

## **Effects of Fiscal Policy on Credit Markets: Evidence in Japan using a Cointegration Analysis**

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*Using a set of cointegration and error correction models with Threshold Autoregressive (TAR) or Momentum Threshold Autoregressive (MTAR) asymmetric adjustment, we investigate whether the effects of fiscal policy (i.e. expansionary or contractionary government spending shocks) on interest rates in Japan are asymmetric or not. Forty years (from 1969 to 2008) of quarterly series on the price of oil, the Consumer Price Index (CPI) used to compute the real price of oil, government spending (nominal value of Federal Government spending of goods and services for consumption and investment) and the interest rate (Three-month Treasury Bill Rate) obtained from the International Monetary Fund's International Financial Statistics CD-ROM were used for the different tests. Empirical results show that the effects of fiscal policy on interest rates are asymmetric in Japan. Furthermore, the impulse response functions indicate that the results are consistent with a dynamic asymmetry in the behavior of fiscal policy movements.*

**JEL Classification Codes:** C22, E6, E24.

**Keywords:** Cointegration, Asymmetric adjustment, monetary policy, Threshold Autoregressive model, Momentum Threshold Autoregressive model.

### **1. Introduction**

Recent years have been characterized by a lively debate on the economic role that governments should play. In recent years Japan has been plagued by a liquidity trap and the government used fiscal policy to correct the problem. In this paper we will investigate the impact of Japanese government fiscal policy on interest rates. After a brief literature review, we introduce the Threshold and Momentum models of cointegration. Next, we present the results of the study and some discussions and implications of the findings.

### **2. Literature Review**

Many authors have investigated the effects of government fiscal policy on output and several views have sprung. We do not have the pretensions of presenting all those views but rather we will discuss just a few of them.

Kandil (2001), using quarterly data for the United States, studied the asymmetry in the effects of government spending shocks. She showed that while interest rates increased in the face of expansionary government spending shocks, there is no evidence of a reduction in the face of contractionary shocks. Consequently, the increased government

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spending crowds out private investment. Moreover, Kandil (2001) showed that there is evidence of a reduction in private consumption as agents anticipate a future increase in taxes to finance the increased government spending. As a result, output growth and price inflation are decreasing despite expansionary government spending shocks, on average, over time. In view of this evidence, Kandil (2001) suggested that public finance considerations ought to dominate attempts to stimulate demand using government spending when near full-equilibrium capacity utilization in the economy. In contrast, contractionary government spending shocks are not offset by an increase in private spending. Hence, demand contraction is pronounced, slowing output growth and price inflation in the face of a reduction in government spending. For Kandil (2001), the implication is that concerns over the pronounced contractionary effects of a reduction in government spending ought to dominate public finance considerations near full-equilibrium.

Sorensen and Yosha (2001) examined the business cycle behavior of state fiscal policy to determine whether policy is asymmetric and, if so, to identify the causes. They concluded that state revenue and expenditure display significant asymmetry over the business cycle, with nearly offsetting effects on the budget surplus. As a result, state fiscal policy tends to mute economic booms to roughly the same degree it mitigates slowdowns. The asymmetries in revenue and expenditure appeared to be associated with balanced budget rules, although their fundamental causes cannot be clearly identified. Sorensen and Yosha (2001) tried to find out the potential causes of asymmetric fiscal policy. They pointed out that asymmetry in fiscal policy may arise for a variety of reasons, including market constraints, balanced budgets rules, lack of budget discipline in upturns, and incumbent political parties trying to influence voting patterns or forcing the hand of future governments.

Sorensen, Wu and Yosha (2001) found evidence of political business cycle asymmetry in the fiscal policy of U.S. state and local governments during the period 1978-1994.

More recently, using sign restrictions on the impulse-response functions and identifying the unexpected variation in government spending by a positive response of expenditure for up to four quarters after the shock, Mountford and Uhlig (2005) find a negative effect in residential and non-residential investment.

In addition, Afonso (2008c) suggests that there is some evidence in favor of the existence of expansionary fiscal consolidations, for a few budgetary items (general government final consumption, social transfers, and taxes).

