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**Structural Break, Nonlinearity and the Hysteresis hypothesis:
Evidence from new unit root tests.**

Zarina Oflaz¹

Abstract

We have checked whether there exists hysteresis hypothesis in 25 OECD countries covering the period from September, 1983 to September, 2013 by using the unit root test which includes nonlinearity and structural break in their testing process. In particular, LNV test proposed by Leybourne, Newbold and Vougas (1998), EG test developed by Enders and Granger (1998), the nonlinear KSS unit root test of Kapetanios et al., (2003), Sollis' (2009) test, combined KSS-LNV test developed by Omay and Yildirim (2014), combined LNV-Sollis test by Omay et al (2017), modified unit root test with Fractional Frequency Flexible Fourier Form by Omay (2015) and EST test by Corakci et al., (2017).

Keywords: Hysteresis hypothesis; Structural Break; Nonlinearity.

JEL Codes: C23, J64.

1. Introduction

The fact that labor markets in the OECD countries have never recovered the full employment levels after the several oil price shocks occurred between 1978 and 1983. Afterward, the labor market performance declined significantly due to global financial and economic crisis in 2008 and 2009. Nowadays, the unemployment rates stay at the higher level compared to 2008- 09 crisis level in many OECD countries. Evidently, that high unemployment levels may lead to the decline of output and have a negative effect on the economic development of the country.

The main difficulty behind the high unemployment rates is whether the rise caused by cyclical factors that can be returned to its equilibrium state as the economic recovery becomes stronger or by hysteresis phenomenon that cannot be eliminated even after the full recovery of the economy. Thus, the remedy for the non-negligible labour market issues is directly related to the mentioned question.

Experts distinguish two competing hypothesis describing the unemployment behaviour: the *natural rate hypothesis (NRH)* and the *hysteresis hypothesis (HH)*.

The first hypothesis states there is a temporary effect of the cyclical factors to the unemployment rate in the short-run and the economy will converge to a unique and stable equilibrium unemployment rate in the long-run. The equilibrium unemployment rate is named as the natural rate of unemployment, which is regulated according to the structure of and frictions in the labour market. The NRH indicates that, high current unemployment rates can be caused by one of the following scenarios. When the economy exhibits a negative growth, the disturbances which are producing cyclical fluctuations of the unemployment rates may push those above its natural rate. The theory supposes that in this case the economy will turn back to its low natural rate over the long run. The second case shows that shocks that alter the structure of the labor market may increase the natural rate itself, therefore, the unemployment rate will stay at the new high level of equilibrium unemployment rate permanently. It is important to emphasize, that in accordance with the hypothesis, the long-run Phillips curve, which represents the relationship between the inflation rate and unemployment rate, is vertical, i.e. these macroeconomic variables are unrelated. Thus, the natural rate is independent of monetary policy actions.

The NRH was quite eligible theory to describe the labor market fluctuations in developed countries, however the sharp rise of the oil prices had a negative and permanent effects on the actual rate of unemployment. This circumstance leads to the necessity of the natural rate theory reconsideration.

The second hypothesis, namely *hysteresis*, is defined as a process which exhibit a dependency on past shocks, i.e. reversing the shock does not recover initial properties of the system even over the long-run. The memory is described as nonlinear, selective, and with remanence (H. Hallett, L. Piscitelli, 2002). Remanence means that the temporary shocks may cause permanent outcomes, thus moving the equilibrium rate from one level to another. According to (H. Hallett, L. Piscitelli, 2002), selectivity determines only the shocks ‘which are ‘not dominated’ by shocks which come later’.

According to HH, cyclical unemployment would be transformed into structural one (B. Bartlett, 2013), that will cause the permanent consequences of unemployment rate rise. Various factors may contribute to structural unemployment. For example, the skills of employees no longer valuable since the industries which hired them do not exist; the human workforce is replaced by computers and automatic machinery.

NRH requires that the unemployment series are stationary or mean reverting around a unique equilibrium unemployment rate. However, the hysteresis hypothesis determines the unemployment series as nonstationary and containing a unit root, meaning that the temporary shocks result the unemployment rate to fluctuate and have no tendency for the unemployment rate to return to its equilibrium level. A following equation is used in order to review these hypotheses

$$u_t = \alpha + \beta u_{t-1} + \varepsilon_t \quad (1)$$

where the unemployment series u_t follows an autoregressive process of order 1, i.e. AR(1) and ε_t is a white noise error term. For the NRH, a unit root in u_t occurs in equation (1) when $H_0 : \beta = 1$. Thus, the natural rate hypothesis holds in its weak form when the null hypothesis is rejected against the alternative $H_0 : \beta < 1$. For the NRH to hold in its strong form the null $H_0 : \beta = 0$ should be tested since the unemployment rates are assumed not to be a function of the past unemployment rate. Otherwise, if the null $H_0 : \beta = 1$ cannot be rejected, the unemployment rates have a hysterical behavior.

2. Empirical Literature

The majority of the previous studies conducted the modelling of the unemployment hysteresis within a linear unit-root framework. Commonly used linear unit root tests, like Dickey-Fuller (ADF) test, Phillips-Perron (PP) test, KPSS test of Kwiatkowski et al. (1992) and/or the DF-GLS test of Elliott et al. (1999) were implemented to examine for the hysteresis effects in the unemployment rate series. The unemployment rates in the US and its states were detected to be non-stationary using the ADF and the PP tests (Breitung (1994), Hatanaka (1996). Song and Wu (1998) and Leon-Ledesma (2000)). Applying the classical linear univariate tests, it has been failed to reject the hysteresis hypothesis in most EU countries (Blanchard and Summers, 1986; Groenewold and Taylor, 1992; Roed, 1996; Leon-Ledesma, 2000; Camarero and Tamarit, 2004).

It should be noted that various ADF test statistics are biased towards the non-rejection of the unit root hypothesis when the structural breaks take place. There are a large number of studies considering the structural breaks. The plenty number of these studies succeeded to reject the unit root null (Camarero and Ordonez; 2006; Lee and Chang; 2008; Furuoka, 2014a), however the majority of the studies were unable to reject the hysteresis effect in the unemployment rates (Chang et al., 2007; Chang; 2011; Cuestas et al., 2011; Cheng et al., 2014; Furuoka, 2014b). Moreover, having low power in small samples, particularly when the unemployment series are near-unit root processes, the conventional univariate unit root tests are less reliable and accurate.

The mentioned studies consider the linear behaviour of the unemployment rates and believe that those demonstrate symmetric adjustment paths. However, nonlinear behaviour of the key macroeconomic variables is important from a theoretical perspective. The observations, which wages are rigid downward, have an important meaning for labour markets. Moreover, business cycles have downturns which are steeper than upturns in the main economic variables,

such as output and employment. Thus, another group of scientists test hysterical unemployment behaviour within a nonlinear framework. There are some studies, which assume that the unemployment rates follow a threshold Autoregressive (TAR) process (Coakley et al., 2001; Perez-Alonso and Sanzo, 2011), smooth transition autoregressive (STAR) process (Skalin and Teräsvirta, 2002; He and Sandberg; 2012) or a Markov switching process with multiple equilibria (Bianchi and Zoega, 1998; Leon-Ledesma and Mc Adam, 2004).

Furthermore, new univariate nonlinear unit root tests have been developed. For example, TAR models have been used in order to construct new unit root test. (Enders and Granger, 1998; Caner and Hansen, 2001), also, ESTAR model have been used to establish unit root tests (Kapetanios et al., 2003). The mentioned studies have determined that most European countries are distinguished by the non-stationary unemployment rates, however for the US other studies have rejected the hysteresis hypothesis (Gustavsson and Österholm, 2006; Ghosh and Dutt, 2008; Yilanci, 2008; Lin et al., 2008; Chang and Lee; 2011; Tiwari; 2014, Tartıcı 2015).

In our study we give a brief review of different nonlinear unit root tests, in particular, LNV test proposed by Leybourne, Newbold and Vougas (1998), EG test developed by Enders and Granger (1998), the nonlinear KSS unit root test of Kapetanios et al., (2003), Sollis' (2009) test, combined KSS-LNV test developed by Omay and Yildirim (2014), combined LNV-Sollis test by Omay et al (2017), modified unit root test with Fractional Frequency Flexible Fourier Form by Omay (2015) and EST test by Corakcı et al., (2017).

The aim of our study is to apply the aforementioned models to test the hysteresis hypothesis in the unemployment rates of 25 OECD countries and compare the implementation of the unit root tests. Thus, we have improved upon the Tartıcı (2015) study by increasing the country and sample period. On the other hand, we also have also used more recent unit root test in order to increase the understanding of the stationarity properties of the unemployment data.

The paper is organized as follows: the next section outlines the theoretical framework of the developed nonlinear unit root tests. The Section 3 and 4 present data description and results of applied tests. Finally, in the last section we report the main conclusions.

3. The model and testing procedure

Equation (1) can be improved in many different ways by adding the deterministic function, allowing for structural breaks, producing asymmetric mean reversion, and assuming the serial dependence of the error terms. In this section we a short review of existing nonlinear unit root tests.

TAR threshold Autoregressive type of unit root test. EG test

Enders and Granger. (1998), have proposed the unit root test employing Tong's (1983) threshold autoregression (TAR) model. Let the unemployment rate, u_t , follow the p th order TAR model:

$$\Delta u_t = \alpha + \rho_1 I_t u_{t-1} + \rho_2 (1 - I_t) u_{t-1} + \sum_{j=1}^{p-1} \hat{\delta}_j \Delta u_{t-j} + \varepsilon_t \quad (2)$$

$t = 1. \dots T$, where I_t is the Heaviside indicator function defined as

$$I_t = \begin{cases} 1, & \text{if } u_{t-1} \geq 0 \\ 0, & \text{if } u_{t-1} < 0 \end{cases}$$

ε_t are zero-mean iid error terms. The null hypothesis $H_0: \rho_1 = \rho_2 = 0$, indicates the unit root process. Under the alternative hypothesis $H_1: \rho_1 < 0$ and $\rho_2 < 0, \rho_1 \neq \rho_2$, authors consider the stationarity with asymmetric adjustment towards a constant mean or deterministic trend.

Symmetric ESTAR test. KSS test

Kapetanios *et al.*, (2003) developed the unit root test against the nonlinear but globally stationary process. Let us consider the following model, namely an exponential smooth transition model (ESTAR)

$$u_t = \beta u_{t-1} + \gamma u_{t-1} \left[1 - e^{-\theta u_{t-d}^2} \right] + \varepsilon_t \quad \text{for } t = 1, \dots, T. \quad (3)$$

where u_t is a stochastic process with unknown parameters γ , ε_t is a white noise disturbance. The transition function is of the exponential form, where $d \geq 1$ is the delay parameter, the slope parameter θ is positive, which defines the transition speed between the two extreme regimes, namely “upper” and “lower” regimes, which correspond to the two extreme values of the transition function.

The Equation (3) can be rewritten with the assumption that $\beta = 1$. Thereafter, Kapetanios *et al.*, (2003) set $\phi = 0$, indicating that the process is in the middle regime, also $d=1$ is selected to maximize goodness of fit. The ESTAR model with aforementioned assumptions is given as follows;

$$\Delta u_t = \gamma u_{t-1} \left[1 - e^{-\theta u_{t-1}^2} \right] + \varepsilon_t. \quad (4)$$

In order to decide whether the process is linear or unit root the null hypothesis $H_0: \theta = 0$ is implemented against the alternative $H_1: \theta > 0$, meaning that the process has the nonlinear globally stationary behaviour. By using the first order Taylor approximation this test become more feasible

$$\Delta u_t = \sum_{j=1}^p \rho_j \Delta u_{t-j} + \delta u_{t-j}^3 + e_t \quad (5)$$

where e_t includes original shocks ε_t as well as the error term resulting from Taylor approximation. The t-statistic, namely the KSS test, for $H_0: \delta = 0$ against $H_1: \delta < 0$ is defined as

$$t_{NL} = \frac{\hat{\delta}}{s.e.(\hat{\delta})}, \quad (6)$$

with the OLS estimate, $\hat{\delta}$ and the standard error of $\hat{\delta}$, $s.e.(\hat{\delta})$.

