Jurnal
EKONOMI
PEMBANGUNAN
Kajian Ekonomi Negara Berkembang
Hal: 201 – 212

UNIT ROOT TEST WITH ONE ENDOGENOUS STRUCTURAL BREAK EVIDENCE FROM INDONESIAN TIME SERIES DATA

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Abstract

This paper examines the robustness of the ADF (Augmented Dickey-Fuller) unit root test to the presence of one structural break. The ADF test results show one variables out of six to be stationary. To check their robustness, two separate additive outlier (AO) models are employed: one allowing for one endogenously-determined break in the intercept and the other in the trend. These two tests can not reject the unit root null hypothesis for all the variables. However, when an innovational outlier (IO) model, that allows for one endogenously-determined break is estimated, the null hypothesis can be rejected for 3 more series. The estimated break dates mostly correspond to the 1998 financial crisis in Indonesia.

Keywords: *unit root; stationarity; structural break, additive & innovational outlier*
JEL classification: C1; C22

INTRODUCTION

In a widely cited study, Nelson and Plosser (1982) show that the null hypothesis of a unit root could not be rejected for eleven out of fourteen macroeconomic series of the US economy. Numerous subsequent studies confirm their study results (Phillips and Xiao, 1998). Accordingly, most macroeconomic series are characterised by differenced stochastic process instead of trend stationary process.

However, it is also acknowledged that macroeconomic series may experience various breaks in their realisations due to shocks arising from crises or radical policy changes[1]. Perron (1989) argues that if breaks are present in the true data generating process, but are not incorporated in the econometric model specification, the conclusion is likely biased towards non-rejecting the null hypothesis of a unit root. After allowing for one exogenously determined break, Perron (1989) provides results that largely reverse Nelson and Plosser (1982) conclusion. That is, most of macroeconomic series under their study better characterised as a stationary process. Nonetheless, Zivot and Andrews (1992) who amend the weakness of Perron (1989) framework by determining the break date endogenously results in conclusions that largely reverse Perron's conclusions.

This paper presents the results of unit root test on six main macroeconomic series of the Indonesia economy. Since most of these series, through visual inspection, show at least one break, this study applies unit root tests that allow for one break at unknown time. Before doing so, however, the traditional unit root test without structural break is implemented for the sake of comparison. The paper is organised as follows. Section II briefly reviews the literature on the unit root tests with one structural break. Section III outlines an empirical framework of unit root test with one structural break. Section IV implements the unit root test that allows for two structural breaks at unknown times and presents its results. Finally, Section V summarises the findings.

UNIT ROOT TEST WITH ONE STRUCTURAL BREAK: A BRIEF REVIEW

Pioneering unit root tests that allow for one structural break, Perron (1989) postulates that the break is exogenous to the realisation of the underlying data-generating mechanism of the series and hence the break occurs at a known time. He extends the Dickey–Fuller type test by augmenting it with dummy variables that represent structural breaks in the economy[2]. Perron (1989) study rejects eleven of the fourteen series for which Nelson and Plosser (1982) fail to reject the unit root hypothesis. Thus, his conclusion is quite the reverse of Nelson-Plosser conclusion, i.e., most macroeconomic variables, including aggregate and per capita GNP, are trend stationary.

Various studies challenge the validity of the assumption that the break is exogenous. Banerjee, Lumsdaine, and Stock (1990), Christiano (1992), and Zivot and Andrews (1992) argue that associating a break with a particular external event, such as the oil price shock of 1973, is problematic because there are other important events, such as the 1964 tax cut and the Vietnam War, which are also reasonable candidates for such a break. As an alternative, they propose that the date of the break is treated as endogenous and let the data generating process of the series determine the date of the break.

Zivot and Andrews (1992) apply a modified Dickey-Fuller type test that allows for one endogenously-determined break to examine the robustness of Perron (1989) results. In doing so they choose the break point that minimises the value of t statistic for the null unit root hypothesis not to be rejected[3]. Zivot and Andrews (1992) fail to reject the unit root hypothesis at the five per cent level for four of the ten Nelson and Plosser series for which Perron rejects. Further, they also fail to reject the unit root null hypothesis at the five per cent or ten per cent level for the post-war quarterly US real GNP series.

Responding to those studies, Perron (1997) modifies his 1989 models by allowing the break date to be determined endogenously. With some differences in empirical implementation, he obtains results which are consistent with and supportive of his previous results. Unlike the results of Zivot and Andrews (1992), he could reject the unit root hypothesis for all the series, except the GNP deflator, which he also rejected in the 1989 study. The difference in results arising from his and Zivot and Andrews (1992) studies warrant a scrutiny as to what factors might be responsible.

