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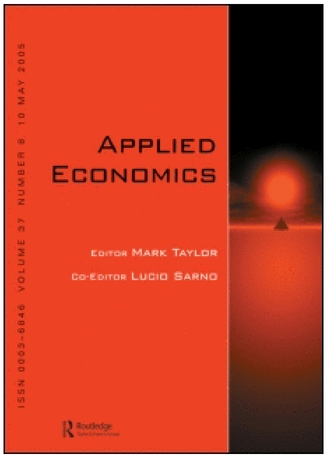

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Evidence on nonlinear error correction in money demand: the case of Taiwan

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This paper proposes a nonlinear error-correction model based upon smooth transition regression methodology. The model is specified such that the short-run adjustment toward long-run equilibrium is nonlinear and that the error correction is a smooth function of long-run deviation. Empirical results obtained from estimating M2 money demand in Taiwan support the hypothesis of a nonlinear error-correction process and provide better interpretation of change in the demand for money.

I. INTRODUCTION

The role of the money demand function as a cornerstone of economic policy has been challenged since the mid-1970s. The traditional approach to the demand for money, from the quantity theory to the Baumol–Tobin model to the portfolio model, has yielded evidence of serious specification errors.[1] In particular, an unstable money demand function not only erodes the grounds for using monetary policy to stabilize the economy, but also misdirects academic research in important issues such as the income elasticity of money demand. Economists continue to search for a specification of the money demand function that gives a reliable relationship with other macro variables. This is an intriguing and important issue in monetary economics.

In the literature, there are fewer disputes on whether a stable long-run demand function actually exists. Friedman and Schwartz (1991) have argued that the underlying characteristics of the money demand function rarely change over a long period of time. In the long run, the demand for money depends mainly on real income, interest rates, and, in an open economy, exchange rates. The challenge is in the search for short-run dynamics in the demand for money that will lead to a steady state.[2] The early approach to formulating the short-run dynamics applied either a partial adjustment process to the desired real money balance or a Koyck transformation of the stock-adjustment process. The specification of the short-run dynamics improved significantly in terms of econometric method when Hendry and Ericsson (1991) first applied an error-correction model to the money demand function. Nevertheless, all these formulations assume a linear and symmetric short-run correction process along which the money demand converges to a long-run equilibrium.

Money and other macroeconomic variables, such as income and the interest rates, may be cointegrated and have a stable long-run equilibrium relation. However, some investigators, such as Muscatelli and Spinelli (1996), and Wolters *et al.* (1998), argue that the error correction to the short-run dynamics may not follow a linear process. For instance, the variations in money are more volatile in an economic downturn than in an upswing.[3] Accordingly, nonlinear adjustment models are needed to formulate the inherently asymmetric fluctuation of money. At the theoretical level, economists have proposed two different models to explain nonlinearity in money

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[1] See Goldfeld (1976) and Hendry and Ericsson (1991), among others.

[2] Some studies such as Barnett (1980) and Spindt (1985) have focused on the construction of a new monetary aggregate that gives a more stable demand for money function.

[3] The earliest literature on the asymmetric business cycles can be traced back to Mitchell (1927).

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demand.[4] The target-bounds model assumes that an agent sets target, upper bound, and lower bound for the real balance. The short-run adjustment will take place whenever the real balance reaches the bounds. The error-correction under this type of model is assumed to be stable and smooth. The other type of model, the buffer stock model, assumes that short-run adjustment occurs only if the real balance deviates significantly from the long-run equilibrium. The speed of adjustment therefore depends on the size of the deviation. Several studies in the literature have applied these micro-based theoretical models to their empirical work. Muscatelli and Spinelli (1996), and Ericsson *et al.* (1998) include a cubic error-correction term as a regressor in their study of the money demand function in Italy and in the United Kingdom, respectively. They find that the nonlinear error-correction specification better describes the short-run dynamics and improves the overall goodness-of-fit. Wolters *et al.* (1998), and Lutkepohl *et al.* (1997) apply a smooth transition error-correction model (STECM) to investigate the issues of instability and nonlinearity of money demand in Germany. Both studies have found a clear structural instability that occurred in 1990, the year of German unification, and a significant asymmetric effect of the inflation rate on the demand for money. Other STECM studies by Teräsvirta and Eliasson (1998), and by Sarno (1999) also find similar nonlinear dynamics in the demand for money in UK and Italy.

