reject the null hypothesis of a unit root using a battery of tests for industrial production and oil supply data[4]. Given that WTI, industrial production and oil supply are all found to be I(1), we can proceed to decompose the data into its trend and cyclical components and compute the cross correlations between these components based on the procedure by Sinclair (2009).

First, considering the demand side, we find that there is a positive correlation between the permanent component of oil prices and industrial production. Also a positive correlation is found between the transitory component of oil prices and industrial production. Both correlations are significant. When calculating the correlations between transitory and permanent components for both the data series, we find they turn out to be negative and significant. To explain this we can use the model of an inverse net demand function that is defined as the excess of the supply function over the demand function; the net demand function is the change in inventories which comprises of oil prices, and demand and supply curve shifters (see Pindyck 1999 for details). Consider the case where a demand shock is permanent. Assuming that there is a permanent increase in demand as a result of increased industrial production, if the expectation is that the shock will be permanent then there will be no change in inventories and production will increase to meet the increase in demand. Under these circumstances there will be a permanent increase in prices which explains the positive significant correlation. Now consider the case where the demand shock is expected to be transitory. If there is a transitory increase in demand, then the expectation is that this increase will be short-lived. In this case in the short run, inventories are drawn down and price increases, which is relatively smaller than the amount by which price increases when there is a permanent demand shock. Hence a transitory increase in demand is accompanied by a transitory increase in price. However, when the demand shock dissipates (as we would expect from a transitory shock) the inventories build up and the demand for production in excess of consumption falls, leading to a price decline. Hence a transitory demand shock is negatively correlated with the permanent component of oil price.

Now considering the supply side, we find that there is a negative and significant correlation between the permanent component of oil prices and oil supplies. Also a positive significant correlation is found between the transitory component of oil prices and oil supply. Again, this can be explained using the inverse net demand function due to Pindyck (1999). If there were to be a permanent negative supply shock (for example, oil supplies are restricted by cartels in the face of increased demand), then that will lead to an increase in prices which will be based on a shift in the inverse net demand function. This explains the negative correlation between the permanent component of oil prices and the permanent component of oil supplies. Conversely, if there is a temporary negative supply shock, then carryover stocks build up leading to a temporary increase in price which explains the negative correlation between the transitory components of the two series. Since the supply shock is deemed temporary, the shock ultimately dissipates and the price falls back to its original level. This explains the positive correlation between a negative transitory shock in oil supplies and the permanent component of the oil price.

5. CONCLUSION

This paper primarily examines whether shocks to energy prices are permanent or transitory. The results have significant consequences for energy related econometric analysis, forecasting, investment decisions and macroeconomic policy-making. Using novel and appropriate unit root tests on energy prices shows that we cannot reject that shocks to energy prices are not transitory. This result is contrary to recent empirical studies where the conclusion is that shocks to energy prices are transitory in nature. Dvir and Rogoff (2009) offer an informal historical narrative as well as a

4. The results are available from the author on request.

formal model that explains why energy prices are likely to be highly persistent and volatile. The narrative is based on the reasoning that during the period under consideration, which is from the 1980s to recent times, there has been intense industrialization in the East, mainly by China which had an effect on the global economy as a whole. They argue that rapid industrialization is a period which is considered as a transition that stretches over a considerable period of time and is therefore characterised by growth shocks which are highly persistent. This in turn leads to persistent changes in oil prices. On the supply side, due to the monopoly of the OPEC countries, there was uncertainty regarding the future supply of oil. This is due to the fact that in the 1970s oil production in the U.S. had reached maximum capacity and excess capacity existed in the OPEC countries. During this time, OPEC countries began to act collusively to restrict energy commodity supplies in a bid to extract large rents (Smith 2008). Hence, in a period of rising demand, the OPEC was able to restrict access to additional supplies of oil which may have contributed to the persistence in oil and other energy prices. The formal model by Dvir and Rogoff (2009) predicts that rational storage behaviour creates higher volatility in oil prices particularly in the presence of uncertainty regarding the trend. Based on this narrative and formalised model of the structure of the oil market by Dvir and Rogoff (2009) we can offer plausible reasons to support the findings of high persistence as well as changing volatility in energy prices.

We add to the analysis further, by considering the decomposition of a benchmark oil price, (WTI in this case), the decomposition of a demand side variable, (industrial production), and a supply side variable (oil supply) into their permanent and transitory components. The cross correlations between these individual components manages to capture commodity market dynamics, with storage, in the spirit of the model by Pindyck (1999). We do not elaborate the analysis further as it is beyond the scope of the paper to expand on this issue of demand and supply led shocks. Besides, as Dvir and Rogoff (2009) have highlighted: demand and supply shocks in the canonical storage model are isomorphic and the emphasis should be on building an understanding of the structure of the oil industry rather than identifying the source of the shocks. However, separate studies using structural models with newly developed measures of global economic activity have been employed to explain relationships between the oil price shocks and the economy. Out of several studies, the classic study by Kilian (2009) is a case in point. The upshot is that this remains a pertinent question and would be a subject of further research.

What can be learned from the observed nature of persistence in energy price shocks? After the oil price shock in the early 1970s, the price controls imposed by the US government may have been necessary if the prices were not to revert back to the long term trend. Similarly, the OPEC production quotas of the early 1980s may have not have been redundant if oil prices were anticipated to exhibit a stochastic trend. The findings have strong implications for investment decisions. If shocks to energy prices were indeed to have a permanent effect, then considering the long investment horizon in oil fields, it would imply that we cannot easily ignore the significant deviations from trend. The cumulative evidence of volatility alongside the persistence found in energy prices highlights the additional aspect of ‘uncertainty in persistence’, which may delay irreversible investment decisions Hamilton (1996). If shocks to energy prices are permanent, then there are possibilities for energy prices to create instabilities of a sort, in particular to emerging countries that are reliant on energy commodities as intermediate inputs. This could lead to increased vulnerability in emerging economies as energy shocks are more likely to translate to disused industrial capacity (Rogoff 2006). An examination of the time series properties of energy prices remains a crucial issue for the macroeconomic policy maker, and this study provides a more likely characterisation of the dynamic properties of the data.

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