Asymmetric ESTAR test

Sollis, (2009) have developed the unit root test applying an extension of the ESTAR model, which is able to produce both symmetric and asymmetric mean reversion

$$\Delta u_t = G_t(\gamma_1, u_{t-1}) \{S_t(\gamma_2, u_{t-1}) \rho_1 + (1 - S_t(\gamma_2, u_{t-1})) \rho_2\} u_{t-1} + \varepsilon_t \quad (7)$$

where the transition functions are of the exponential and logistic forms

$$\begin{aligned} G_t(\gamma_1, u_{t-1}) &= 1 - e^{-\gamma_1 u_{t-1}^2}, & \gamma_1 &\geq 0 \\ S_t(\gamma_2, u_{t-1}) &= \frac{1}{1 + e^{-\gamma_2 u_{t-1}}}, & \gamma_2 &\geq 0 \end{aligned} \quad (8)$$

ε_t is a zero mean i.i.d. error term. The model is asymmetric, if $\rho_1 \neq \rho_2$, and symmetric otherwise. Here, we assume only asymmetric version of ESTAR.

A first order Taylor expansion for ESTAR is used since there are unknown parameters under the null hypothesis, $H_0: \gamma_1 = 0$.

$$\Delta u_t = \rho_1 \gamma_1 u_{t-1}^3 S_t(\gamma_2, u_{t-1}) + \rho_2 \gamma_1 u_{t-1}^3 (1 - S_t(\gamma_2, u_{t-1})) + \eta_t \quad (9)$$

Finally, an augmented version of ESTAR is

$$\Delta u_t = \phi_1 u_{t-1}^3 + \phi_2 u_{t-1}^4 + \phi_t \quad (10)$$

The null hypothesis becomes as follows

$$H_0: \phi_1 = \phi_2 = 0 \quad (10)$$

against the alternative hypothesis, $H_1: \phi_1 < 0$ and $\phi_2 < 0$, $\phi_1 \neq \phi_2$ of stationary asymmetric nonlinear ESTAR for testing the unit root null hypothesis. Solis, (2009) derived the asymptotic distribution of an F-test and asymptotic critical values of the test statistics.

LNV test

Leybourne, Newbold and Vougas, (1998) proposed LNV test based on the following regression models with the logistic smooth transition function

$$\begin{aligned}
 \text{Model A} \quad u_t &= \alpha_1 + \alpha_2 S_t(\gamma, \tau) + v_t \\
 \text{Model B} \quad u_t &= \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + v_t \\
 \text{Model C} \quad u_t &= \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + v_t
 \end{aligned} \tag{11}$$

where the logistic smooth transition function is defined as follows²

$$S_t(\gamma, \tau) = \left[1 + \exp\{-\gamma(t - \tau T)\} \right]^{-1} \quad \gamma > 0 \tag{12}$$

The test statistics is established by two steps.

First step:

$$\begin{aligned}
 \text{Model A} \quad \hat{v}_t &= u_t - \hat{\alpha}_1 - \hat{\alpha}_2 S_t(\hat{\gamma}, \hat{\tau}) \\
 \text{Model B} \quad \hat{v}_t &= u_t - \hat{\alpha}_1 t - \hat{\beta}_1 - \hat{\alpha}_2 S_t(\hat{\gamma}, \hat{\tau}) \\
 \text{Model C} \quad \hat{v}_t &= u_t - \hat{\alpha}_1 t - \hat{\beta}_1 - \hat{\alpha}_2 S_t(\hat{\gamma}, \hat{\tau}) - \hat{\beta}_2 t S_t(\hat{\gamma}, \hat{\tau})
 \end{aligned} \tag{13}$$

At the first step the deterministic component of the selected model is estimated applying a nonlinear least squares (NLS) algorithm, the NLS residuals are obtained.

Second step: Calculate the ADF statistic

$$\Delta \hat{v}_t = \hat{\rho} \hat{v}_{t-1} + \sum_{i=1}^k \hat{\delta}_i \Delta \hat{v}_{t-i} + \hat{\eta}_t \tag{14}$$

The following unit root hypotheses are considered:

Null hypothesis	$u_t = \mu_t, \mu_t = \mu_{t-1} + \varepsilon_t, \mu_0 = \psi$
Alternative hypothesis	Model A, Model B or Model C
Null hypothesis	$u_t = \mu_t, \mu_t = \kappa + \mu_{t-1} + \varepsilon_t, \mu_0 = \psi$
Alternative hypothesis	Model B or Model C.

KSS-LNV test

Omay and Yildirim, (2014) proposed new nonlinear unit root test by combining aforementioned KSS test by Kapetanios et al (2003) and LNV test by Leybourne, Newbold and Vougas (1998). The expansion of the existing models is that the structural change is modeled as

² For the Nonlinear optimization algorithm which is used in LNV type of de-trending we have followed the suggestion which are given in Omay and Emirmahmutoğlu (2017).

a smooth transition between different regimes. while Leybourne et al. (1996) consider those as a rapid structural break. The authors establish the linear unit root null hypothesis against nonlinear stationary process around smoothly varying trend and intercept under the alternative hypothesis. The *First Step* is similar to the first step of the LNV model.

Second Step: the KSS statistic is computed

$$\Delta \hat{\varepsilon}_t = \hat{\rho} \hat{\varepsilon}_t^3 + \sum_{j=1}^k \hat{\delta}_j \Delta \hat{\varepsilon}_{t-j} + \hat{\eta}_t \quad (15)$$

The null hypothesis $H_0: \rho = 0, \text{ for all } i$, holds for linear nonstationary process against the alternative, $H_a: \rho < 0, \text{ for some } i$, (Nonlinear and Stationary around nonlinear trend and intercept).

LNV-Solis test

Omay et al, (2017) proposed a test which permits for simultaneous structural change and asymmetric nonlinear adjustment towards the equilibrium level. Three aforementioned smooth transition models, see Equations (11, 12) are considered with logistic transition function which model gradual structural breaks. In order to model adjustments towards the equilibrium asymmetric ESTAR (AESTAR) nonlinearity is used, see Equations (7,8). As in Kapetanios et al., (2003) and Sollis (2009), the transition functions were replaced by their first-order Taylor series approximation, see Equations (9,10).

Omay et al., (2017) assumes a serial correlation of the error terms, assuming that those enter in a linear way as follows

$$\Delta u_t = G_t(\gamma_1, u_{t-1}) \left\{ S_t(\gamma_2, u_{t-1}) \rho_1 + (1 - S_t(\gamma_2, u_{t-1})) \rho_2 \right\} u_{t-1} + \sum_{j=1}^p \delta_j \Delta u_{t-j} + \varepsilon_t \quad (16)$$

with zero mean error terms, ε_t . Therefore, the auxiliary regression equation becomes

$$\Delta u_t = \phi_1 u_{t-1}^3 + \phi_2 u_{t-1}^4 + \sum_{j=1}^p \delta_j \Delta u_{t-j} + \varepsilon_t \quad (17)$$

The null hypothesis $H_0: \phi_1 = \phi_2 = 0$, tests for a unit root process against the alternative hypothesis, $H_1: \phi_1 < 0 \text{ and } \phi_2 < 0, \phi_1 \neq \phi_2$, of globally stationary AESTAR nonlinearity.

The null hypothesis testing can be performed in two steps. First, estimate the deterministic component of the preferred model using NLS algorithm and then collect residuals $\hat{u}_t$. Second, employing these residuals estimate the auxiliary regression equation by ordinary least squares and test the null hypothesis by F-test. The critical values for the test are displayed in Omay et al., (2017).

Unit root test with FFFFF

The unit root test with Fractional Frequency Flexible Fourier Form (FFFFF) was improved by Omay, (2015) by combining the methods of Becker et al. (2004) and Enders and Lee (2012b). The Fourier approach is capable to model multiple smooth breaks when the deterministic function is of unknown form and even is not periodic.

Consider the Dickey-Fuller test

$$u_t = d(t) + \phi_1 u_{t-1} + \lambda t + \varepsilon_t \quad (18)$$

where ε_t is a stationary disturbance term with constant variance σ^2 . $d(t)$ is a deterministic function of t .

Omay, (2015) claims that the deterministic function can be approximated by applying the Fourier transform

$$d(t) = \alpha_0 + \alpha \sin\left(\frac{2\pi kt}{T}\right) + \beta_k \cos\left(\frac{2\pi kt}{T}\right) \quad (19)$$

where k is a particular frequency and T is the number of observations. According to the literature, the single frequency is recommended to use since it leads to an optimal approximation to a model with structural change. Fractional frequency is employed instead of integer ones, that allows to establish an appropriate nonlinear trend into the unit root test.

The testing regression for fractional single frequency is

$$\Delta u_t = \rho u_{t-1} + c_1 + c_2 t + c_3 \sin\left(\frac{2\pi k^{fr} t}{T}\right) + c_4 \cos\left(\frac{2\pi k^{fr} t}{T}\right) + e_t. \quad (20)$$

The critical values for the test are tabulated in Omay, (2015).

ÇEO test

Corakcı et al., (2017) consider three regression models, see Equations (11) and employ an exponential smooth transition (EST) function which allows for temporary breaks.

$$S_t(\gamma, \tau) = 1 - \exp\left[-\gamma(t - \tau)^2\right], \quad \gamma > 0 \quad (21)$$

The null hypothesis tests for linear non-stationarity against the alternative including nonlinear stationary (stationary around smoothly changing trend and intercept). The calculation of the test statistics is based on the two step procedure as in Leybourne, Newbold and Vougas (1998)³.

³ The panel extension of EG, Solis (2004, 2009) and LNV can be found in Çorakcı *et al* (2017) (Hereafter, ÇEO), Omay *et al* (2017) (Hereafter, OÇE), Emirmahmutoğlu and Omay (2014), Omay *et al* (2017) (Hereafter, OHS), respectively. For the further study we can use these panel unit extensions, as well.

Data description

Data includes monthly unemployment rates of 25 OECD countries. The unemployment rates in 14 countries are observed from September, 1983 to September, 2013; Chile (from January, 1986), Spain (from April, 1986), Mexico (from January, 1987), Finland (from January, 1988), Norway (from January, 1989), Korea (from January, 1990), Germany (from January, 1991), Austria (from January, 1993), Hungary (from January, 1996), Poland (from January, 1997), Greece (from April, 1998).

In the following Table 1, the descriptive statistics and the number of observations for each country are observed.

Table 1. Descriptive statistics

Country	Values number	Minimum value	Maximum value	Median	Mean	Variance	Standard deviation
Australia	369	3.97	11.21	6.84	7.14	3.65	1.91
Austria	249	3.40	5.50	4.30	4.29	0.02	0.05
Belgium	369	6.30	11.00	8.30	8.41	1.57	1.25
Canada	369	5.90	12.70	7.90	8.50	2.87	1.70
Chile	333	5.44	14.54	7.99	8.31	2.80	1.67
Denmark	369	3.10	9.90	5.80	6.14	2.53	1.59
Finland	309	2.90	17.60	8.70	9.29	12.97	3.60
France	369	6.90	11.30	9.20	9.33	1.13	1.06
Germany	273	5.20	11.50	8.20	8.17	2.56	1.60
Greece	186	7.20	27.50	10.55	12.38	27.84	5.28
Hungary	213	5.50	11.30	7.60	8.14	3.86	1.97
Ireland	369	3.70	17.10	13.10	11.00	22.85	4.78
Italy	369	5.80	12.50	8.90	9.10	2.27	1.51
Japan	369	2.00	5.50	3.70	3.65	1.14	1.07
Korea	285	1.90	8.20	3.30	3.45	1.52	1.23
Luxembourg	369	1.40	6.00	2.90	3.27	1.44	1.20
Mexico	321	2.10	7.60	3.60	3.82	1.16	1.08
Netherlands	369	2.50	8.30	5.10	5.16	2.21	1.49
Norway	297	2.30	6.80	3.70	4.22	1.59	1.26
Poland	201	6.80	20.30	11.10	13.24	19.07	4.37
Portugal	369	3.90	17.60	7.20	7.76	9.28	3.05
Spain	330	7.90	26.60	15.95	15.87	23.02	4.80
Sweden	369	1.30	10.50	6.50	6.08	6.73	2.59
UK	369	4.60	11.30	7.70	7.59	4.43	2.11
USA	369	3.80	10.40	5.80	6.29	2.57	1.60

According to the Table 1, Greece and Spain display the highest unemployment rates, 27.5% and 26.6%, respectively. The unemployment rates in Finland, Poland, Ireland, Spain, and

Greece shows considerably unstable behavior while Austria, Japan, Korea, Luxembourg, Mexico, Norway are characterized by the quite balanced macroeconomic situation.