Due to the unavailability of reliable data and information, current research on the formulation of a nonlinear error-correction model within the context of the money demand function is limited to industrial countries. However, a newly industrialized country such as Taiwan offers an excellent opportunity to test the new econometric method. Taiwan has a fast-changing financial market and has stood apart from the devastation of the Asian financial crisis that began in July 1997. During the past two decades, it underwent a large number of deregulation measures and financial innovation, including deregulation of capital flows and allowing banks to be privately owned. A study of the money demand in such a buoyant economy provides valuable information to policy makers in both the domestic and international entities. Liang *et al.* (1982) exemplify the research on the demand for money in Taiwan in using a traditional specification and a partial adjustment for the short-run dynamics. They verify that 'the case of missing money' occurred in Taiwan, a conclusion that is similar to the findings of Goldfeld (1976) for the USA.[5] In the search for a better specification of money demand in Taiwan, several subsequent studies have found evidence of cointegration and short-run error-correction dynamics. Chien (1992) has shown the existence of a cointegration relationship among real M1B, real GNP, the three-month time deposit rate, the saving deposit rate, and the real stock price index. Arize (1994) and Lee (1996) both apply the error-correction model to the estimation of the money demand. However, none of these studies takes into consideration the nonlinearity in the demand for money.

This paper applies STECM as proposed by Granger and Teräsvirta (1993) and Terävirta (1994) to the estimation of Taiwan's demand for money. The model allows the demand for money to fluctuate between two distinct regimes as represented by the upper and lower bands of the monetary targets. It also allows the transition between these two regimes to be smooth, so that a continuum of states exists. It further follows the econometric methods of Lin and Teräsvirta (1994), Eitrheim and Teräsvirta (1996), and Teräsvirta and Eliasson (1998), in testing the adequacy of STECM for the Taiwan's demand for money. The rest of the paper is organized as follows: Section II introduces the smooth transition error-correction model. Section III tests the time series properties and establishes a long-run equilibrium that is represented by a cointegrating relation between the real money balance and some macro variables. Section IV presents the results of the Lagrange Multiplier (LM) test on the linearity and the estimation of a nonlinear STECM. Section V offers concluding remarks.

II. THE SMOOTH TRANSITION ERROR-CORRECTION MODELS

Consider an M2 series, denoted as y_t, and a set of k macroeconomic variables, $X_t = (x_{1t}, x_{2t}, \ldots, x_{kt})'$, such as GNP, interest rates, inflation rate, etc. Suppose both y_t and X_t are $I(1)$ series, and are linearly cointegrated with only one cointegrated relation,

$$y_t = \beta' X_t + z_t \qquad t = 1, 2, \ldots, T \tag{1}$$

where z_t, an $I(0)$ series, represents the deviation from the long-run equilibrium, $\beta' X_t$. The standard error-correction model (ECM) has the form,

$$\Delta y_t = \pi_0 + \pi_1' W_t + u_t \tag{2}$$

where $W_t = (z_{t-1}, \Delta y_{t-1}, \ldots, \Delta y_{t-p}, \Delta X_t', \ldots, \Delta X_{t-p}')'$ and $u_t \sim \text{n.i.d.}(0, \sigma^2)$. The coefficient vector π_1 has a dimension of $m \times 1$, where $m = (k+1)(p+1)$. The short-run adjustment, Δy_t, towards the long-run equilibrium is linear in z_{t-1}.

It is argued that the short-run adjustment of the demand for money is inherently nonlinear and asymmetric. The demand elasticities (say π_1) with respect to income, the

[4] See Sarno (1999) for a review.

[5] In conducting a post-sample forecast, Goldfeld has observed a systematic over-prediction of US real money balance after 1974, which is in contrast to the random prediciton errors in the periods following.

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interest rate, and inflation during an economic expansion period are different from those elasticities in a recession period. However, instead of assuming a discrete regime-switching model, this paper follows Granger and Teräsvirta (1993) and Terävirta (1994) in formulating a smooth transition error-correction model (STECM).

Consider a STECM,

$$\Delta y_t = \pi_0 + \pi_1' W_t + (\delta_0 + \delta_1' W_t) F(z_{t-d}; r, c) + u_t \tag{3}$$