Biau and Girard (2005) find a cumulative multiplier of government spending larger than one, and positive reactions of private consumption and private investment in France.

De Castro and Hernandez de Cos (2006) use data for Spain and show that, while there is a positive relationship between government expenditure and output in the short-term, in the medium and long-term expansionary spending shocks only lead to higher inflation and lower output.

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Heppke-Falk et al. (2006) use cash data for Germany, and find that a positive shock in government spending increases output and private consumption, although the effect is relatively small.

Giordano et al. (2007) show that, in Italy, government expenditure has positive and persistent effects on output and on private consumption.

Afonso and Sousa (2009) worked on a paper that provides a detailed evaluation of the effects of fiscal policy on the composition of GDP, namely, by estimating the impact of government spending and government revenue shocks on private consumption and private investments as in Gali et al. (2007).

Second, Afonso and Sousa (2008) asked how asset markets (via stock prices and housing prices) are affected by fiscal policy shocks.

Third, Afonso and Sousa (2009) looked at the impact of fiscal policy on the external sector through the effects on exchange rate in line with Monacelli and Perotti (2006).

Fourth, Afonso and Sousa (2009) analyzed the effects of fiscal policy shocks on the growth rate of monetary aggregates, therefore assessing the potential interaction between fiscal policy and monetary policy in the spirit of Davig and Leeper (2005), Chung et al. (2007), and Gali and Monacelli (2008).

Fifth, Afonso and Sousa (2009) looked at the impact of fiscal policy on the labor market, namely, by assessing its impact on wages and productivity. In addition an important contribution of Afonso and Sousa's paper is the use of quarterly fiscal data, which allowed them to identify precisely the effects of fiscal policies. They analyzed empirical evidence from the U.S., U.K., Germany and Italy, respectively for the periods 1970:3-2007:4, 1964:2-2007:4, 1980:3-2006:4, and 1986:2-2004:4. The most relevant findings of Afonso and Sousa's paper can be summarized as follows.

Government spending shocks (i) have, in general, a small effect on GDP; (ii) do not impact significantly on private consumption; (iii) have a negative impact on private investment; (iv) have a varied effect on housing prices that ranges from a positive and persistent effect to a negative effect and gradual recovery according to the country under consideration, a pattern that depends on the effect on (long-term) interest rates; (v) lead to a quick fall in stock prices; (vi) do not impact significantly on the price level and the average cost of refinancing the debt; (vii) have a small and positive effect on the growth rate of monetary aggregates; (viii) lead to a depreciation of the real effective exchange rate; and (ix) have a positive and persistent impact on productivity.

On the other hand, government revenue shocks: (i) have positive (although lagged) effects on GDP and private investment, as a result of the fiscal consolidation; (ii) a positive effect on both housing prices and stock prices that later mean reverts, but the exact impact depends on the effects on (long-term) interest rates; (iii) in general, do not

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have an impact on the price level; and (iv) lead to an appreciation of the real exchange rate

The objective of this paper is to investigate the asymmetric effects of government spending shocks (i.e. fiscal policy) on variables in the credit market (such as interest rate and the price level).

Specifically, we will examine whether expansionary and contractionary government spending shocks have the same effects on the variables mentioned above.

Employing newly developed techniques, the analysis will provide evidence on asymmetry in the effects of policy shocks.

### 3. Data and Methodology

We will modify and improve Kandil's (2001) methodology which is very similar to that of Cover (1992). We will use new and more appropriate estimation techniques such as unit-root and cointegration tests with possibility of error correction. Since Kandil (2001) used data containing over 30 years of observations, it is possible, perhaps likely, that the data are non-stationary.

Therefore, estimation of equations based on such data would subject the results to "spurious" concerns raised by Granger and Newbold (1974). Another concern pertaining to time-series considerations as pointed out by Hendry (1986), Granger (1998) and Hakkio and Rusch (1989), is that estimating economic relationships such as productivity, unemployment and wage to the exclusion of the long-run association between the variables may create biases. This is particularly the case if the variables are cointegrated. Therefore, a better way of proceeding is to investigate whether there exists a long-run relationship between the variables.