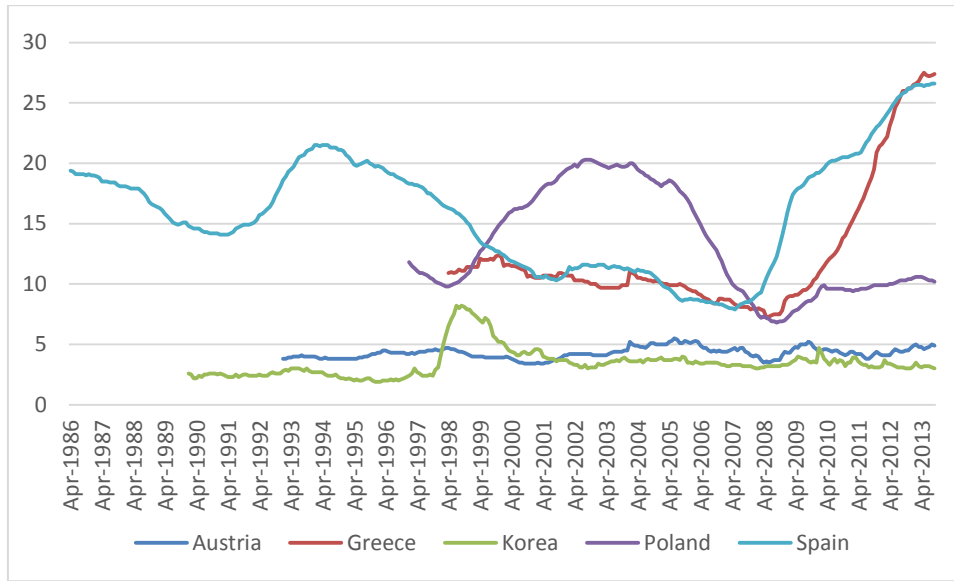


Figure 1. The unemployment rates of 5 countries: Austria, Greece, Korea, Poland, and Spain

In Figure 1, we compare the unemployment rates (in percent) of the selected countries. We observe that both Spain and Greece suffers a significant rise of the unemployment rate from 2008 until 2013, in comparison the unemployment rates of Korea and Austria are quite stable and low. The joblessness in Poland increases sharply from about 10% to over 20% between 1998 and 2003, then the decline of those follows until 2008, later the unemployment rate remains at about 10% from 2010 to 2013.

An Empirical Application

In this section we apply all the unit root tests which include the nonlinearity and structural break in their testing procedure to examine the hysteresis hypothesis in OECD countries. The empirical results of the conducted unit root tests are presented and discussed. Moreover, all applied unit root tests assume that the error terms are serially uncorrelated.

Table 2. Unit Root Test Results

Country	ADF	LNV	Sollis	KSS	EG	LNV-KSS	EST test	FFFFF test	LNV-Sollis
Australia	-1.193	-3.628	0.465	-1.629	1.287	-2.055	-1.641	-3.740***	5.409
Austria	-1.493	-1.198	-0.606	-1.174	2.367	-0.927	-1.439	-1.235	2.313
Belgium	-2.106	-2.452	1.364	-3.324*	6.643*	-1.768	-3.596	-3.467	2.747
Canada	-2.321	-2.549	-1.030	-3.500*	3.148**	-3.073	-2.377	-2.729	8.508**
Chile	-3.336**	-2.610	-0.084	-3.025*	3.932**	-3.392	-2.664	-3.382**	5.231
Denmark	-1.745	-2.564	0.934	-3.000*	2.763	-3.127	-2.531	-4.074**	8.352**

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Finland	-1.694	-3.566	-0.337	-4.194*	4.238**	-3.362	-2.186	-2.904	10.260**
France	-2.031	-2.408	2.305	-3.489*	5.829*	-3.368	-3.607	-4.480*	6.530
Germany	-0.638	-1.400	0.947	-2.063**	2.984	-1.109	-1.698	-2.242	0.234
Greece	-2.791	-1.537	-0.594	-2.815**	2.022	-1.653	0.564	-2.456	5.177
Hungary	-0.703	-0.703	-1.276	-0.369	0.671	-0.369	-0.703	-3.787***	0.667
Ireland	-1.045	-2.441	-0.728	-1.774	2.101	-2.031	-2.141	-3.200	12.183*
Italy	-0.042	-1.198	0.802	-1.673	1.264	0.073	-1.633	-2.666	5.094
Japan	-1.073	-2.888	-0.936	-2.083**	1.044	-3.622***	-1.521	-3.206	7.771***
Korea	-1.736	-2.216	-0.387	-3.344*	2.336	-2.376	-2.379	-2.840	3.902
Luxembourg	0.714	-2.560	0.606	-1.042	0.494	-2.191	-0.073	-2.914	7.159***
Mexico	-1.155	-1.819	0.436	-3.592*	2.666	-3.083	-0.888	-2.921	7.886***
Netherlands	-1.647	-2.155	0.780	-2.801**	4.495**	-2.144	-3.209	-3.539	9.299**
Norway	-0.498	-2.321	0.514	-1.219	0.450	-1.529	-0.842	-1.973	0.302
Poland	-1.312	-1.312	-0.076	-2.014	2.078	-2.014	-1.905	-1.960	1.410
Portugal	0.416	-2.747	-0.982	-2.017***	0.760	-1.896	-0.849	-3.224	4.553
Spain	-1.198	-2.134	-1.377	-2.098***	3.211***	-2.077	-2.322	-3.686**	3.831
Sweden	-0.904	-2.900	-2.478	-2.032***	2.771	-3.059	-1.979	-3.255	3.884
UK	-1.558	-2.415	0.493	-1.982***	2.765	-2.970	-3.265	-3.662***	3.997
USA	-2.554	-2.592	0.310	-3.446	3.726***	-2.696	-3.078	-3.712***	9.695**

Note 1: ADF %10, %5 and %1 significance level 3.09, 3.39, 3.74. LNV %10, %5 and %1 significance level -3.85, -4.16, -4.76. Sollis %10, %5 and %1 significance level 3.50, 4.30, 6.07. KSS %10, %5 and %1 significance level -1.92, -2.22, -2.82. EG %10, %5 and %1 significance level 3.10, 3.82, 5.53. LNV-KSS %10, %5 and %1 significance level -3.50, -3.81, -4.42. EST test %10, %5 and %1 significance level -3.98, -4.27, -4.81. FFFFF test %10, %5 and %1 significance level for $k < 1$ -3.60, -3.90, -4.50, for $k=1.1$ -3.42, -3.74, -4.39, for $k=1.2$ -3.33, -3.67, -4.31, for $k=1.3$ -3.26, -3.62, -4.29, for $k=1.4$ -3.20, -3.55, -4.22, for $k=1.5$ -3.13, -3.48, -4.14, for $k=1.6$ -3.07, -3.42, -4.10, for $k=1.9$ -2.94, -3.30, -3.99. LNV-Sollis %10, %5 and %1 significance level 7.01, 8.10, 10.53.

Note 2: *, **, *** indicates the %1, %5 and %10 significance levels, respectively.

In order to assess the efficiency of the nonlinear unit root test, we first demonstrate the results of the conventional ADF test. The ADF test, the linear unit root test, could reject the hysteresis hypothesis only for Chile at the %5 significance level. In other words, the unemployment rates in other 24 OECD countries show a hysterical behaviour.

The KSS test which is nonlinear and allows for state dependency and producing symmetric mean reversion displayed the most frequent rejection of the null hypothesis. The test could reject the null hypothesis for 18 OECD countries, in particular Belgium, Canada, Chile, Denmark, Finland, France, Korea, Mexico, USA at the %1 significance level; Germany, Greece, Japan, and Netherlands at the %5 significance level; UK, Sweden, Spain, and Portugal at the %10 significance level. According to the EG test, which is also nonlinear and state dependent, but allows for asymmetry adjustment, the null hypothesis is rejected only in 8 cases, including Spain, USA, Canada, Chile, Finland, Netherlands, Belgium, and France. In contrast to the KSS test, the Sollis test, which produce asymmetric mean reversion, was not able to reject the null hypothesis in any OECD country. Up to now, the symmetric adjustment assumption shows the superior results.

However, the considered tests verify only the possibility of nonlinear adjustment in the unemployment rate series. However, the unemployment rates might experience the structural breaks during the given time period. To see whether allowing for structural breaks alters the results of the unit root tests, we implement three different tests, namely LNV test, EST test and

FFFFF test. The results of the LNV and EST unit root tests suggest that the null hypothesis of a unit root is not rejected. On the other hand, the FFFFF test rejects the null hypothesis in 8 OECD members, namely Australia, Chile, Denmark, France, Hungary, Spain, UK, and USA. The implementation of the unit root tests with structural breaks decreased considerably the number of countries which do not contain a unit root.

In the recent studies, new combined unit root tests, namely the LNV-KSS test and the LNV-Sollis test, were investigated. The tests are nonlinear, allowing state dependency and structural breaks. While the LNV-KSS test could reject the hysteresis hypothesis only for Japan at the %10 significance level, the LNV-Sollis unit root test rejects the hysteresis hypothesis in 9 cases, namely Ireland at %1 significance level; Canada, Denmark, Finland, Netherlands, USA at the %5 significance level; Japan, Luxembourg, Mexico at the %10 significance level. It worse to mention, that the LNV-Sollis test allows an asymmetric nonlinear adjustment towards the attractor in contrast to the LNV-KSS test, which assumes a symmetric adjustment.

Overall, most of the conducted unit root tests verified the hysteresis hypothesis for a majority of the OECD countries considered in our study. However, the KSS test could reject the hysteresis hypothesis in the unemployment rates for 18 OECD countries. According to this result, we suppose that the unemployment rates in those countries follow a nonlinear but globally stationary process.

Conclusion

In this study we observed the validity of the hysteresis hypothesis for 25 OECD countries. We have applied unit root tests allowing for nonlinearity and/or structural breaks. The results recommend that the assumption of more complex dynamics in the unemployment rates results in more frequent rejection of the null hypothesis. In particular, the FFFFF test that allow for structural breaks was able to reject the null hypothesis of unit root for 8 OECD countries, while the LNV-Sollis that allow for both structural break and asymmetric nonlinear adjustment towards the attractor, reject the hysteresis hypothesis in 9 OECD countries. On the other hand, the unit root test that allow only for nonlinear symmetric adjustment, namely the KSS test, shows a more frequent rejection of the null hypothesis. The test demonstrates that the unemployment rates in 18 OECD countries follow a non-hysterical process. Therefore, we suppose that the unemployment rates in those countries follow a nonlinear but globally stationary process.

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**Panel Co-integration between Economic Growth and Non-tax revenue
among Indian States**

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Abstract

Present study tries to explore the stationarity and co-integration between Non-tax revenue and State Domestic Product of twenty major states of Indian federal system under panel data structure. LLC (2002) and IPS (2003) tests of stationarity have been used to detect unit root in the panel series. Kao (1999) test of panel co-integration shows that the SDP and NTR and NTR and SDP for the twenty states for the period under study are co-integrated. The results of the study suggest that state domestic product of the states are causing the non-tax revenue of the states and the non-tax revenue of the states are also causing state domestic product of the states for Indian federal system.

Keywords: State Domestic Production, Non-tax Revenue, Panel Unit Root, Panel Co-integration.

JEL Codes: H2; C23.

1. Introduction

Inadequate revenue sources, uncontrolled growth of current expenditure and failure of central government transfers to grow as fast as the states 'own revenues' have been the major causes of fiscal imbalance at state level. In fact, states' expenditures have been growing which has resulted as shrinking of capital expenditure. Other factors responsible for poor fiscal health of states are up-trend of non-developmental consumption expenditure, high cost of Government employees and implementation of the Sixth Central Pay Commission. From the revenue receipts side, poor tax collections, poorly targeted cost recovery policy and lack of appropriate tax reforms have adversely affected the income buoyancy of the states. Accordingly, states receive revenue from revenue account and capital account. Non-tax revenue of the states comprises interest receipts, receipts from general services like lotteries, fees received for providing various social and economic services etc. But the rates and fees charged for such services are insignificantly low.