The empirical models used by Zivot and Andrews (1992) and Perron (1997) are both derived from those of Perron (1989). They consist of three models. Model A, B and C are designed to estimate a time-series process that experiences a shift in the intercept, a change in the trend slope, and a combination of them, respectively, in which the date of the break is determined endogenously. Following the notations of Zivot and Andrews (1992), the three models are specified below, which are nothing but Dickey-Fuller type tests augmented by the incorporation of break dummies[4].

<u>Model A</u>

$$y_t = \hat{\ }^{A} + \hat{d}^{A} D_t(\hat{T}B) + \hat{\ }^{A} DU_t(\hat{T}B) + \hat{\ }^{A} t + \hat{\ }^{A} y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^{A} \Delta y_{t-j} + \hat{e}_t \quad \ldots \ 1$$

Model B

$$y_t = \hat{\ }^B + \hat{\ }^B t + \hat{\ }^B DT_t^*(\hat{TB}) + \hat{\ }^B y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^B \Delta y_{t-j} + \hat{e}_t \quad \text{.............} \ 2$$

Model C

$$y_t = \hat{\ }^C + \hat{\ }^C DU_t(\hat{TB}) + \hat{\ }^C t + \hat{d}^C D_t(\hat{TB}) + \hat{\ }^C DT_t(\hat{TB}) + \hat{\ }^C y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^C \Delta y_{t-j} + \hat{e}_t \quad \text{..............} \ 3$$

where *TB* is the break date, D_t is pulse dummy with 1 for $t = TB$ and zero otherwise, $DU_t = 1$ if $t > TB$ and *0* otherwise is a post-break intercept dummy, and $DT_t^* = t - TB$ if $t > TB$ and *0* otherwise is a post-break slope dummy. In Model C of Perron (1997), $DT_t = t$ if $t > TB$ and *0* otherwise.

The main differences between Zivot and Andrews (1992) and Perron (1997) lie in the empirical implementation. The following discussion elaborates their points of differences.

a. Unlike Zivot and Andrews (1992), Perron (1998) distinguishes breaks that occur instantly from those that occur slowly over time[5]. He applies *AO* (additive outlier) models to the former cases and *IO* (innovational outlier) models to the latter cases. While the *IO* regression model constitutes one-step empirical procedure as applied by Zivot and Andrews (1992) to all three cases, the *AO* models involve two-step procedures as outlined in equation (4). These two steps involve: (*i*) detrending the series by regressing it on the trend components (constant, time-trend, and break dummy), and (*ii*) applying the ADF test without trend function to the residuals of the first step. Vogelsang and Perron (1998), through simulations for finite samples, show that, regardless of the DGP following *IO* or *AO*, the application of *AO* model is superior when the break date is chosen using the significance of the trend break parameters, because the test size is invariant to the change in the magnitude of the breaks[6]. They also found that the test size is more sensitive to slope shifts than to intercept shifts[7]. They conclude that the *AO* statistics should be used if large slope shifts are suspected regardless of whether the *AO* or *IO* model better approximates the true DGP since size is not inflated in either case. For this reason, Perron (1997) applies *AO* framework to model B that involves the following two steps.

$$(1)\ y_t = \hat{\ }^B + \hat{\ }^B t + \hat{\ }^B DT_t^*(\hat{TB}) + \tilde{y}_t$$

$$and$$

$$(2)\ \tilde{y}_t = \hat{\ }^B \tilde{y}_{t-1} + \sum_{i=1}^{k} \hat{c}_i^B \Delta\tilde{y}_{t-i} + e_t \quad \text{......} \ 4$$

b. While Zivot and Andrews (1992) employ the general-to-specific method based only on *t* statistic (*t*-Sig) in selecting the truncation lag parameter (*k*), Perron (1997) applies the same method based on both *t* statistic (*t*-Sig) and *F* test (*F*-Sig) for selecting the truncation lag parameter (*k*).

c. Zivot and Andrews (1992) estimate the break date by minimising the value of *t* statistic for $\alpha=1$ in all three models. Perron (1997), on the other hand, applies three different methods in estimating the break date. These methods are: (*i*) minimising the value of *t* statistic for $\alpha=1$, (*ii*) maximising the absolute value (due to unknown sign of the break *a priori*) of the *t* statistic on the trend break parameters ($|t_{\hat{\theta}}|$ for model A and $|t_{\hat{\gamma}}|$ for model B and C), and (*iii*) minimising (by imposing known sign of the trend break *a priori*) the value of *t* statistic on the trend break parameters ($t_{\hat{\theta}}$ for model A and $t_{\hat{\gamma}}$ for models B and C). Perron (1997) conducts simulations using finite samples to examine possible effects of these three different methods of estimating break dates on the power of the test. Based on these simulations, he shows that when the break date is estimated by method (*iii*) the critical values for t are smaller in absolute value than those when the break date is determined by method (*i*) in all three models. Thus, the break date selection method (*iii*) provides more powerful test than method (*i*).