Building upon Kandil's (2001) empirical models, several models will be used to test the asymmetric effects of government spending shocks. To test the asymmetric effects of government spending shocks in the credit market, the following model will be estimated:

$$r_t = \alpha_0 + \alpha_1 g_t + \alpha_2 o_t + u_t \quad (1)$$

Where  $r$  is the log value of the nominal interest rate;  
 $g$  is the log value of government spending;  
 $o$  is the log value of energy price;  
 $u_t$  is the disturbance terms.

In terms of Equation (1), if the output, the government spending, and the energy price are all  $I(1)$  and the linear combination  $r_t - \gamma_0 - \gamma_1 g_t - \gamma_2 o_t = u_t$  is stationary, then the variables are cointegrated of order (1.1). The vector  $x_t$  is  $(r_t, 1, g_t, o_t)'$  and the cointegrating vector  $\gamma$  is  $(1 - \gamma_0 - \gamma_1 - \gamma_2)$ . According to Enders (1995), the system is in

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long-run equilibrium when  $\gamma x_t = 0$ . The deviation from long-run equilibrium-called the equilibrium error- is  $u_t$  so that  $u_t = \gamma x_t$ .

The equations mentioned above will be estimated for long-run relationship and for cointegration allowing for TAR and MTAR adjustment following Engle and Granger's (1987) methodology. The testing procedure is described in the following section.

### 3.1 Threshold and Momentum Models of Cointegration

The Engle and Granger (1987) methodology as applied to the fiscal policy model begins by positing a long-run equilibrium relationship of the forms given for example in equation (1).

The next step in the Engle and Granger procedure focuses on the OLS estimate of  $\rho$  in the following regression equation:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \quad (2)$$

Where the estimated regression residuals from (1) are used to estimate (2).

Rejecting the null hypothesis of no cointegration (i.e., accepting the alternative hypothesis  $-2 < \rho < 0$ ) implies that the residuals in (2) are stationary with mean zero. As such, equation (2) is an attractor such that its pull is strictly proportional to the absolute value of  $\mu_{t-1}$ . The change in  $\mu_t$  equals  $\rho$  multiplied by  $\mu_{t-1}$  regardless of whether  $\mu_{t-1}$  is positive or negative.

The implicit assumption of symmetric adjustment is problematic if the money-supply shock adjustment is asymmetric. A formal way to introduce asymmetric adjustment is to let the deviations from the long-run equilibrium in equation (1) behave as a Threshold Autoregressive (TAR) process. Thus, it is possible to replace (2) with:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \varepsilon_t \quad (3)$$

Where  $I_t$  is the Heaviside indicator such that:

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq 0 \\ 0 & \text{if } \mu_{t-1} < 0 \end{cases} \quad (4)$$

Asymmetric adjustment is implied by different values of  $\rho_1$  and  $\rho_2$ ; when  $\mu_{t-1}$  is positive, the adjustment is  $\rho_1\mu_{t-1}$ , and if  $\mu_{t-1}$  is negative, the adjustment is  $\rho_2\mu_{t-1}$ . A sufficient condition for stationarity of  $\{\mu_t\}$  is for:  $-2 < (\rho_1, \rho_2) < 0$ . Moreover, if the  $\{\mu_t\}$  sequence is stationary, the least squares of estimates of  $\rho_1$  and  $\rho_2$  have an asymptotic multivariate normal distribution if the value of the threshold is known (or consistently estimated). Thus, if the null hypothesis  $\rho_1 = \rho_2 = 0$  is rejected, it is possible to test for symmetric adjustment (i.e.,  $\rho_1 = \rho_2$ ) using a standard  $F$ -test.

Since adjustment is symmetric if  $\rho_1 = \rho_2$ , the Engle-Granger test for cointegration is a special case of (3).