Despite marked increase in the absolute level of collection of non-tax revenues, it has remained static at less than one per cent of gross state domestic production. The state government during the period 1995-96 to 2005-06 has surely enhanced their own revenue mobilisation by more than 200 per cent but as their revenue expenditures have gone up simultaneously, their capacity to finance revenue expenditures out of own revenue receipts has declined by more than 16 per cent during the same period. Non-tax revenue grew at a compound annual rate of 7.6 per cent during the period 10 years ending 2009-10 (Economic Survey 2012-13). Further, the states do not have the same ability to finance their growing fiscal imbalance as availed by the Central Government. The states do not have independent power to borrow from open market, nor from Reserve Bank of India because of the regulations of the overdraft facility. Given these constraints on borrowing, the burden of adjusting the imbalances in the state finance has tended to fall mainly on capital account and maintenance of expenditures with adverse implications in terms of infrastructural constraints, declining productivity of state public sector enterprises and ultimately deceleration in the long term growth of economy of the states. These issues have been brought out by several studies like Aiyar and Kurup (1992), Rao and Sen (1993), Srivastava *et al.* (1998), George (2002), Srivastava (2002), Gaur (2002).

It has been found that most of the studies in the field of public finance have concentrated on revenue and expenditure of the state governments by using co-integration and error correction

mechanism like Anderson *et al.* (1986), Artis and Buti (2000), Baghestani and McNown (1994), Chang *et al.* (2002), Garcia and Henin (1999), Hassan and Lincoln (1997), Hondroyiannis and Papapetrou (1996), Joulfaian and Mookerjee (1991), Kollias and Makrydakis (1995), Manage and Marlow (1986), Miller and Russek (1990), Owoye (1995), Gounder *et al.* (2007), Kollias and Paleologous (2007) and Pandey and Dixit (2009).

The relationship between Non-tax revenue (NTR) and State Domestic Production (SDP) has been an important issue of discussion among scholars and economists throughout the world. The existence of nexus in between NTR and SDP can be examined in several ways like growth rate relating to SDP and NTR, proportion of NTR to SDP, several policies relating to accelerate SDP and NTR, etc. So far as inter-state non-tax revenue and state domestic production in India is concerned, limited studies are available. Purohit and Purohit (2009) pointed out the structure of non-tax sources of the states. They presented a detailed analysis of six selected services from social and economic services and concluded that non-tax revenue sources are insignificant sources of revenue in the states' budge. It is important to note that if proper attention is paid in pricing of the government services, non-tax revenue can be the major source of revenue.

Nagpal (2017) in her study analysed the trend of non-tax revenue for the state of Haryana for the period 1980 to 2015. She observed that non-tax revenue has not been able to contribute much to the growing expenditure nor has it been able to keep pace with the other sources of revenue.

Sarma, Pradhan and Bohra (1993) in their study pointed out for many states and many countries in the world, most of the government services are provided free or remain undercharged as these services are regarded 'pure public' goods in character and these are expected to be financed through taxation or borrowings. In case of Rajasthan quarter of the State's own non-tax revenues comprise interest receipts during the period of study i.e. 1974-1993.

Mohanty and Patra (2016) concluded that the per capita non-tax revenue has favorable effect on per capita revenue expenditure of the sub- national governments during the period 2010 to 2014 for the Indian states. But the need is to upgrade the efficiency of revenue collection. As far as policy implications are concerned the government should raise the non-tax revenue by increasing the number of users and the marginal pricing provision of goods. Bagchi (1992)

observed that non-tax revenue has accounted for a small and declining share of total revenue in India.

The review of above literature reveals the fact that no study has been conducted under panel framework for the Indian states for the period 1980-81 to 2011-12. Therefore the present study tries to explore the stationarity and co-integration between Non-tax revenue and State Domestic Product of major states of Indian federal system in time series as well as panel data structure for the period 1980-81 to 2011-12.

Objectives of the Study:

- I. To test the panel stationarity of State Domestic Production and Non-tax revenue of the major states of the Indian federal system in terms of total and growth rate.
- II. To test the panel co-integration in between SDP and NTR for the Indian federal system of twenty major states in terms of total and growth rate.

Hypothesis of the Study:

To test the first hypothesis we are considering the following AR (1) process for N states and T time periods

$$Y_{it} = \rho_i Y_{it-1} + \varepsilon_{it} \quad (1)$$

X_{it} , represents the exogenous variables, including any fixed effect or individual trend, ρ_i is the autoregressive coefficient, and the errors ε_{it} are assumed to be mutually IID. The null and alternative hypotheses are

Ho: If $|\rho_i| < 1$, Y_i is said to be trend-stationarity

Ha: if $|\rho_i| = 1$ then Y_i contains a unit root

To test the second hypothesis between state domestic production (SDP) and non-tax revenue(NTR) for the twenty major Indian states two variables are integrated of the same order say I(1), estimated with OLS the long run equilibrium equation

$$Y_{it} = \beta_0 + \beta_1 X_{it} + \omega_{it} \quad (2)$$

From the above equation save the residuals ω_{it} , as are estimates of the equilibrium error, w_{it} .

For the state domestic production and non-tax revenue for the twenty major Indian states to be co-integrated if the equilibrium errors are stationarity at level. Thus the present equation is to be estimated:

$$\Delta\omega_{it} = \delta\omega_{it-1} + \nu_{it} \quad (3)$$

And finally the hypothesis to be tested:

H₀: $\delta = 0$, for non-stationarity of ω_{it} , i.e. for non-co-integration, if $t_\delta > \tau$

H_a: $\delta < 0$, for stationarity of ω_{it} , i.e. for co-integration, if $t_\delta > \tau$

Econometric Methodology:

Two types of tests can be distinguished based on the panel unit root test (Kappler 2006). First type of test considers a homogeneous alternative (Levin Lin and Chu 2002, Breitung 2000 and Hadri 2000) and the second type of test considers heterogeneous alternative (Im, Pesaran and Shin 2003, Maddala and Wu 1999 and Choi 2001). In the present study, LLC (2002) and IPS (2003) tests of stationarity have been used.

Panel Unit Root Tests: Levin, Lin, and Chu (2002)

Levin Lin and Chu (LLC) consider the following basic ADF specification:

$$\Delta Y_{it} = \alpha Y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta Y_{it-j} + X'_{it} \delta + \varepsilon_{it} \quad (4)$$

Where we assume a common $\alpha = \rho - 1$, but allow the lag order for the difference terms, p_i to vary across cross-sections³.

For a given set of lag orders, we begin by estimating two additional sets of equations, regressing both ΔY_{it} and Y_{it-1} on the lag terms ΔY_{it-j} (for $j = 1, 2, 3, \dots, p_i$) and the exogenous variables X_{it} . The estimated coefficients from these two regressions will be denoted $(\hat{\beta}, \hat{\delta})$ and

³ The hypotheses for the test; $H_0 : \alpha = 0$, i.e., there is a unit root against $H_a : \alpha < 0$, i.e. there is no unit root.

(β, δ) , respectively. We define $\Delta \bar{Y}_{it}$ and $\bar{Y}_{it-1}$ by removing the autocorrelations and deterministic components⁴ and proxies by standardizing both $\Delta \bar{Y}_{it}$ and $\bar{Y}_{it-1}$, dividing by the regression standard error⁵.

An estimate of the coefficient α may be obtained from the pooled proxy equation:

$$\Delta \bar{Y}_{it} = \alpha \bar{Y}_{it-1} + \eta_{it} \quad (5)$$

LLC shows that under the null, a modified t -statistic for the resulting $\hat{\alpha}$ is asymptotically normally distributed

$$t_{\alpha it}^* = \frac{t_{\alpha} - (N T)^{\frac{1}{2}} S_N \hat{\sigma}^{-2} se(\hat{\alpha}) \mu_{mT^{\frac{1}{2}}}}{\sigma_{mT^{\frac{1}{2}}}} \rightarrow N(0,1) \quad (6)$$

Where t_{α} is the standard t -statistic for $\hat{\alpha} = 0$, $\hat{\sigma}^2$ is the estimated variance of the error term η , $se(\hat{\alpha})$ is the standard error of $\hat{\alpha}$, and

$$\bar{T} = T - (\sum_i p_i / N) - 1 \quad (7)$$

The average standard deviation ratio, S_N , is defined as the mean of the ratios of the long-run standard deviation to the innovation standard deviation for each individual. Its estimate is derived using kernel-based techniques. The remaining two terms, $\mu_{mT^{\frac{1}{2}}}$ and $\sigma_{mT^{\frac{1}{2}}}$ are adjustment terms for the mean and standard deviation.

The LLC method⁶ requires a specification of the number of lags used in each cross-section ADF regression, p_i as well as kernel choices used in the computation of S_N .

⁴ $\Delta \bar{Y}_{it} = \Delta Y_{it-1} - \sum_{j=1}^{p_i} \hat{\beta}_{ij} \Delta Y_{it-j} - X'_{it} \delta$ and $\bar{Y}_{it-1} = Y_{it-1} - \sum_{j=1}^{p_i} \hat{\beta}_{ij} \Delta Y_{it-j} - X'_{it} \delta$

⁵ $\Delta \bar{Y}_{it} = (\Delta Y_{it} / s_i)$ and $\bar{Y}_{it-1} = (\bar{Y}_{it-1} / s_i)$

⁶ LLC test depends upon the assumption of independence of across cross-sections. And this test is not applicable if cross-sectional correlation is present.

Panel Unit Root Tests: Im, Pesaran, and Shin (2003)

Im, Pesaran, and Shin (IPS)⁷ begin by specifying a separate ADF regression for each cross section:

$$\Delta Y_{it} = \alpha Y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta Y_{it-j} + X'_{it} \delta + \varepsilon_{it} \quad (8)$$

After estimating the separate ADF regressions⁸, the average of the t -statistics for α_i from the individual ADF regressions, $t_{iT_i}(p_i)$:

$$\bar{t}_{NT} = \left(\sum_{i=1}^N t_{iT_i}(p_i) \right) / N \quad (9)$$

The lag order is always zero ($p_i = 0$ for all i), simulated critical values for $\bar{t}_{NT}$ are provided for different numbers of N and T (equation containing either intercepts, or intercepts and linear trends).

$$W_{\bar{t}_{NT}} = \frac{\sqrt{N} \left\{ \bar{t}_{NT} - N^{-1} \sum_{i=1}^N E(\ell_{iT}^0(p_i)) \right\}}{\sqrt{N^{-1} \sum_{i=1}^N \text{Var}(\ell_{iT}^0(p_i))}} \rightarrow N(0,1) \quad (10)$$

The expressions for the expected mean and variance of the ADF regression t -statistics, $E(\ell_{iT}^0(p_i))$ and $\text{Var}(\ell_{iT}^0(p_i))$, are given by Im, Pesaran, and Shin (IPS) for various values of T and p and differing test equation assumptions⁹.

⁷ IPS suggests an average of the ADF tests when u_{it} are serially correlated with different serial correlation properties across cross-sectional units.

⁸ The null hypothesis and alternative hypothesis may be written as, $H_0 : \alpha_i = 0$, for all i and

$H_a : \begin{cases} \alpha_i = 0 & \text{for } i = 1, 2, 3, \dots, N_1 \\ \alpha_i < 0 & \text{for } i = N + 1, N + 2, N + 3, \dots, N \end{cases}$; (Where i may be reordered as necessary) which may be

interpreted as a non-zero fraction of the individual processes is stationarity.

⁹ Second generations of panel unit root test are surveyed extensively in Omay (2014). Moreover, the new panel unit root test extending IPS are Çorakçı et. al (2017) Omay et. al.(2017), see also these papers.

Panel Co-integration Tests:

A Panel Residual-based test has been performed for panel co-integration (Kao1999) between state domestic production and non-tax revenue for the twenty major Indian states for the period 1980-81 to 2011-12.

Kao (1999) test specifies cross-section specific intercepts and homogeneous coefficients on the first-stage regressors in the panel co-integration. In the bivariate case for panel SDP & panel NTR the model is of the form

$$Y_{it} = \alpha_i + \beta X_{it} + e_{it} \quad (11)$$

Where $Y_{it} = Y_{it-1} + \varepsilon_{it}$ and $X_{it} = X_{it-1} + \xi_{it}$.

The estimated residuals are

$$e_{it} = Y_{it} - \alpha_i - \beta X_{it} \quad (12)$$

i.e. e_{it} Containing a unit root. For this residual we run an ADF test. (Tests can be calculated from the fixed effects residuals)

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + v_{it} \quad (13)$$

The Kao test used the null hypothesis of no co-integration for the panel¹⁰.