If the sign of the slope or intercept changes is not known *a priori* and then the break date is estimated using method (*ii*). In such a case the power of the test is found to be greater than the one given by method (*i*) only in model C and similar in the other two models.

METHODOLOGY

Empirical Implementation of Unit Root Tests with One Structural Break

Based on those comparisons, the following is a framework for choosing empirical procedures to implement unit root tests that leads to tests with good power and least size distortion:

a. In cases where a shift in trend slope is assumed and the sign of the shift is known *a priori* the possible best framework is *AO* of model B in which the break date is estimated by minimising (for negative sign) or maximising (for positive sign) the value of t statistic for the coefficient on the post-break slope dummy.
b. In cases where the type and sign of the shift with relatively small magnitude is not known *a priori*, the possible best framework is *IO* of model C in which the break date is estimated by maximising the absolute value of t statistic for the coefficient on the post-break slope dummy.
c. When a shift in intercept is assumed, *IO* or *AO* frameworks applied to model A where the break date is estimated by maximising the absolute value of t statistic for the coefficient on the post-break intercept possibly performs better than the one where the break date is estimated by minimizing the value of t statistic for $\alpha=1$.
d. Still related to point (c), if the magnitude of the intercept shift is assumed to be big enough *AO* framework for model *A* is more appropriate where the break date is estimated by maximising the absolute value of t statistic for the coefficient on the post-break intercept.

Since the structural break, especially the 1998 financial crisis, experienced by the Indonesian economy arguably is of larger magnitude than those of developed economies and hence the resulting shifts in both intercepts and slopes of its data serries realisations are also possibly large in magnitude. Hence, in this study the unit root tests that allow for one break whose date is determined endogenously we should rely on the additive outlier (*AO*) framework both for models A and B. While the *AO* framework of model *A* involves two steps of equation (5) below, the *AO* framework of model *B* follows equation (4).

$$\begin{aligned}(1)\ y_t &= \hat{\ }^{A} + \hat{d}^{A} D_t(\hat{T}B) + \hat{\ }^{A} t \\ &\quad + \hat{\ }^{A} DU_t(\hat{T}B) + \tilde{y}_t \\ &\qquad \textit{and} \qquad \ldots\ldots\ldots\ldots 5\end{aligned}$$

$$(2)\ \tilde{y}_t = \hat{\ }^{A} \tilde{y}_{t-1} + \sum_{i=1}^{k} \hat{c}_i^{A} \Delta\tilde{y}_{t-i} + e_t$$

The break date is estimated by maximising the absolute value of t statistic for the coefficient on the intercept break ($|t_{\hat{\theta}}|$) for model *A* and by minimising the t statistic for the coefficient on the post-break slope ($|t_{\hat{\gamma}}|$) for model *B*. However, to the extent that the true data generating process is supposedly of model *C* and thus the test power weakens if instead models A and B are adopted, then model *C* is applied using the innovational outlier (*IO*) framework (equation 3) in which the break date is estimated by maximising $|t_{\hat{\gamma}}|$. In all cases the truncation lag parameter *k* is calculated using both *t*-Sig and *F*-Sig methods.

Data

This study involves 6 macroeconomic variables of the Indonesian economy. They are Interbank Call Money Rate (*R1*), Nominal exchange rate (*E*), monetary aggregates represented by *M1* and *M2*, real GDP (*Y*), and the Consumer Price index (*CPI*). While *E* and *R1* series are taken from the International Financial Statistics IMF CD-ROM, the rest of the series are taken from the Indonesian Financial Statistics published monthly by Bank Indonesia and posted on its website <www.bi.go.id>. In addition, the real GDP (Y) series is obtained from the Indonesia Central Bureau of Statistics in quarterly frequency and is converted into monthly frequency by employing the distributive method. All of these series are in monthly frequency from 1984:12 to 2003:12. All variables are expressed in natural logarithm, except the interest rates.