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Since the exact nature of the non-linearity may not be known, it is also possible to allow the adjustment to depend on the change in  $\mu_{t-1}$  (i.e.  $\Delta\mu_{t-1}$ ) instead of the level of  $\mu_{t-1}$ . In this case, the Heaviside indicator of (4) becomes:

$$I_t = \begin{cases} 1 & \text{if } \Delta\mu_{t-1} \geq 0 \\ 0 & \text{if } \Delta\mu_{t-1} < 0 \end{cases} \quad (5)$$

Even though Hansen (1997) shows that setting a Heaviside indicator using  $\Delta\mu_{t-1}$  can perform better than the specification using pure TAR adjustment, Enders and Granger (1998) and Enders and Siklos (2001) show that the series exhibits more “momentum” in one direction than the other. They call this model Momentum-Threshold Autoregressive (M-TAR) model. Respectively, the  $F$ -statistics for the null hypothesis  $\rho_1 = \rho_2 = 0$  using the TAR specification of (4) and the M-TAR specification of (5) are called  $\Phi_\mu$  and  $\Phi_\mu^*$ . As there is generally no presumption as to whether to use (4) or (5), the recommendation is to select the adjustment mechanism by a model selection criterion such as the  $AIC$ .

If the errors in equation (3) are serially correlated, it is possible to use an augmented threshold model for the residuals. In this circumstance, equation (3) is replaced by:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1-I_t)\rho_2\mu_{t-1} + \sum_{i=1}^p \beta_i\Delta\mu_{t-i} + \varepsilon_t \quad (6)$$

The distributions of  $\Phi_\mu$  and  $\Phi_\mu^*$  depend on the number of observations, the number of lags in equation (3) and the number of variables in the cointegrating relationship. Enders and Siklos (2001) report critical values using cointegrating vectors containing up to three variables. Since the model used in this study contains four variables in the cointegrating relationship, we develop first the critical values for the four-variable case.

### 3.2 Data

The International Monetary Fund’s International Financial Statistics provides the core data for Japan. To test the effects of fiscal policy on the credit market, forty years of quarterly series (from 1969 to 2008) on the price of oil, the Consumer Price Index (CPI) used to compute the real price of oil, government spending (nominal value of Federal Government spending of goods and services for consumption and investment) and the interest rate (Three-month Treasury Bill Rate) are collected.

### 3.3 Impulse Response Functions

According to Enders (1995), the shape of the impulse response functions and the results of the variance decompositions can indicate whether the dynamic responses of the variables conform to theory. We shall examine different responses of output to government spending shocks, tax revenues shocks and real price of oil shocks.

## 4. Empirical Results

### 4.1 Results for the Test of Stationarity

Since by definition cointegration necessitates that the variables be integrated of the same order, we pretested the variables for their order of integration. The Dickey-Fuller (DF) test and the Augmented Dickey-Fuller (ADF) test are used to infer the number of unit roots (if any) in the output, money supply, price of oil and interest rate series for each country. Lag 8, which is determined by the Akaike Information Criterion (AIC), is used in the ADF test. Table 1 provides the summary of the findings. Both the DF and ADF tests fail to reject the null hypothesis (all the t-statistics are less than the DF critical values) of any of the variables. These findings suggest that the series are all I (1). As a result, the series are non-stationary and we proceed by taking first differences of the series and test for cointegration.

**Table 1: Unit Root Test Results**

	<i>DF Test</i>		<i>ADF Test</i>		
	<i>Lag</i>	<i>t</i>	<i>Lag</i>	<i>AIC</i>	<i>t</i>
<i>Japan</i>					
<i>R</i>	0	-0.56	8	-466.69	-1.18
<i>Po</i>	0	-1.78	8	-319.77	-1.98
<i>Gx</i>	0	2.39	8	-564.79	1.65

The Dickey Fuller (DF) critical values for T = 100 are -3.51, -2.89, -2.58, at 1%, 5% and 10% levels respectively.

r = interest rate; po = price of oil; gx = Government Expenditures.

### 4.2 Results of the Long Run Relationships in the Credit Market

We use OLS to estimate the long run relationships of the variables interest rate, government spending (or government expenditures) and price of oil as implied by Equation (1) above. The long run cointegrating equations for Japan are as follow (the t-statistics are in parentheses):

$$r_t = 5.119 - 0.631g_t + 0.840o_t \quad (7)$$

(5.11) (-3.69) (4.59)

Where *r* is the log value of the nominal interest rate;  
*g* is the log value of government spending;  
*o* is the log value of energy price.