Data for the Study:

In the present study data has been taken from Handbook of Statistics on Indian Economy and State Finance for twenty major states; Andhra Pradesh, Assam, Bihar, Gujarat, Haryana, Himachal Pradesh, Jammu & Kashmir, Karnataka, Kerala, Madhya Pradesh, Maharashtra, Manipur, Nagaland, Orissa, Punjab, Rajasthan, Tamil Nadu, Tripura and Uttar Pradesh (Handbook of Statistics on Indian Economy 2012-13). For a proper comparison of stationarity over time for the states, the revised series of SDP should be extended backwards. For this

¹⁰ See also Omay et. al. (2017) for nonlinear panel co-integration extension in VECM form.

purpose, first we computed the price correction factor¹¹. In the present study, panel stationarity has been tested for State Domestic Product (SDP) and Non-tax revenue (NTR) in panel data (in terms of level and growth rate). A Panel Residual-based test has been performed for panel co-integration (Kao1999) between state domestic production and non-tax revenue for the twenty major Indian states.

Results and Discussion:

In the present study, Levin Lin and Chu (LLC 2002) and Im, Pesaran and Shin (IPS 2003) tests of stationarity have been used. In Table 1, LLC panel unit root test presented for the SDP and NTR.

Table 1: Panel Unit Root -LLC test for SDP and NTR.

Model	At Level	Coefficient	t-value	t-star	P>t
State Domestic Product	No constant and no trend	0.079	12.754	12.296	1.00
	With constant	0.082	9.356	15.056	1.00
	With constant and trend	0.000	0.009	6.632	1.00
	At First Difference				
	No constant and no trend	-0.142*	-2.838	-2.779	0.00
	With constant	-0.564	-7.742	2.560	0.99
	With constant and trend	-1.333	-14.654	2.173	0.99
	At Second Difference				
	No constant and no trend	-2.726*	-21.342	-20.752	0.00
	With constant	-2.797	-22.004	3.396	1.00
	With constant and trend	-2.857	-22.291	9.569	1.00
Non-taxRevenue	At Level				
	No constant and no trend	0.157	6.725	6.448	1.00
	With constant	0.137	4.781	15.755	1.00
	With constant and trend	0.010	0.230	14.058	1.00
	At First Difference				
	No constant and no trend	-0.497*	-5.874	-5.752	0.00
	With constant	-0.784	-8.429	8.457	1.00
	With constant and trend	-1.235	-12.177	10.446	1.00
	At Second Difference				
	No constant and no trend	-2.865*	-21.157	-20.620	0.00
	With constant	-2.919	-21.569	13.657	1.00
	With constant and trend	-3.010	-22.254	20.207	1.00

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 12-13. Statistical estimate on the basis of equations 4, 5, 6 and 7 given in section Econometrics Methodology

¹¹ Defined as the ratio of implicit deflator for 2004-05 series to 1999-2000 series for the year 2004-05, 1999-2000 series to the 1993-94 series for the year 1999-2000 similarly for 1993-94 series to 1980-81 series for the year 1994.

* Significant at 1 percent level, STATA11.0 used for LLC test

LLC test suggests presence of unit root at level, first difference and second difference in trend and intercept model. For both the series at no constant and no trend LLC test suggests that the series are stationarity at first difference. Table 2 presents LLC test for panel SDP and NTR in terms of growth rate. The result suggests that both the series are stationarity at no constant and no trend

Table 2: Panel Unit Root -LLC test for growth in SDP and growth in NTR.

Model	At Level	Coefficient	t-value	t-star	P>t
Growth in State Domestic Product	No constant and no trend	-1.079*	-13.814	-13.415	0.00
	With constant	-1.380*	-16.736	-2.638	0.00
	With constant and trend	-1.458	-17.310	0.683	0.75
	At First Difference				
	No constant and no trend	-2.730*	-24.711	-24.010	0.00
	With constant	-2.745	-24.820	0.567	0.71
	With constant and trend	-2.779	-25.080	6.370	1.00
	At Second Difference				
	No constant and no trend	-3.740*	-31.417	-30.555	0.00
	With constant	-3.739	-31.419	6.893	1.00
	With constant and trend	-3.743	-31.456	15.482	1.00
Growth in Non-taxRevenue	At Level				
	No constant and no trend	-1.269*	-15.080	-14.644	0.00
	With constant	-1.373	-15.929	-1.784	0.04
	With constant and trend	-1.432	-16.380	1.213	0.89
	At First Difference				
	No constant and no trend	-2.752*	-24.488	-23.787	0.00
	With constant	-2.753	-24.508	-0.819	0.21
	With constant and trend	-2.760	-24.571	4.681	1.00
	At Second Difference				
	No constant and no trend	-3.665*	-30.227	-29.378	0.00
	With constant	-3.665	-30.212	2.632	1.00
	With constant and trend	-3.664	-30.225	9.748	1.00

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 12-13. Statistical estimate on the basis of equations 4, 5, 6 and 7 given in section Econometrics Methodology

* Significant at 1 percent level, STATA11.0 used for LLC test

Table 3 shows that IPS test on panel SDP and NTR for twenty major states in India. Panel SDP for twenty states in India are stationarity at first difference for intercept and trend while for Panel NTR stationarity at first difference for constant and also for intercept and trend.

Table 3: Panel Unit Root -IPS test for SDP and NTR.

Model	At Level	t-bar	W[t-bar]	P-value
State Domestic Product	With constant	1.330	13.776	1.00
	With constant and trend	-0.541	8.430	1.00
	First Difference			
	With constant	-1.875	-1.957	0.03
	With constant and trend	-3.055*	-4.991	0.00
	Second Difference			
	With constant	-4.553*	-14.872	0.00
	With constant and trend	-4.523*	-12.488	0.00
	At Level			
	With constant	0.366	9.044	1.00
Non-taxRevenue	With constant and trend	-0.378	9.298	1.00
	First Difference			
	With constant	-2.269*	-3.893	0.00
	With constant and trend	-2.772*	-3.479	0.00
	Second Difference			
	With constant	-4.513*	-14.678	0.00
	With constant and trend	-4.558*	-12.666	0.00

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 2012-13. Statistical estimate on the basis of equations 8, 9 and 10 given in section Econometrics Methodology

* Significant at 1 percent level, STATA11.0 used for IPS test

Table 4 presents IPS test for panel SDP and NTR in terms of growth rate. Table 4 shows that SDP and NTR are stationarity at level for constant and also for constant and trend.

From the above we came to know that both the series are stationarity at same order hence panel based co-integration test developed by Kao (1999) can be applied. Kao (1999) test specifies cross-section specific intercepts and homogeneous coefficients on the first-stage regressors in the panel co-integration. In the bivariate case for panel SDP & panel NTR the model is of the form

$$SDP_{it} = \alpha_i + \beta NTR_{it} + e1_{it}$$

$$NTR_{it} = \alpha_i + \beta SDP_{it} + e2_{it}$$

$$SDPG_{it} = \alpha_i + \beta NTRG_{it} + e3_{it}$$

$$NTRG_{it} = \alpha_i + \beta SDPG_{it} + e4_{it}$$

The estimated residuals are

$$e1_{it} = SDP_{it} - \alpha_i - \beta NTR_{it}$$

$$e2_{it} = NTR_{it} - \alpha_i - \beta SDP_{it}$$

$$e3_{it} = SDPG_{it} - \alpha_i - \beta NTRG_{it}$$

$$e4_{it} = NTRG_{it} - \alpha_i - \beta SDPG_{it}$$

Table 4: Panel Unit Root -IPS test for growth in SDP and growth in NTR

Model	At Level	t-bar	W[t-bar]	P-value
Growth in State Domestic Product	With constant	-3.454*	-9.709	0.00
	With constant and trend	-3.487*	-7.298	0.00
	First Difference			
	With constant	-5.217*	-18.067	0.00
	With constant and trend	-5.165*	-15.788	0.00
	Second Difference			
	With constant	-6.605*	-24.741	0.00
	With constant and trend	-6.462*	-22.464	0.00
	At Level			
	With constant	-3.408*	-9.484	0.00
Growth in Non-taxRevenue	With constant and trend	-3.432*	-7.002	0.00
	First Difference			
	With constant	-5.103*	-17.519	0.00
	With constant and trend	-5.002*	-14.954	0.00
	Second Difference			
	With constant	-6.213*	-22.854	0.00
	With constant and trend	-6.072*	-20.457	0.00

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 2012-13. Statistical estimate on the basis of equations 8, 9 and 10 given in section Econometrics Methodology

* Significant at 1 percent level, STATA11.0 used for IPS test

For this residual we run Dickey Fuller test. The Kao test used the null hypothesis of no co-integration for the panel. The Kao test for panel data of SDP and NTR rejects the null hypothesis of no co-integration.

Our results so far indicate that there is a bi-directional relationship between NTR and states economic growth for the panel data. This means that changes in Non-tax revenue affect State Domestic Product, i.e. efforts made in collection of non-tax revenue of the states may provide returns in terms of higher State Domestic Product. The Kao test for panel data also suggests that State Domestic Product affects the Non-tax revenue for the selected Indian States.

Table 5 : Kao test for panel co-integration: Regression results for SDP and NTR

Model	Constant	Variable	R square	ADJ_ R square	R. MSE	F statistics
$SDP_{it} = \alpha_i + \beta NTR_{it} + e1_{it}$	37543.92 (11.04)	0.135 (22.22)	0.4362	0.4353	68549	493.66
$NTR_{it} = \alpha_i + \beta SDP_{it} + e2_{it}$	69364.57 (3.87)	3.224 (22.22)	0.4362	0.4353	330000	493.66
$SDPG_{it} = \alpha_i + \beta NTRG_{it} + e3_{it}$	5.77 (19.35)	0.001 (0.29)	0.0001	-0.0015	7.0246	0.08
$NTRG_{it} = \alpha_i + \beta SDPG_{it} + e4_{it}$	16.70 (6.43)	0.083 (0.29)	0.0001	-0.0015	49.867	0.08

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 2012-13.; Estimated on the basis of Equation 11.; Software used STATA 11.0

Table 6 : Kao test for panel co-integration: Residual Based

DF Test for Residual	τ - statistics	1% critical value
$e1_{it} = SDP_{it} - \alpha_i - \beta NTR_{it}$	-7.295*	-3.430
$e2_{it} = NTR_{it} - \alpha_i - \beta SDP_{it}$	-10.024*	-3.430
$e3_{it} = SDPG_{it} - \alpha_i - \beta NTRG_{it}$	-31.582*	-3.430
$e4_{it} = NTRG_{it} - \alpha_i - \beta SDPG_{it}$	-28.527*	-3.430

Note: Data has been taken from RBI Handbook of Statistics on Indian Economy 2012-13 and State Finance 2012-13.; Residual calculated on the basis of equation 12.; Software used STATA 11.0

* Significant at 1 percent level.

Conclusion

The present study has focused on the nexus of the non-tax revenue and state domestic product of twenty major Indian states. The study is based on the panel data and panel unit root test has been applied on SDP and NTR in the present study. The LLC and IPS tests show stationarity of the series at first difference for trend and intercept. Kao test of panel co-integration shows that the SDP and NTR and NTR and SDP for the twenty states for the period under study are co-integrated. The results of the study suggest that state domestic product of the states are causing the non-tax revenue of the states and the non-tax revenue of the states are also causing state domestic product of the states for Indian federal system.

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Business Cycle Analysis for the Midland-Odessa Petroplex

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Abstract

A business cycle coincident index (BCI) for the Midland-Odessa region is developed using a dynamic single factor modeling methodology. The model assumes that co-movements of metropolitan economic indicators have a common element that can be summarized as a single underlying and unobservable variable known as the “state of the economy.” The model utilizes a Kalman filter smoothing approach which smooths the index across time and across indicators, resulting in index movements that are less pronounced during expansions and recessions. Indicator series used to estimate the Midland-Odessa BCI are: employment, the unemployment rate, real retail sales, and total real wages. The estimated BCI exhibits movements that are correlated with national economic contractions and expansions, movements in oil prices, and an existing Midland-Odessa business cycle index.

Keywords: Regional Business Cycles; Kalman Filter; Coincident Indicators

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1. Introduction

Knowledge regarding current conditions of the Midland-Odessa Petroplex economy is difficult to acquire. Regional stakeholders, such as businesses, local government agencies, and consumers will all benefit from more reliable local economic information. While there are many useful macroeconomic variables that can provide insight to national economic conditions, those data do not always reflect local business conditions (Phillips, 2005). One tool that can be used to gauge metropolitan economic health is a business cycle index (BCI). A BCI is a weighted combination of selected economic variables that can provide insights to the current state of the economy (Phillips and Cañas, 2008).