RESULTS

The results of the traditional ADF unit root test are reported in Table 1. Based on the AIC model selection criterion for determining the optimal lag lengths, the unit root null hypothesis is rejected at 5 per cent level for one series only, namely *R1*.

Table 2 presents the results of unit root test that allows for one structural break at unknown time using the model with the break in the intercept (Model *A*). The empirical framework is *AO* (additive outlier) in which the break date is estimated by maximising the absolute value of $t_{\hat{\theta}}$. As shown in the table, each series has two estimation results. While the *first line* is associated with the lag order *k* being determined based on the *t*-Sig method, the *second line* is due to *k* being selected based on the *F*-Sig method. Contrary to expectation, this test procedure can not reject the unit root null hypothesis for all the variables. Thus, this finding reverses some of the traditional ADF unit root test result.

Table 1: ADF Unit Root Test

	Series	TI/I[1]	$t_{\hat{\alpha}}$	S/N
1.	E	TI	-2.382	N
2.	R1	I	-3.598*	S
3.	M1	T	-3.171**	N
4.	M2	I	-1.346	N
5.	Y	I	-1.965	N
6.	CPI	T	-2.729	N

[1] *TI* and *I* denote the regression where the trend function includes both time-trend and intercept and intercept only, respectively (the inclusion of *T* or *I* is determined by the result that most likely rejects the unit root null hypothesis). * and ** denote the rejection of the unit root hypothesis at 5% and 10% level respectively.
S denotes stationary and N denotes nonstationary.

Table 2: Unit Root Test with One Unknown Break in Intercept (AO Model A)[1)]
Break Date is Estimated by Maximizing the Absolute Value of $t_{\hat{\theta}}$

$$(1)\; y_t = \hat{\ }^{A} + \hat{d}^{A} D_t(\hat{T}B) + \hat{\ }^{A} t + \hat{\ }^{A} DU_t(\hat{T}B) + \tilde{y}_t\;;\quad (2)\; \tilde{y}_t = \hat{\ }^{A}\tilde{y}_{t-1} + \sum_{i=1}^{k}\hat{c}_i^{A}\Delta\tilde{y}_{t-i} + e_t$$

No	Variable	$\hat{T}B$		$\hat{\beta}$	$t_{\hat{\beta}}$	$\hat{\theta}$	$t_{\hat{\theta}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	S/N
1.	E	1997:11	14	0.004673	21.17472	1.02810	32.84214	0.85376	-3.28732	N
		1997:11	15			1.02810	32.84214	0.82873	-3.65525	N
2.	R1	1999:06	15	0.143767	8.72362	-21.8175	-8.50125	0.81222	-3.80482	N
		1999:06	15			-21.8175	-8.50125	0.81222	-3.80482	N
3.	M1	1997:11	12	0.014045	123.2597	0.174205	10.77736	0.79981	-3.04445	N
		1997:11	15			0.17420	10.77736	0.76078	-3.32155	N
4.	M2	2001:11	14	0.019742	190.1297	-0.34685	-15.75839	0.88165	-2.78633	N
		2001:11	14			-0.34685	-15.75839	0.88165	-2.78633	N
5	Y	1998:03	14	0.006017	76.23875	-0.33383	-29.3574	0.92844	-1.36724	N
		1998:03	14			-0.33383	-29.3574	0.92844	-1.36724	N
12.	CPI	1998:03	13	0.006720	117.3693	0.442366	53.62939	0.74817	-3.08550	N
		1998:03	15			0.44237	53.62939	0.74087	-2.89951	N

Note: *1)* The critical value at 2.5%, 5%, and 10% are -4.40, -4.17, and -3.90, respectively (Vogelsang and Perron, 1998); Both *t*-Sig and *F*-Sig methods are used in estimating the truncation parameter (*k*) (the first and second lines in each cell are associated with the former and the latter respectively).
S denotes stationary and N denotes non-stationary.

As for the estimated break dates, both lag order determination methods (*t*-Sig and *F*-Sig) provide exactly the same result for each series. Except for *M2* the estimated break dates coincides with the period of financial crisis. Further, all estimated coefficients on the dummy for intercept shift ($\hat{\theta}$) are significant at 5 per cent level.

When model *B* is estimated using also *AO* framework where the break date is estimated by minimising the value of $t_{\hat{\gamma}}$, the results do not change. In this case only the lag order is determined by using *t*-Sig method. The null hypothesis for *R1* series which previously cannot be rejected at any level can now be rejected at 10 per cent. In comparison, model *A* results in more sensible estimated break dates than those resulting from model *B* estimation as far as their coincidence with the financial crisis period is concerned[8].