Economic theory predicts that the coefficients on the government spending (*g<sub>t</sub>*) should be positive and less than one. It predicts also that the coefficient on the price of oil

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should be positive and less than one. As anticipated, all the cointegrating parameters in the above equations turn out to be consistent with the economic theory except for the cointegrating parameter of the government spending variable. The cointegrating parameters of the variables are all statistically significant at conventional statistical significance levels.

The residuals found in the above equations (7) are used to proceed with the cointegration and asymmetric adjustment tests.

### **4.3 Cointegration and Asymmetric Tests with Possibility of TAR or MTAR Adjustment in the Credit Market**

To test for cointegration and asymmetry, we saved the residuals of the long-run relation equations (7) in the sequence  $\{\mu_{1t}\}$ . For each type of asymmetry, we set the indicator function  $I_t$  according to Equation (4) or Equation (5) and estimated an equation in the form of Equation (6). The AIC was used to select the most appropriate lag length  $p$  and adjustment mechanism (i.e. TAR versus MTAR adjustment). The sample value of the F-statistic for the null hypothesis  $\rho_1 = \rho_2 = 0$  was compared with the appropriate critical value reported in Table 2 and/or 3. If the alternative hypothesis is accepted (i.e., the null hypothesis of no cointegration is rejected), we then used Chan's (1993) methodology to find the consistent estimate of the threshold. After all, there is no reason to presume that the threshold is identically equal to zero. Once the threshold is properly estimated, we test for symmetric versus asymmetric adjustment (i.e., we test the null hypothesis  $\rho_1 = \rho_2$ ) using the usual F-statistic.



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**Table 2: Distribution for the F-Statistic for the Null Hypothesis  $\rho_1 = \rho_2$  in the 4-variable case**

	LAGGED CHANGES											
	1 LAG			2 LAGS			3 LAGS			4 LAGS		
	90%	95%	99%	90%	95%	99%	90%	95%	99%	90%	95%	99%
THE TAR MODEL: $\phi_\mu$												
50	8.98	10.53	13.91	8.28	9.63	12.85	7.92	9.31	12.39	7.36	8.64	11.56
100	8.45	9.77	12.63	8.09	9.36	12.05	7.92	9.13	11.82	7.61	8.84	11.39
150	8.40	9.65	12.42	8.17	9.41	11.97	8.04	9.22	11.88	7.84	9.00	11.61
200	8.53	9.71	12.44	8.33	9.57	12.17	8.27	9.47	12.13	8.11	9.24	11.74
250	8.54	9.79	12.58	8.46	9.66	12.28	8.41	9.62	12.05	8.31	9.45	12.01
500	8.81	10.03	12.73	8.74	9.92	12.60	8.69	9.86	12.40	8.70	9.92	12.44
THE MTAR MODEL: $\phi_\mu^*$												
50	9.85	11.47	14.91	9.05	10.53	13.81	8.75	10.18	13.58	8.14	9.53	12.58
100	9.55	10.89	13.81	9.12	10.46	13.33	8.91	10.21	13.03	8.62	9.84	12.66
150	9.28	10.63	13.44	9.01	10.28	12.99	8.86	10.08	12.76	8.62	9.88	12.46
200	9.19	10.46	13.17	8.98	10.28	13.04	8.83	10.07	12.68	8.72	9.95	12.50
250	9.10	10.36	13.09	8.96	10.17	12.90	8.87	10.16	12.82	8.72	9.90	12.54
500	9.05	10.30	12.99	8.99	10.24	12.81	8.92	10.18	12.76	8.80	10.03	12.54

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**Table 3: Distribution for the F-Statistic for the Null Hypothesis  $\rho_1 = \rho_2$  in the 4-variable case**