While Midland and Odessa are two separate metropolitan statistical areas (MSAs), strong economic ties and close geographic proximity intertwine the two MSAs as one metropolitan economy. The Midland-Odessa metropolitan economy is part of the larger Texas Permian-Basin region that contains 20 counties. The Permian Basin economy is driven by the oil and gas industry which sustains over 44,000 jobs. It annually generates \$113.6 billion in economic output (Ewing et al., 2014). Midland-Odessa holds roughly half of the Permian Basin population and is the most densely populated area of the region.

Not surprisingly, Midland-Odessa economic activity is strongly influenced by crude oil price movements. The region experienced rapid expansion during the 1970s and early 1980s due to high oil prices, but lost many of those gains when prices plummeted in 1986 (EIA, 2002). The Petroplex economy was also damaged by the economic downturn and financial market crisis in 2008. The recovery spurred by oil price increases that began in 2009 allowed employment to increase by approximately 50 thousand jobs between 2010 and late 2014. During this same period, the unemployment rate hit a low of 2.8 percent in late 2014 (FRBD, 2017). A significant decline in oil prices after July 2014 caused unemployment to increase throughout the ensuing months.

Broad based measures of the Midland-Odessa economy are relatively scarce. The objective of this study is to develop a business cycle index (BCI) for this metropolitan economy. A dynamic single factor methodology is used to develop the BCI. This method has been successfully employed to estimate business cycle indices for other regions and seems to be well suited toward doing so for the Petroplex, also (Arias et al., 2016; Fullerton and Subia, 2017).

Subsequent sections are as follows. Section two provides a brief review of the evolution of BCI estimation methodologies as well as the more recent Stock and Watson (1991) approach. The third section describes the theoretical model and methodology used. Section four discusses the data employed and the empirical results obtained. Section five summarizes the study and offers concluding remarks.

2. Literature Review

Rudimentary business cycle indexes date from the early 1800s, with interest in the topic intensifying during the 1920s and beyond (Kitchin, 1923). Two key elements were eventually identified to play central roles in the definition of a business cycle: co-movements among the individual indicators and the differentiation between expansions and contractions (Burns and Mitchell, 1946; Diebold and Rudebusch, 1996). The Burns and Mitchell (1946) index was maintained by the U.S. Department of Commerce (USDC) until 1995 (Phillips, 2005).

The Conference Board (CB) now maintains the U.S. coincident index which employs four indicators: employees on nonagricultural payrolls; personal income less transfer payments measured in 1996 dollars; industrial production; manufacturing, wholesale, and retail trade sales in 1996 dollars (CB, 2012). The index is calculated by computing the month-to-month changes of each indicator and then adjusting these monthly contributions to equalize the volatility of each indicator. The “component factor,” used to equalize the volatility of the monthly contribution, is inversely related to the standard deviation of the month-to-month changes in each component (CB, 2012).

Stock and Watson (1991) employ an econometric approach to index estimation that provides a statistical rationale for the estimation of indicator indices. The model developed is a parametric version of a single index model proposed by Sargent and Sims (1977). The model assumes that the co-movements of indicators have a common element that can be summarized as a single underlying and unobservable variable, the “state of the economy” (Stock and Watson, 1991). It utilizes a Kalman filter smoothing approach which smooths the index across time and across indicators which results in index movements that are less pronounced during expansions and recessions (Phillips and Cañas, 2008). The single factor model approach produces an index that is highly correlated with the CB coincident index.

Clayton-Matthews and Stock (1998/1999) use the single factor method to estimate coincident and leading indexes for Massachusetts. Five criteria are used to select the specification for the coincident index model. The estimated index should reflect recent economic history as well as the cyclical history of the regional economy. The estimated filter should assign weights on the indicator series that concentrate around a zero lag and each indicator should make a significant contribution to the output index. The one-step ahead forecast error for each indicator should be uncorrelated with itself, the forecast errors of the other indicators, and the other indicators. The output of the Kalman filter should be smooth to reflect the assumed smooth state of the economy. Characteristics of the index should be robust.

Phillips (2005) applies the Stock-Watson model to estimate a coincident index for Texas. The index is estimated using non-farm employment, quarterly real gross state product (RGSP), and unemployment. The growth rate for the index is equal to the weighted average change of the component series in order to make it comparable to a different Texas coincident index estimated using the CB method. One-step ahead forecast errors are tested to assess whether the white noise components of the error terms of each of the indicators are uncorrelated with themselves, forecast errors of other indicators, and each of the other indicators (Clayton-Matthews and Stock, 1998/1999). The new coincident index passes the white noise specification test, but the CB method index does not. Finally, the dynamic single factor Texas index is also found to reliably predict downturns with relatively few false signals and with good timing characteristics (Neftci, 1982).

It should be noted that the dynamic single factor model is not the only modern method for estimating a business cycle index. Other candidates include the Hamilton (1989) regime switching approach, the Terasvirta and Anderson (1992) smooth transition procedure, and the Kim and Nelson (1998) dynamic single factor with regime switching technique. Each of the alternative methods has been shown to provide useful information and insights with respect cycle movements in various economies. The approach deployed in this study has been successfully applied to the development of coincident business cycle indices for other metropolitan economies in Texas. Given that, it makes a logical candidate for estimating a similar gauge for the Midland-Odessa Petroplex.

Phillips and Cañas (2008) use the dynamic single factor method to estimate a coincident index for four border metropolitan economies in Texas. Establishment employment, unemployment rate, real wage, and real retail sales data are deployed as indicators with the average growth rate of real personal income in each respective area used to re-trend the series. More recently, the same methodology, indicators, and re-trending procedure is used by Fullerton and Subia (2017) to estimate a coincident index for the Lubbock metropolitan economy which is located fairly close to Midland-Odessa. To date, the latter metropolitan economy has not been analyzed using these techniques. Material below develops a dynamic single factor model coincident index for the regionally important Petroplex economy.

3. Theoretical Model

Indicator series used to estimate the Midland-Odessa BCI are: employment, unemployment rate, real retail sales, and total real compensation (BLS, 2017a). Because the methodology produces a stationary index, the coincident index is re-trended using the growth rate of real metropolitan personal income. The average growth rate of the index is re-trended as equal to the average growth in real metropolitan personal income over the course of the sample (Phillips and Cañas, 2008). Metropolitan personal income data are obtained from the Bureau of Economic Analysis (BEA, 2017), retail sales data are obtained from the Texas Comptroller of Public Accounts (TCPA, 2016), wage data are obtained from the Texas Workforce Commission (TWC, 2017), and employment data are obtained from the Federal Reserve Bank of Dallas (FRBD, 2017). Data series used in this sample begin in 1990. Indicator series for the coincident index are chosen to represent broad components of the overall economy: consumption (retail sales), income (wages), plus employment and unemployment (production) (Clayton-Matthews and Stock, 1998/1999). Comparability to indicator series used by the Conference Board for the national coincident index was also considered. Other selection criteria for indicator series includes: timeliness, relatively high frequency, and availability across MSAs.

Employment data produced by the Bureau of Labor Statistics (BLS, 2017b,c) are early benchmarked with data from TWC (2016). The data are seasonally adjusted with a two-step seasonal adjustment procedure (Phillips and Nordlund, 2012). Early benchmarking and two-step seasonal adjustments are completed by FRBD (2017). The unemployment rate is seasonally adjusted using the X-12 procedure (BLS, 2017b, c).

Wage and compensation data are seasonally adjusted using the X-12 procedure and adjusted for inflation to first quarter 2016 dollars. Retail sales data are obtained from the Texas Comptroller of Public Accounts. The Standard Industrial Classification (SIC) system was used to measure retail sales prior to 2002. Beginning in 2002, the North American Industry Classification System (NAICS) is used to measure retail sales. In order to avoid the bias that would otherwise occur from combining the two data sets, data prior to 2002 are converted into NAICS using the 2002 NAICS to 1987 SIC concordance provided by the U.S. Census Bureau (USCB, 2002). The data are then seasonally adjusted using the X-12 procedure and adjusted for inflation to first quarter 2016 dollars. Wage and retail sales data are reported quarterly with a lag of about three quarters. Metropolitan personal income is adjusted for inflation to first quarter 2016 dollars. Those data are reported yearly with a lag of two years.

The basic form of the model is:

$$Y_t = \beta + \gamma(L)\Delta C_t + \mu_t \quad (1)$$

$$D(L)\mu_t = \varepsilon_t \quad (2)$$

$$\phi(L)\Delta C_t = \delta + \eta_t \quad (3)$$

To estimate the coincident index, the Stock and Watson (1991) procedure is used. The employment, wage, and retail sales series are first transformed using natural logarithms and then first differenced. The unemployment rate is reported in percentages and is differenced without transformation. The resulting series are then normalized by subtracting respective means and dividing by respective standard deviations. This steps insures $\beta = 0$ in Equation (1) and $\delta = 0$ in Equation (3).

Thus, $Y_t = \Delta X_t$ are the stationary first difference in natural logs of the indicators, C_t is the log of the unobserved state of the economy, and μ_t is the idiosyncratic component of each indicator. $\phi(L)$ is a scalar lag polynomial and $D(L)$ is a lag polynomial matrix, L denotes the lag operator. $\phi(L)$ and $D(L)$ are assumed to have finite orders p and k , respectively. The disturbance terms μ_t and ε_t are assumed to be serially uncorrelated and this assumption is tested as part of the empirical analysis. It is assumed that all off-diagonal values of the $D(L)$ matrix are equal to zero and that all diagonal values are non-zero numbers, assumptions that are also tested below. This allows the μ_t in different equations to be mutually uncorrelated with each other at all leads and lags.

The scale of the $\gamma(L)$ coefficients is fixed by setting the variance of η to unity, while the timing of the coincident index is fixed by setting all but one of the elements $\gamma(L)$ to zero in one of the indicator equations in Equation (1). One- and two-period lags are included for each indicator so that the variables are allowed to influence the index from zero to two lags (Phillips, 2005). Lags that are not significant at the 5% level are dropped. Significant error terms for each of the univariate equations are determined by examining correlograms and Q-statistics of the residuals estimated for the various lag structures. Autoregressive lags of the coincident index itself may also be added in order to correct for any autocorrelation in the index and reduce month-to-month noisiness.

Equations (1) - (3) form the dynamic single-factor, multiple indicator model (Sargent and Sims, 1977). If the Y_t indicators move in conjunction with the economy, then the common co-movement C_t can be interpreted as the current state of the economy (Stock and Watson, 1991). Equation (1) shows how each of the indicator series contributes to the underlying growth process (Phillips and Cañas, 2008). ΔC is the common co-movements in the growth of the indicators, Y . Equation (2) models the idiosyncratic components of each on the indicator series. The idiosyncratic components, μ , are stationary, mean zero, autoregressive stochastic processes (Clayton-Matthews and Stock, 1998/1999). Equation (3) models the growth of the underlying state of the economy (Phillips and Cañas, 2008). Growth in the state of the economy is modeled as a stationary autoregressive process.

Maximum likelihood estimates of the parameters for Equations (1) - (3) and estimation of the filtered state are obtained by representing Equations (1) - (3) in state space form and using a Kalman filter (Clayton-Matthews and Stock, 1998/1999). As shown in Stock and Watson (1991), the state space equation, also called the transition equation, is formed by combining Equations (2) and (3). Because one objective is to estimate the level of C_t using information up to time t , it is convenient to augment these equations by the identity $C_{t-1} = \Delta C_{t-1} + C_{t-2}$ (Stock and Watson, 1991). The Transition equation is given by,

$$\begin{pmatrix} C_t^* \\ \mu^* \\ C_{t-1} \end{pmatrix} = \begin{pmatrix} \phi^* & 0 & 0 \\ 0 & D^* & 0 \\ Z_c & 0 & 1 \end{pmatrix} \begin{pmatrix} C_{t-1}^* \\ \mu_{t-1}^* \\ C_{t-2} \end{pmatrix} + \begin{pmatrix} Z_c' & 0 \\ 0 & Z_\mu' \\ 0 & 0 \end{pmatrix} \begin{pmatrix} \eta_t \\ \varepsilon_t \end{pmatrix} \quad (4)$$

where $C_t^* = (\Delta C_t \Delta C_{t-1} \dots \Delta C_{t-p+1})'$ $\mu^* = (\mu_t' \mu_{t-1}' \dots \mu_{t-k+1}')'$

$$\begin{aligned} \phi^* &= \begin{pmatrix} \phi_1 \dots \phi_{p-1} & \phi_p \\ I_{p-1} & 0_{(p-1) \times 1} \end{pmatrix} \\ D^* &= \begin{pmatrix} D_1 \dots D_{k-1} & D_k \\ I_{n(k-1)} & 0_{(k-1) \times n} \end{pmatrix} \\ Z_c &= \begin{bmatrix} 1 & 0_{1 \times (p-1)} \end{bmatrix} \\ Z_\mu &= \begin{bmatrix} 1 & 0_{1 \times (p-1)} \end{bmatrix} \end{aligned}$$

The asterisk indicates that the variable represents a vector or matrix of variables and is used for notational brevity.