Inspecting the series plots in Figure 1 raises suspicion that model *A* and *B* might not be the true DGP for such series as *M2*, *Y*, and *CPI* thereby resulting tests with low power, which in turn fails to reject the null hypothesis. As seen in the figure, the plots of these series follow more closely a broken trended line with changing slope after the break than just a broken trended line. Therefore, model *C* might be the correct DGP for these series.

Figure 1: Plots of Monthly Data Series
with Two Endogenously-Determined Break Dates (TB1 and TB2)

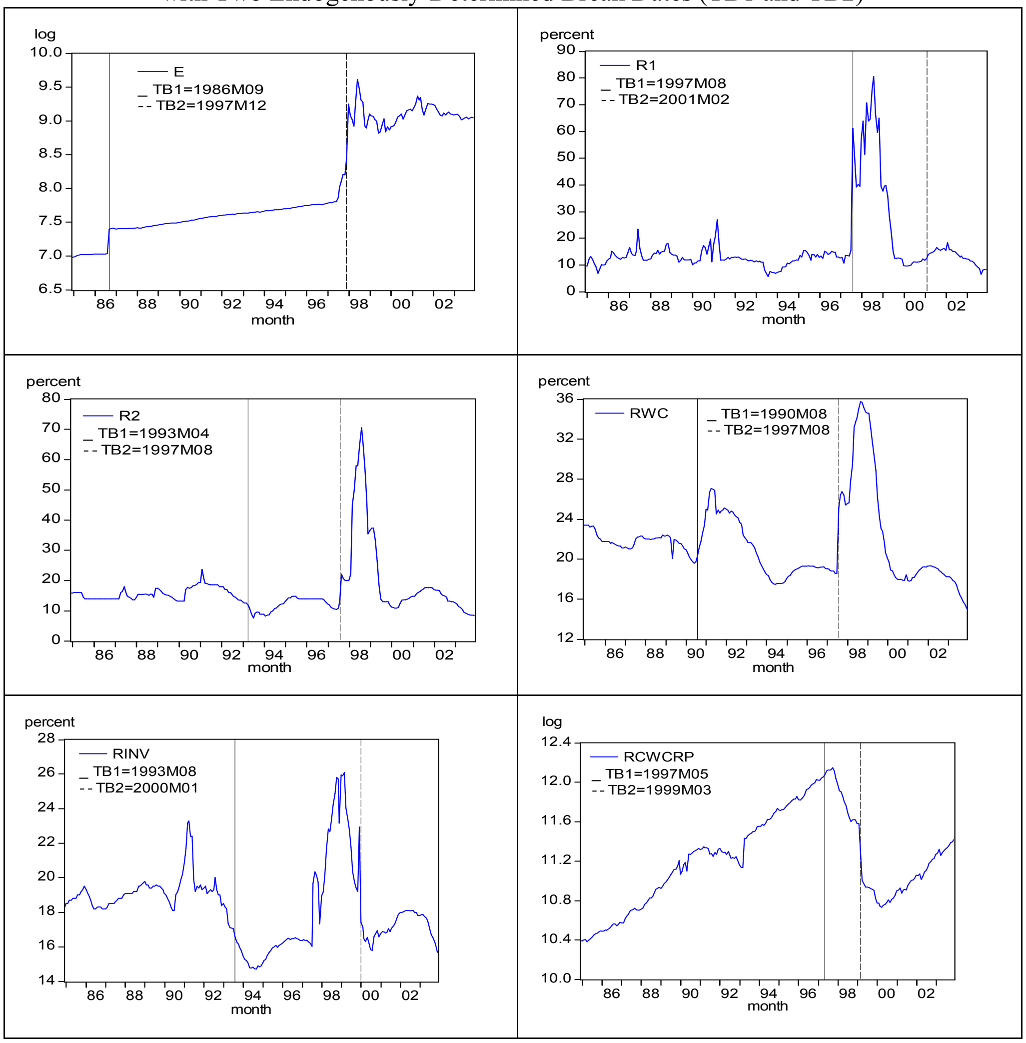

Figure 1: Continued

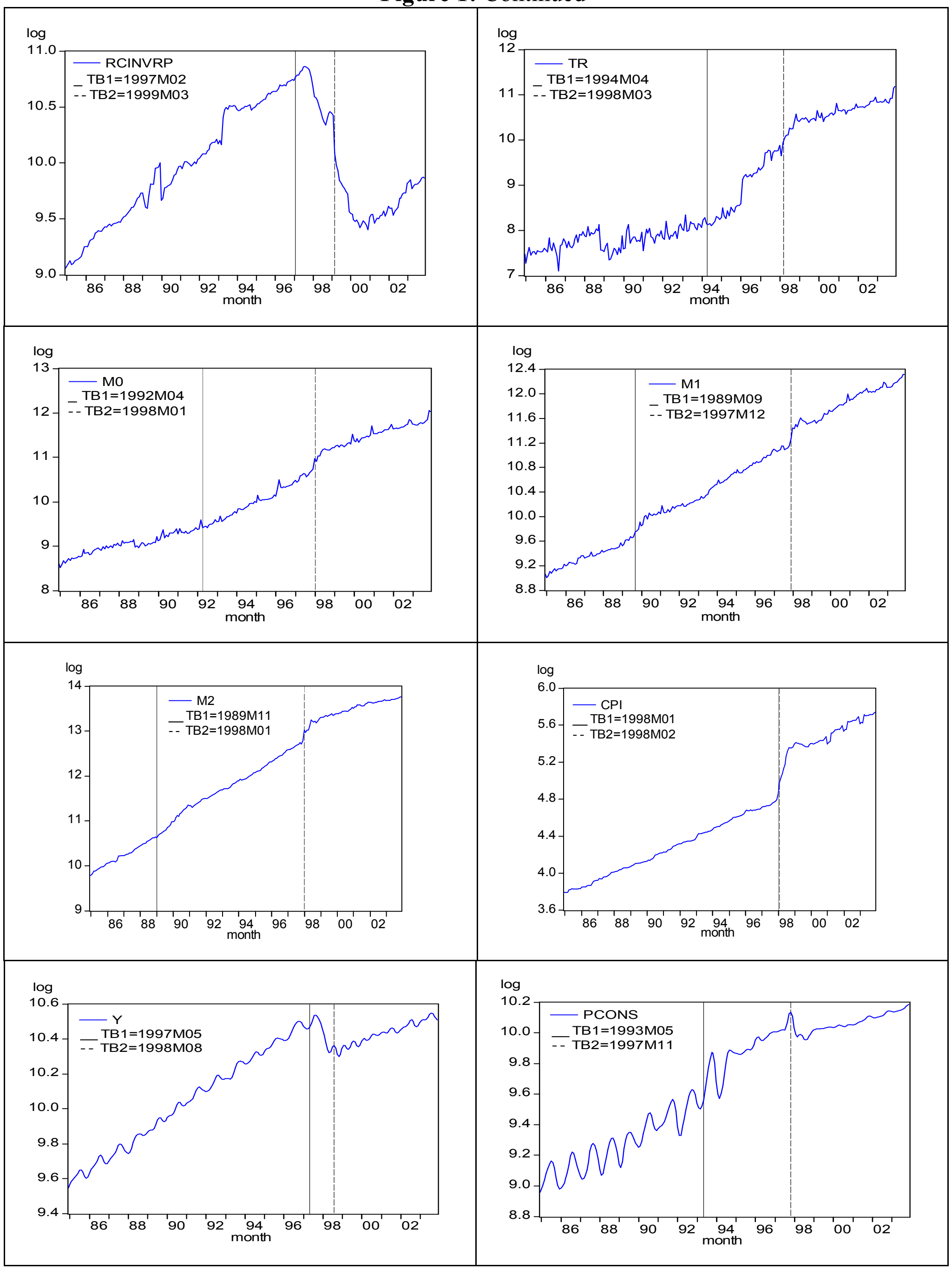

Table 3: Unit Root Test with One Unknown Break in Slope (AO of Model B)
Break Date is Estimated by Minimizing the Value of $t_{\hat{}}$

$$(1)\ y_t = \hat{}^{B} + \hat{}^{B} t + \hat{}^{B} DT_t^{*}(\hat{T}B) + \tilde{y}_t \ ;\ (2)\ \tilde{y}_t = \hat{}^{B} \tilde{y}_{t-1} + \sum_{i=1}^{k} \hat{c}_i^{B} \Delta \tilde{y}_{t-i} + e_t$$