	LAGGED CHANGES											
	5 LAGS			6 LAGS			7 LAGS			8 LAGS		
	90%	95%	99%	90%	95%	99%	90%	95%	99%	90%	95%	99%
THE TAR MODEL: $\phi_\mu$												
50	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A
100	7.45	8.60	11.15	7.19	8.29	10.71	7.02	8.15	10.57	6.80	7.89	10.22
150	7.65	8.80	11.26	7.48	8.60	10.91	7.40	8.56	10.95	7.19	8.29	10.66
200	7.95	9.13	11.65	7.78	8.93	11.38	7.76	8.91	11.37	7.61	8.71	11.14
250	8.22	9.40	12.01	8.03	9.19	11.71	8.00	9.13	11.68	7.87	8.96	11.34
500	8.63	9.83	12.36	8.53	9.73	12.20	8.49	9.67	11.98	8.48	9.66	12.12
THE MTAR MODEL: $\phi_\mu^*$												
50	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A	N/A
100	8.40	9.66	12.25	8.17	9.36	12.08	8.05	9.26	11.85	7.75	8.91	11.52
150	8.47	9.67	12.34	8.29	9.46	12.07	8.16	9.33	11.93	7.97	9.10	11.66
200	8.60	9.82	12.39	8.42	9.60	12.23	8.32	9.51	12.05	8.19	9.35	11.81
250	8.63	9.84	12.40	8.53	9.76	12.20	8.45	9.62	12.12	8.33	9.49	12.02
500	8.77	10.04	12.65	8.72	9.87	12.43	8.72	9.88	12.41	8.58	9.78	12.19

Note: NA indicates not available. We do not provide the critical values for the model with more than 5 lags using only 50 observations.

Table 4 reports the estimated values for the  $\rho_i$  and the sample F-statistics for the null hypothesis  $\rho_1 = \rho_2 = 0$  as well as the F-statistics for the null hypothesis  $\rho_1 = \rho_2$  using the lag length and adjustment mechanism selected by the AIC.

The lag length is selected such that the Akaike Information Criterion (AIC) is minimized. As such AIC selects 8 lags for Japan.

**Table 4: The Estimated Adjustment Equations (interest rates)**

Countries	$\rho_1$	$\rho_2$	$\phi^a$	$\rho_1 = \rho_2^b$	Lags <sup>c</sup>	AIC	Flag
Japan	0.079 (1.02)	-0.048 (-1.63)	8.97**	3.43**	8	245.04	MTAR

<sup>a</sup> Entries are the sample values of  $\phi_\mu$  or  $\phi_\mu^*$  for the adjustment process shown in column 8.

\*,\*\*,\*\*\* indicate significance levels of 10%, 5% and 1% respectively. t-statistics are in parentheses.

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<sup>b</sup>Entries in this column are the values for the sample F-statistic for the Null Hypothesis that the adjustment equations are equal.

<sup>c</sup>Entries in this column are the number of lags of  $\{\Delta\mu_t\}$  selected by AIC.

The estimated  $\phi^*$  for the null hypothesis  $\rho_1 = \rho_2 = 0$  is 8.97 in Japan. The critical value reported in Table 3 at the 5% significance level with 150 observations and 8 lags is 8.29. As such, we reject the null hypothesis of a unit root in favor of stationarity with asymmetric adjustment.

The point estimates of the  $\rho_i$  suggest that negative deviations from the long-run equilibrium are eliminated much faster than positive deviations. Specifically in Japan, 7.9% of a negative deviation is eliminated within a quarter while 5% of a positive deviation is eliminated during the same time frame.

The F-statistic for asymmetric adjustment ( $\rho_1 = \rho_2$ ) is given in Table 4. For the Japan, the F-statistic (3.43) rejects symmetric adjustment at the 5% significance level in favor of asymmetric *MTAR* adjustment. According to Lutkepohl (1994) the coefficients of cointegration relations cannot be interpreted as elasticities. This is because the ceteris paribus cannot be meaningful. The error correction specifications can be more informative.