The measurement equation is obtained by writing Equation (1) as a linear combination of the state vector:

$$Y_t = \beta + \begin{bmatrix} YZ_c & Z_\mu & 0_{n+1} \end{bmatrix} \begin{bmatrix} C_t^* \\ \mu_t^* \\ C_{t-1} \end{bmatrix} \quad (5)$$

Equation (6) and (7) can be written as

$$\alpha_t = \mu_\alpha + T_t \alpha_{t-1} + R \zeta_t \quad (6)$$

$$Y_t = \beta + Z \alpha_t + \xi_t \quad (7)$$

Where $\alpha_t = (C_t^* \mu_t^* C_{t-1})'$, $\xi_t = (\eta_t \varepsilon_t)'$

The Kalman filter is then applied to the state space representation of the model. Let $\alpha_{t|\tau}$ denote the estimate of α_t based on $(y_1, \dots, y_\tau)$, let $E[\xi_t \xi_t'] = H$, $E[\zeta_t \zeta_t'] = \Sigma$ and $P_{t|\tau} = E[(\alpha_{t|\tau} - \alpha_t)(\alpha_{t|\tau} - \alpha_t)']$. Given this notation, the prediction equations for the Kalman filter are:

$$\alpha_{t|t-1} = \mu_\alpha + T_t \alpha_{t-1|t-1} \quad (8)$$

$$P_{t|t-1} = T_t P_{t-1|t-1} T_t' + R \Sigma R' \quad (9)$$

The forecast of Y_t at time $t-1$ is $Y_{t|t-1} = \beta + Z \alpha_{t|t-1}$ and the updating equations for the filters are:

$$\alpha_{t|t} = \alpha_{t|t-1} + P_{t|t-1} Z' F_t^{-1} v_t \quad (10)$$

$$P_{t|t} = P_{t|t-1} - P_{t|t-1} Z' F_t^{-1} Z P_{t|t-1} \quad (11)$$

where $F_t = E[v_t v_t'] = Z P_{t|t-1} Z' + H$ and $v_t = Y_t - Y_{t|t-1}$.

This estimation procedure yields three versions of the state vector: $\Delta C_{t|t-1}$, the prediction estimates; $\Delta C_{t|t}$, the filtered estimates; and $\Delta C_{t|T}$, the smoothed estimates. The prediction estimates, $\Delta C_{t|t-1}$, use information from the prior period to estimate the state of each period (Clayton-Mathews and Stock, 1998/1999). These estimates are used to form the one-step ahead prediction errors, $\hat{\varepsilon}_{t|t-1} = \Delta X_t - \Delta X_{t|t-1}$ which are based on the initial parameter estimates (Clayton-Mathews and Stock, 1998/1999). These prediction errors are the fitted residuals from Equations (1) and (2), where the estimates for $\Delta C_{t|t-1}$ are used in place of the unobserved ΔC . The one-step ahead prediction errors are later used to test the validity of the assumption that the error terms of the component series are mutually uncorrelated with each other at all leads and lags (Clayton-Mathews and Stock, 1998/1999).

As shown in Equation (3), the Kalman filter models each of the indicator series as left-hand-side variables with the unobserved coincident index on the right hand side. Given this structure, quarterly variables are modeled as a function of current and previous values of the monthly latent series. This allows quarterly data to enter the equations with monthly data. This can be demonstrated by:

$$\Delta X_t = \Upsilon(L)\Omega(L)\Delta C_t + \mu_t \quad (12)$$

where $\Omega(L) = 1 + 2L + 3L^2 + 2L^3 + L^4$ and $\Delta X_t = X_t - X_{t-3}$, with t representing months.

Because the methodology produces indexes that are stationary with unit variances, it is necessary to make two adjustments. First, the variance of the growth rate of the index is scaled to the average variance of the growth rates in the component series. Second, the average growth rate of the index is re-trended to equal the average growth in real metropolitan personal income over the course of the sample (Phillips and Cañas, 2008).

The index is estimated using the Dynamic Single-Factor Model Software (DSFM) package (Clayton-Matthews, 2005). The structure of the model, estimation, and transformation from estimated state of the economy to metropolitan coincident index allows extracting the relevant business cycle signal from potentially noisy data. The following section provides an analysis of the estimated coincident index for the Midland–Odessa Petroplex economy.

4. Empirical Analysis

Table 1 contains descriptions of the indicator series used to estimate the coincident index. Table 2 reports summary statistics for each indicator series during the sample period. The employment indicator reaches a maximum of approximately 179 thousand and has an average of about 120 thousand. The series follows an upward trend and exhibits cyclical movements. The unemployment indicator exhibits pronounced cyclical movements and a slight downward trend. It has a maximum of about 9.6 percent and an average of 5.35 percent. Real retail sales has a maximum of \$2,055 million and an average of \$1,260 million. It follows an upward trend and displays notable cyclical movements. Real total wages has an upward trend, but displays relatively muted cyclical movements in comparison to the other indicators.

Table 1: Variables, Definitions, and Units of Measure

Indicator Series	Definition	Units	Source
Employment	Monthly MSA total non-farm employment, early benchmarked using preliminary releases of the QWEC from the TWC, and two-step seasonally adjusted.	Thousands	Bureau of Labor Statistics; Federal Reserve Bank of Dallas
Unemployment	Monthly MSA total labor force currently unemployed and seeking employment, seasonally adjusted	Percent	Bureau of Labor Statistics; Federal Reserve Bank of Dallas
Real Retail Sales	Quarterly MSA retail sales as defined by NAICS, seasonally adjusted, 2016Q1 dollars	Millions	Texas Comptroller of Public Accounts
Real Total Wages	Quarterly MSA total wages and salaries for all industries, seasonally adjusted and in Q1 2016 dollars. Included in wages are paid leave, bonuses, stock options, tips, and cash value of meals and lodging.	Millions	Texas Workforce Commission

Table 2: Indicator Series Summary Statistics

	Employment	Unemployment	Real Retail Sales	Real Total Wages
Mean	119.77	5.35	\$1,260.66	\$1,393.17
Median	109.61	5.17	\$1,145.42	\$1,107.07
Maximum	178.93	9.59	\$2,054.80	\$2,810.70
Minimum	88.78	2.67	\$905.66	\$894.61
Standard Deviation	24.76	1.61	\$299.91	\$542.41
Skewness	0.82	0.58	1.16	1.11
Kurtosis	2.48	2.67	3.28	2.96

Notes: Values rounded to two decimal places. Sample period: employment and unemployment 1990M01 – 2016M08. Sample period: real retail sales and real total wages 1990Q1 – 2016Q1.

All four indicator series are positively skewed (Table 2). Real total wages, real retail sales, and employment have overall upward trends, while unemployment has a downward trend. All four series exhibit cyclical patterns caused by shocks to the regional economy. These shocks cause rapid short term changes in the indicator series. The steady trends of the indicators are punctuated with small periods of high growth, resulting in positive skewness in each series.

Table 3 reports the estimated values of the coefficients for the common state, the lags of the autoregressive error term equations, and autoregressive lags of the index. The coefficients of the common state exhibit the hypothesized signs and are all statistically significant at the 1-percent level. The common state enters the employment measurement equation contemporaneously and with lags of one and two months. The common state enters the measurement equations for unemployment and real total wages contemporaneously, only. The common state enters the real retail measurement equations contemporaneously and with a one-period lag.

One measure of smoothness of the coincident index is the sum of the autoregressive coefficients of the index itself. Shocks are more persistent and the cycle is smoother when the sum is closer to one, while remaining less than one (Phillips, 2005). The index estimation includes one first-order autoregressive term with a value of approximately 0.823. The autoregressive term of the index is statistically significant at the 1-percent level.

Table 3: Estimated Coefficients

Estimated coefficients of the common state C_t				
Measurement Equation	Variable	Coefficient	Asymptotic std. error	t-statistic
Employment	C_t	0.552	0.063	8.700***
	C_{t-1}	-0.516	0.084	-6.092***
	C_{t-2}	0.359	0.059	6.039***
Unemployment Rate	C_t	-0.315	0.046	-6.805***
Real Retail Sales	C_t	0.112	0.023	4.832***
	C_{t-1}	-0.074	0.022	-3.273***
Real Total Wages	C_t	0.058	0.006	8.700***
Estimated Coefficients for lags in the autoregressive equations for the error terms				
Measurement Equation	Lag	Coefficient	Asymptotic std. error	t-statistic
Unemployment	-1	-0.056	0.063	-0.886
	-2	0.172	0.058	2.949***
	-3	0.221	0.062	3.536***
	-4	0.136	0.063	2.143**
Real Wages	-1	-0.653	0.156	-4.188***
	-2	0.143	0.198	0.725
	-3	0.346	0.159	2.171**
Estimated Coefficients for the autoregressive equation for the common factor				
Measurement Equation	Lag	Coefficient	Asymptotic std. error	t-statistic
Lag	-1	0.823	0.044	18.582***

Notes: Sample period: employment and unemployment 1990M01 – 2016M08. Sample period: real retail sales and real total wages 1990Q1 – 2016Q1.

The results of the diagnostics test on the one-step ahead prediction errors are shown in Table 4. Although both the unemployment and real wages series have statistically significant autoregressive coefficients, as shown in Table 3, the results of the diagnostics test generally indicate that the error terms of the component series are mutually uncorrelated at all leads and lags. One of the F-statistics is statistically significant at the 5-percent level and none of the other 31 coefficients are statistically significant. These results verify that idiosyncratic components of the indicators are modeled correctly and confirm the assumption that the co-movements of the component series stem from a single underlying variable.

Table 4: F-Statistics for 6-Lag Specification Test

Dependent Variables	eEmployment	eUnemployment	eRetail	eWages
eEmployment	0.581	1.174	0.416	0.314
eUnemployment	0.882	0.342	0.181	0.364
eRetail	1.352	1.507	1.328	0.564
eWages	1.679	0.880	0.717	0.810
Employment	0.507	1.637	0.463	0.171
Unemployment	1.083	0.629	0.276	0.340
Real Retail Sales	1.052	1.485	1.350	0.547
Real Total Wages	0.649	2.237**	0.902	0.598

Notes: Dependent Variable is One-Step Ahead forecast error. * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$. H_0 : Coefficients are jointly zero. Failure to reject the null hypothesis indicates that the idiosyncratic components of indicator series are uncorrelated and implies that the presence of only one single underlying factor is a reasonable assumption. Sample period: employment and unemployment 1990M01 – 2016M08. Sample period: real retail sales and real total wages 1990Q1 – 2016Q1.

Table 5 reports the cumulative dynamic multipliers and the estimated weights for each of the indicators. The multipliers are the average growth rate of each of the indicator series and the weights are the share that each average growth rate contributes to the common co-movement growth rate, ΔC (Murphy, 2005). Employment and real wages have the greatest weights, 53.24% and 20.64% respectively. Unemployment has a weight of 13.65% and real retail sales has the smallest weight of 12.46%.

Table 5: Cumulative Dynamic Multipliers

Indicator	Multiplier	Share*
Employment	1.128	53.243
Unemployment	-0.289	13.651
Real Retail Sales	0.264	12.460
Real Total Wages	0.437	20.644

Notes: Values rounded to two decimal places. Sample period: employment and unemployment 1990M01 – 2016M08. Sample period: real retail sales and real total wages 1990Q1 – 2016Q1.

The average growth rate of the computed index is approximately 0.24% and has been adjusted to match the 0.21% average growth rate of Midland-Odessa annual real total personal income for the 1990 to 2014 sample period. This adjustment is made because the methodology produces a stationary index which shows changes in the business cycle but not the trends and volatility of the region (Phillips and Cañas, 2008).