No	Variable	$\hat{T}B$	1)	$\hat{}$	$t_{\hat{}}$	$\hat{}$	$t_{\hat{}}$	$\hat{}$	$t_{\hat{}}$	S/N
1.	E	2002:10	14	0.0107	31.947	-0.0163	-1.5117	0.9584	-2.0670	N
		2002:10	14	0.0107	31.947	-0.0163	-1.5117	0.9584	-2.0670	N
2.	R1	1998:10	12	0.1284	7.7460	-0.4796	-7.3564	0.8522	-4.1756**	N
		1998:10	12	0.1284	7.7460	-0.4796	-7.3564	0.8522	-4.1756**	N
3.	M1	2001:11	12	0.0152	165.400	-0.0047	-3.6629	0.8476	-2.9408	N
		2001:11	12	0.0152	165.400	-0.0047	-3.6629	0.8476	-2.9408	N
4.	M2	2000:04	6	0.0201	204.795	-0.0122	-19.758	0.8945	-2.9003	N
		2000:04	13	0.0201	204.795	-0.0122	-19.758	0.8834	-2.7617	N
5.	Y	1995:11	14	0.0065	72.753	-0.0059	-30.234	0.8163	-3.4421	N
		1995:11	14	0.0065	72.753	-0.0059	-30.234	0.8163	-3.4421	N
6.	CPI	2003:09	13	0.0091	69.894	0.0208	0.5911	0.9775	-2.2459	N
		2003:09	13	0.0091	69.894	0.0208	0.5911	0.9775	-2.2459	N

Note: $\hat{T}B$ is the estimated break date; $\hat{}$ is the estimated time trend; $t_{\hat{}}$ is *t*-statistic for $\hat{}$; $\hat{}$ is the estimated coefficient on DT^* dummy; $t_{\hat{}}$ is *t*-statistic for $\hat{}$; 1) The truncation parameter ( ) is calculated using *t*-Sig method (first line) and *F*-Sig method (second line); Critical Values at 1%=**-4.91**; 5%=**-4.36**; and 10%=**-4.07** (Perron, 1997)* and ** denote the rejection of the null hypothesis that *α=1* at 5% and 10% levels respectively. S denotes stationary and N denotes nonstationary.

Table 4: Unit Root Test Results Using IO2 Model (Model C)
Break Date is Estimated by Maximizing the Absolute Value of $t_{\hat{\gamma}}$

$$y_t = \hat{}^{C} + \hat{}^{C} DU_t(\hat{T}B) + \hat{}^{C} t + \hat{d}^{C}(\hat{T}B) + \hat{}^{C} DT_t(\hat{T}B) + \hat{}^{C} y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^{C} \Delta y_{t-j} + \hat{e}_t$$

is based on *t*-sig (first line in cells) and on *F*-Sig (second line in cells)

No	Variable	$\hat{T}B$		$\hat{\theta}$	$t_{\hat{\theta}}$	$\hat{\gamma}$	$t_{\hat{\gamma}}$	$\hat{\alpha}$	$t_{\hat{\alpha}}$	S/N
1.	E	1986:09	14	0.70496	2.48003	-0.03994	-2.68708	0.95651	-2.17338	N
		1997:11	15	0.62776	5.69687	-0.00091	-1.76840	0.55789	-7.91667*	S
2.	R1	1997:06	15	66.35484	8.67609	-0.30797	-8.26789	0.42625	-9.49947*	S
		1998:07	13	-9.59148	-0.95186	0.02782	0.56091	0.91828	-1.60419	N
3.	M1	1997:09	12	0.19407	3.67886	-0.00080	-2.88458	0.70499	-4.6345**	N
		1997:09	15	0.45344	6.28450	-0.00240	-6.27224	0.80532	-5.20771*	S
4.	M2	1997:10	6	0.46029	6.51180	-0.00243	-6.50284	0.80208	-5.64324*	S
		1997:10	13	0.01951	2.74155	-0.00018	-4.27968	0.94425	-6.62891*	S
5.	Y	1997:11	7	0.14117	3.49511	-0.00151	-5.08151	0.59002	-6.90365*	S
		1997:11	7	0.14117	3.49511	-0.00151	-5.08151	0.59002	-6.90365*	S
6.	CPI	1997:12	12	0.06481	3.30944	-0.00032	2.75442	0.71318	-11.0499*	S
		1997:12	13	0.06747	3.43540	-0.00028	2.30782	0.72632	-9.89445*	S

Note: The critical values at 2.5%, 5%, and 10% are -5.20, -4.91, and -4.59 (Perron, 1997); * and ** denote the null hypothesis that α=1 is rejected at 5% and 10% levels respectively.
S denotes stationary and N denotes nonstationary.