### 4.4 The Error-Correction Representation: TAR or MTAR Asymmetric Adjustment in the Credit Market (Interest Rate)

Having found evidence supporting asymmetric adjustment in Japan and using the long relation implied in Equation (7), the estimated error-correction equations assuming Momentum Threshold Autoregressive (*MTAR*) adjustment (with t-statistics in parentheses) are:

$$\Delta o_t = A_{11}(L)\Delta o_{t-1} + A_{12}(L)\Delta g_{t-1} + A_{13}(L)\Delta r_{t-1} + 0.1721z_{t-1}^+ - 0.1185z_{t-1}^- \quad (8)$$

(3.24)      (-0.26)

$$\Delta g_t = A_{21}(L)\Delta o_{t-1} + A_{22}(L)\Delta g_{t-1} + A_{23}(L)\Delta r_{t-1} - 0.0152z_{t-1}^+ + 0.0174z_{t-1}^- \quad (9)$$

(-2.83)      (0.68)

$$\Delta r_t = A_{31}(L)\Delta o_{t-1} + A_{32}(L)\Delta g_{t-1} + A_{33}(L)\Delta r_{t-1} - 0.0308z_{t-1}^+ - 0.0171z_{t-1}^- \quad (10)$$

(-0.79)      (-0.58)

Where:

$$z_t^+ = I_t(r_t - 5.119 + 0.631g_t - 0.840o_t),$$

$$z_t^- = (1 - I_t)(r_t - 5.119 + 0.631g_t - 0.840o_t),$$

$I_t$  = Threshold Heaviside Indicator Function,

$A_{ij}(L)$  is a polynomial in the lag operator L, and the lag length is selected using the multivariate version of AIC, which selected 8 lags

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In Equation (10), the interest rate in Japan seems to adjust faster when there is a positive discrepancy from the long-run equilibrium than when there is a negative discrepancy. Specifically, the point estimates imply that the interest rate adjusts by 3% of a positive deviation from long-run equilibrium, but by only 1.7% of a negative gap. The t-statistics imply that neither of the coefficients on the error-correction terms is significant at conventional significance levels.

In equations (8) and (9), only the coefficients of the positive error-correction terms (i.e.  $z^+$ ) are significant at conventional significance levels. These results indicate that adjustments towards the long-run equilibrium are accomplished via changes in the government spending and the real price of oil.

### 5. Conclusion and Policy Implications

This study investigates the asymmetric effects of fiscal policy (i.e. government spending) on interest rates (proxying credit market conditions) in Japan. Forty years of quarterly series on the price of oil, the Consumer Price Index, government revenues, government spending and the interest rate for Japan are obtained from the International Monetary Fund's International Financial Statistics CD-ROM. We use a cointegration analysis developed by Engle and Granger to test the effects of fiscal policy shocks on output.

The Engle-Granger cointegration analysis finds that the variables in the credit market are cointegrated which is an indication of the existence of long-term equilibrium relationships between the interest rate, government spending and the the price of oil in both countries. The cointegrating parameters in the credit market turn out to be consistent with economic theory predictions except for the cointegrating parameter of the government spending variable.

The asymmetric adjustment tests conclusively indicate that the effects of fiscal policy on interest rates are asymmetric.

The impulse response function indicates that the results are consistent with a dynamic asymmetry in the behavior of the government spending movements.

Specifically, a negative government spending shock produces a significantly larger response in interest rates as compared to its positive counterpart. Furthermore, the error-correction representation reveals that the effects of positive government spending movements dwindle rather quickly in Japan. The results for a negative government spending change, however, do not dwindle that quickly. These results show evidence that the interest rate decreases in the face of contractionary government spending shocks in Japan. Furthermore, there is no evidence that the interest rate significantly increases in the face of expansionary government spending shocks.

The results found in Japan are not consistent with the findings of the studies of previous authors such as Kandil (2001).

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Overall the results found in this study and described above will have interesting policy implications. In Japan, expansionary government spending shocks do not have positive effects on the interest rate. The interest rate is not responsive to government spending and output growth. Perhaps one explanation why interest rates are not responsive to government spending is that Japan has been plagued by a liquidity trap and a stubborn deflation for quite some time. As such, public finance considerations ought to dominate attempts to stimulate the economy using government spending during recessions because chances are that the increase in government spending will not crowd out private spending. Furthermore, a vigorous fiscal policy in Japan may help fight against the stubborn price deflation because the price level is very responsive to expansionary government spending shocks. On the other hand, there is evidence that contractionary government spending shocks in Japan will reduce interest rates and output, but not the price level. Therefore, because of extremely low prices and interest rates, contractionary government spending would be an inadequate policy in Japan as it might depress prices and interest rates further.

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