Figure 1 plots the re-trended index of economic activity for the combined Midland-Odessa metropolitan economy. Although the data used to estimate the BCI begin in January or Q1 1990, the estimated BCI begins in January 1991 because employment and real retail sales include lags of the common state, C_t . This first available BCI value is used as the base period. The graph highlights six recessionary periods. Recessions for Midland-Odessa are defined as periods of at least a 0.5% decline in the coincident index with at least three months between the peak and trough (Crone, 2006). Three of the Petroplex downturns coincide with national recessions, but with later start dates (peaks) than those of the macroeconomic downturns.

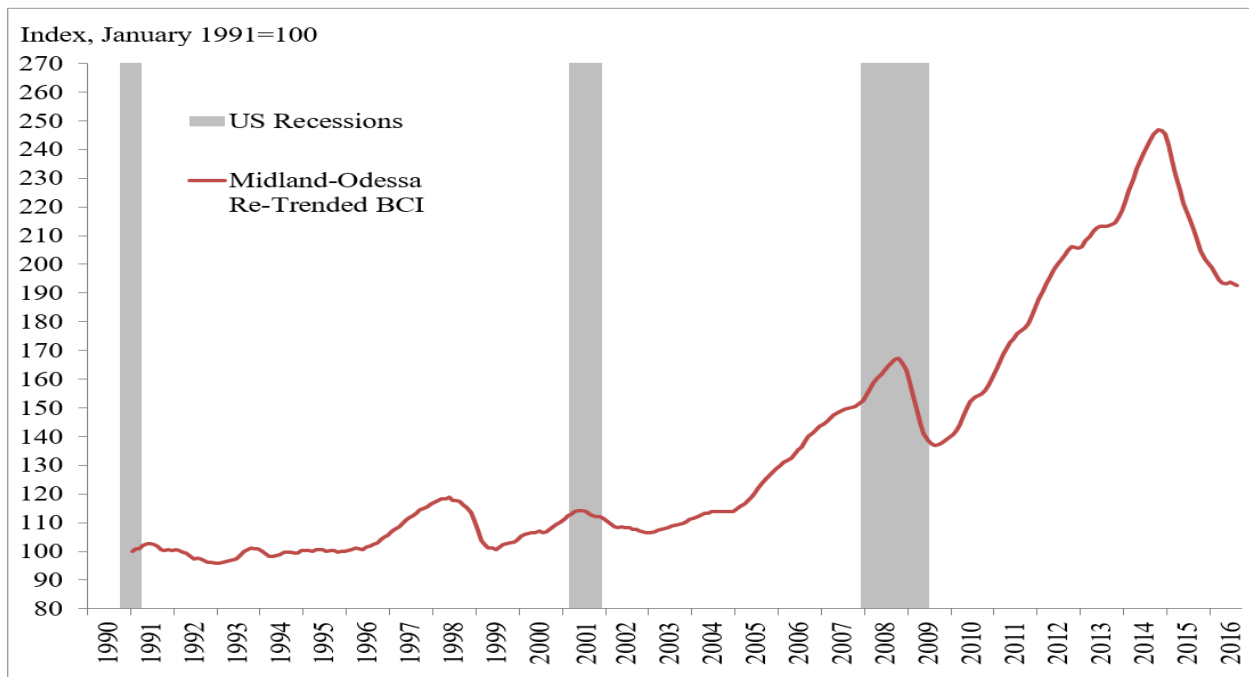


Figure 1: Petroplex Re-Trended BCI

The first local recession that is independent of a national contraction occurred between October 1993 and March 1994. The second independent contraction occurred between May 1998 and June 1999. The third independent recession begins in October 2014 and, because the index is still in decline for this most recent estimation period, a trough is not yet established for it.

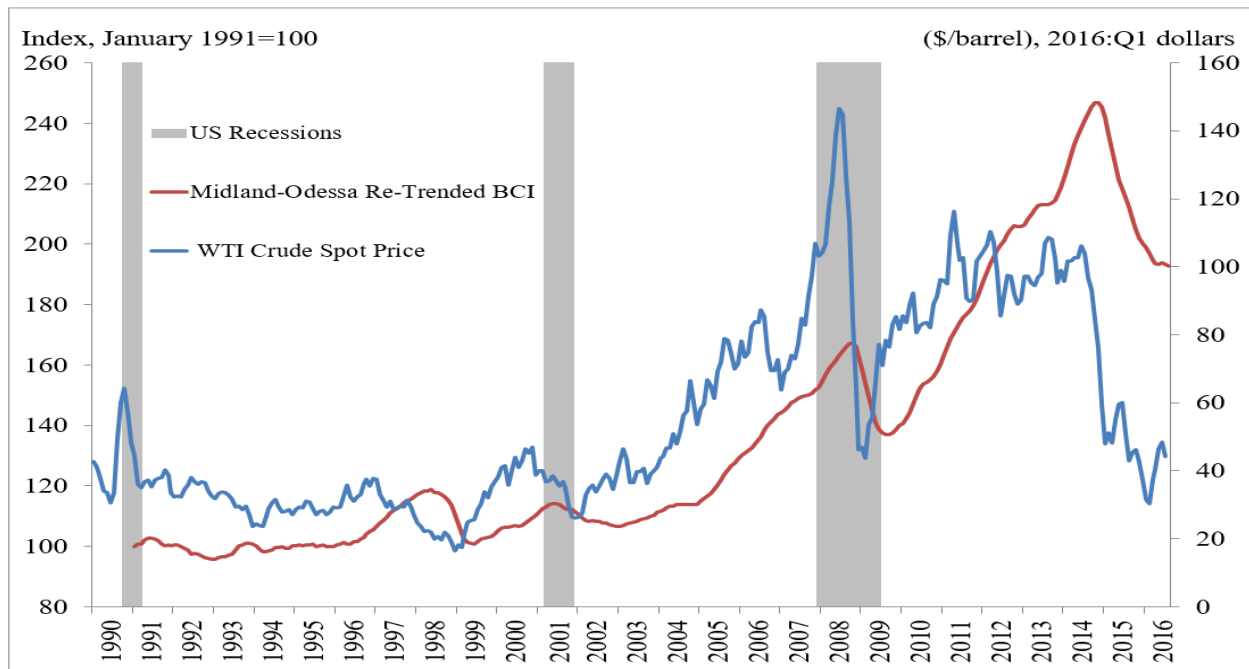


Figure 2: Petroplex Re-Trended BCI and Monthly WTI Oil Price

Figure 2 shows that all three of the Petroplex local recessions are associated with sharp declines in oil prices. Prices declined by 34% between June 1992 and March 1994, by 56% between December 2006 and December 2008, and by 59% between June of 2014 and July 2016.

Notable expansionary periods in the Permian Basin economy are also closely tied to increases in oil prices. The correlation coefficient between the contemporaneous lags of the re-trended BCI and the West Texas Intermediate oil price series during the sample period is approximately 0.668. When oil prices increased by 40% between July 1995 and December 1996, the coincident index increased by 13% between December 1996 and May 1998. Oil prices more than doubled between March 1999 and January 2001 and the index increased by 14% between June 1999 and May 2001. When oil prices more than doubled between late 2003 and mid-2006, the BCI increased by 47% between November 2004 and October 2008. Oil prices also doubled between early 2007 and mid-2008. That development likely contributed to the delayed materialization of the 2008 to 2009 national recession in Midland-Odessa.

Figure 3 shows both the index estimated in this study as well as the Midland-Odessa Regional Economic index. The latter index is published monthly by the Midland Development Corporation (MDC) and is estimated by Ingham Economic Reporting (MDC, 2016). The MDC index tracks the business cycles of the combined Midland-Odessa metropolitan area using component series that include taxable retail sales, spending on automobiles, hotel/motel tax receipts, airline passenger volumes, value of all building permits, housing sales dollar volumes, average home sales prices, unemployment, and employment (Ingham, 2013). The methodology behind the estimation of the MDC index is not disclosed, but one observable difference between its estimation and the index estimated in this study is the greater number of indicators used in the estimation of the MDC index. As seen in Figure 3, the two indexes follow very similar patterns until November 2010 when the index estimated in this study begin to exhibit higher levels of growth. In general, the movements in the two indexes are closely correlated, but the Petroplex BCI tends to mark peaks and troughs earlier than the MDC series. The MDC appears to track business cycles as more of a lagging index relative to the coincident BCI developed in this study.

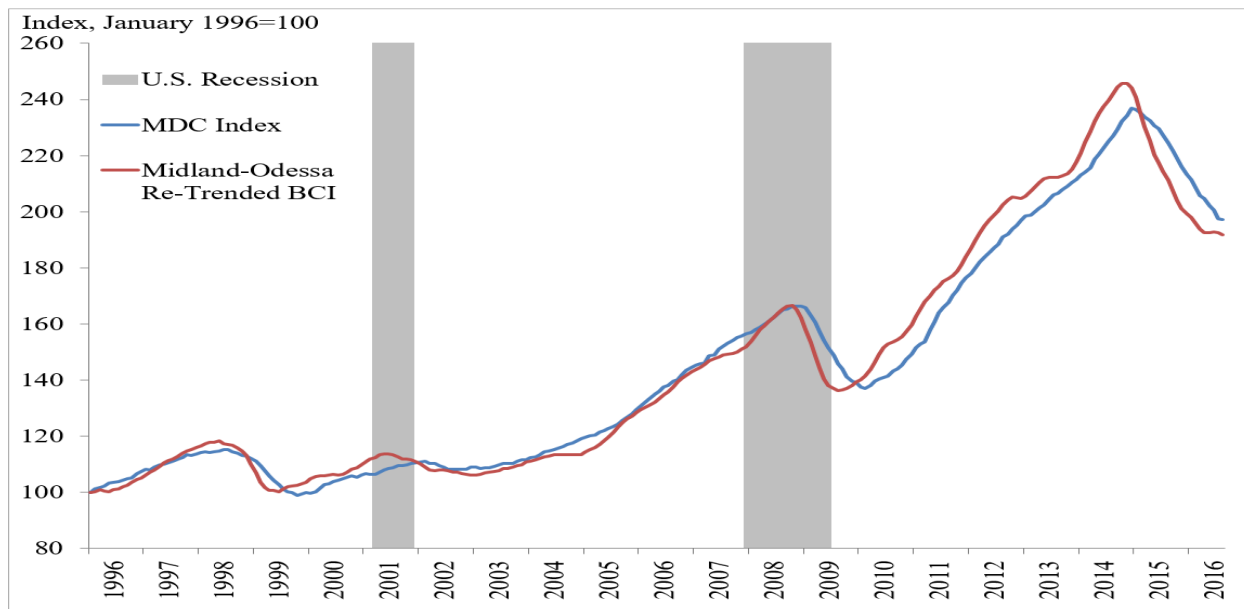


Figure 3: Petroplex BCI and MDC BCI

The Midland-Odessa Petroplex BCI offers a new tool for understanding the state of the regional economy. The timeliness and availability of the indicator series used to estimate the index allows for a more current estimation of economic conditions. It is clear that the dynamics of this economy often deviate from those of the national economy. The Petroplex BCI provides a broad-based means for better tracking business conditions in this unique metropolitan economy.

Conclusion

A dynamic single-factor modeling approach is used to estimate a coincident index for the Midland-Odessa Petroplex economy (Stock and Watson, 1991; Clayton-Matthews, 2005). The methodology assumes that the co-movements of key economic indicators have a single underlying, unobservable factor. Through an econometric approach, this factor is extracted from indicator variables and used to calculate an index that represents current economic conditions. The model also utilizes the Kalman filter smoothing approach, a recursive process that estimates mean square error estimates, facilitates the use of mixed-frequency data, and results in a smoother index with less pronounced recessions and expansions. Indicator series used in this study are: employment, unemployment, inflation adjusted retail sales, and real total compensation. The data used in this study cover the period from January 1990 through August 2016.

The calculated index for Midland-Odessa is generally smooth and its movements are consistent with historical economic conditions of the region. Three of the recessionary periods in Midland-Odessa are associated with national economic contractions. Two additional recessions for this metropolitan economy follow the movement of the West Texas Intermediate crude oil price. Given that the petroleum industry is a well-known driver of the Midland-Odessa economy, the strong correlation between the coincident index and oil prices is as hypothesized.

The metropolitan business cycle index for Midland-Odessa is designed to reliably reflect current economic conditions. The approach employed is one that has been applied to other metropolitan economies in Texas (Phillips and Cañas, 2008). The new index thus provides an additional tool for monitoring Midland-Odessa economic conditions that is directly comparable to the BCI utilized for the other economic regions in Texas and has relatively limited data requirements.

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