As reported in Table 4, the estimation of model *C* using *IO* framework where the break date is estimated by maximising the absolute value of $t_{\hat{}}$ produces markedly different results. There are three clear-cut cases (*M2*, *CPI*, and *Y*) for which the null hypothesis is rejected at 5 per cent level irrespective of the lag order selection method, *t*-Sig or *F*-Sig. This confirms the suspicion that model *C* is more appropriate DGP for these series than models *A* and *B*. Since all estimated important coefficients are significant at 5 per cent level, it is fair to conclude that these three series follow a trend stationary process with a broken trend and changing slope. However, for the remaining three cases the *t*-Sig and *F*-Sig methods of lag order selection provide conflicting results with regard to the rejection of the null hypothesis. Among these cases, there are two (*E* and *M1*) for which the *F*-Sig method could reject the null hypothesis at 5 per cent level but the *t*-Sig could not. The situation is reversed for *R1*. Namely, it is *t*-Sig, rather than *F*-Sig method, that could reject the null hypothesis at 5 per cent level.

Overall, the unit root test that allows for one structural break with unknown date for Indonesian macroeconomic data series provides mixed results. However, if each series is modelled according to its DGP shown by the series plot, the results of the ADF test for *M2*, *CPI*, and *Y* are not robust to one unknown structural break.

CONCLUDING REMARKS

This paper investigates the time series properties of six monthly macroeconomic variables of the Indonesia economy. In particular, it examines whether the degree of integration of these variables is affected by the inclusion of one structural break determined endogenously in the model. These variables are: nominal exchange rate (*E*), interbank call-money rate (*R1*), narrow money (*M1*), broad money (*M2*), consumer price index (*CPI*), and real GDP (*Y*). Estimating a two-step additive outlier (*AO*) model that allows for one structural break either in the intercept or in the trend produces results that cannot reject the unit root null hypothesis for all the variables. This result is true irrespective of whether the general-to-specific method of choosing the lag order is based on *t* or *F* statistics.

The estimation of an innovational outlier (*IO*) model that allows for one structural break in both intercept and trend provides markedly different results. Depending on the lag order selection method, the unit root null is rejected for 3 series when the lag order is endogenously selected based on *t* and *F* statistics, respectively. These two lag order selection methods unambiguously show the rejection of the unit root hypothesis for 3 series (*M2, CPI* and *Y*) only. Further, nearly of the structural breaks dates correspond to the 1998 financial crisis.

However, the fact that those series likely experience more than one structural break, as reflected in their plots, might cause this unit root test to have low power, thereby producing biased conclusions. Therefore, attempts to remedy this by applying a unit root test that allows for two or more structural breaks at unknown dates are worth considering in future studies.

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[1] The examples of policies with break consequences include frequent devaluations, deregulation of both real and financial sectors and policy regime shifts.

[2] His significant contribution possibly lies in his presentation of three different econometric models that capture different effects on the economy of different types of shocks. He calls his first model "crash" model where the Dickey-Fuller's model is augmented by incorporating a break dummy and a post-break intercept dummy that represents a shift in the intercept caused by the 1929 Great Crash. The second model, called "changing growth" model, captures the effect of the oil price shock of 1973 on the economy in a post-break slope dummy that represents a change in the trend slope due to the slowdown in growth following the shock. The third model combines these two effects (changes in slope and intercept) created by the 1929 great crash on some macroeconomic variables. See Perron (1989).

[3] They show that they can no longer use Perron's critical values because the critical values that emerge when the break date is endogenously estimated are at least as large in absolute value as those computed for an arbitrary fixed break date.

[4] The "hats' on the *TB* parameters in 1 to 3 are to emphasise that they correspond to estimated (endogenous) values of the break fraction.

[5] Perron (1989) considers the 1929 Great Crash as an example of structural break that occurred gradually because it lasted several years and hence assuming the DGP is of *IO*, while the 1973 oil price shock as a break that occurred instantly. Accordingly, he modeled these two cases differently by applying *IO* to the former and *AO* to the latter in accordance with the DGP.

[6] Vogelsang and Perron (1998) prove through simulations that the size and power of the *AO* tests when data follow the *IO* model is very similar to when data follow the *AO* model and vice versa.

[7] The test size distortion becomes a problem when the magnitude of (the coefficient on the intercept shift dummy) is 5 to 10 times the standard deviation of the innovation errors, and the magnitude of (the coefficient on the slope shift dummy) is 1 to 2 times of the standard deviation of the innovation errors. See Vogelsang and Perron (1998).

[8] The interval of the estimated structural break dates for model B is longer. It covers a period well before and long after the financial crisis in Indonesia.