The function $F(z_{t-d}; r, c)$ is a continuous transition function with the transition variable z_{t-d} and parameters (r, c) that provides a variety of nonlinear models. For example, a logistic function could be specified,

$$F(z_{t-d}; r, c) = (1 + \exp[-r(z_{t-d} - c)])^{-1} \qquad r > 0 \tag{4}$$

$F(z_{t-d}; r, c)$ is bounded by zero $(F(-\infty; r, c))$ and one $(F(+\infty; r, c))$. The specification allows the 'parameters' of W_t and z_{t-1} in STECM to change monotonically with the delayed disequilibrium error z_{t-d}. The smoothness parameter r determines the speed of the transition. When $r \to 0$, the model reduces to a linear ECM, $\Delta y_t = \pi_0 + \pi_1' W_t + u_t$. When $r \to +\infty, F(z_{t-d}; r, c) = 0$ for $z_{t-d} \leqslant c$ and $F(z_{t-d}; r, c) = 1$ for $z_{t-d} > c$, the STECM becomes a two-regime threshold model. The model specified in Equation 3 with a transition function expressed as Equation 4 is called the logistic STECM (LSTECM). Since $F(z_{t-d}; r, c)$ is not symmetric about c, LSTECM is capable of generating the asymmetric short-run dynamics in two forms. The short-run dynamics takes the form, $\Delta y_t = (\pi_0 + \delta_0) + (\pi_1 + \delta_1)' W_t + u_t$ during the expansion period and $z_{t-d} > c$. The dynamics switches into $\Delta y_t = \pi_0 + \pi_1' W_t + u_t$ during the recession period and $z_{t-d} \leqslant c$. The transition from one state to another is smooth and takes the form $\Delta y_t = \pi_0 + \pi_1' W_t + (\delta_0 + \delta_1' W_t) F(z_{t-d}; r, c) + u_t$.

An alternative specification of STECM assumes the transition function $F(z_{t-d}; r, c)$ to be an exponential function,

$$F(z_{t-d}; r, c) = 1 - \exp[-r(z_{t-d} - c)^2] \qquad r > 0 \tag{5}$$

The model, called the exponential STECM (ESTECM), consists of an error-correction form shown in Equation 3 and a transition function expressed as Equation 5. The ESTECM allows the 'parameters' to change symmetrically about c with z_{t-d}. In the extreme case when $r \to 0$, the model reduces to linear ECM with $\Delta y_t = \pi_0 + \pi_1' W_t + u_t$. When $r \to +\infty$, the model switches to another regime with $\Delta y_t = (\pi_0 + \delta_0) + (\pi_1 + \delta_1)' W_t + u_t$. Since $F(z_{t-d}; r, c)$ is symmetric about c, the ESTECM gives similar short-run dynamics between an expansion and a recession period. It also implies a symmetric transition from one state to another.

Since the linear ECM is nested in the STECM, it is necessary to apply the test for linearity directly to the model of LSTECM and ESTECM. Although it is conceivable to formulate the linearity test by making a null hypothesis $H_0 : r = 0$, the linear ECM is not identified under H_0 because the parameters (δ_0, δ_1, c) can take any values. This problem makes it invalid to apply the Lagrange multiplier (LM)-type test using the asymptotic theory. To deal with this problem, Luukkonen *et al.* (1988) suggest that one can replace the transition function $F(z_{t-d}; r, c)$ with its third-order Taylor approximation about $r = 0$,[6]

$$\Delta y_t = \alpha_0 + \alpha_1' W_t + \beta_1' W_t(z_{t-d}) + \beta_2' W_t(z_{t-d})^2 + \beta_3' W_t(z_{t-d})^3 + \varepsilon_t \tag{6}$$

If one assumes that the delay parameter d is known, the linearity test is equivalent to the test of the hypothesis

$$H_0 : \beta_1 = \beta_2 = \beta_3 = 0 \tag{7}$$

Define an auxiliary regression,

$$\hat{u}_t = \alpha_0 + \alpha_1' W_t + \beta_1' W_t(z_{t-d}) + \beta_2' W_t(z_{t-d})^2 + \beta_3' W_t(z_{t-d})^3 + v_t \tag{8}$$

where $\hat{u}_t$ is the residual obtained from the regression $\Delta y_t = \pi_0 + \pi_1' W_t + u_t$ under the null hypothesis of linearity. The LM-type test of the linearity against STECM (both LSTECM and ESTECM) is to compute the statistic

$$LM_0 = \frac{(SSR_0 - SSR_1)/(3m)}{SSR_1/(T - 4m - 1)} \tag{9}$$

where SSR_0 is the sum of the squared residuals $\hat{u}_t$ and SSR_1 is the sum of the squared residuals $\hat{v}_t$ obtained from Equation 8. The statistic has an asymptotic F distribution with $3m$ and $T - 4m - 1$ degrees of freedom under the null hypothesis of linearity.

Through a sequence of tests on Equation 6 or, equivalently, Equation 8, one can distinguish LSTECM from ESTECM. Consider a sequence of null hypotheses as follows:

$$H_{03} : \beta_3 = 0 \tag{10}$$

$$H_{02} : \beta_2 = 0 \mid \beta_3 = 0, \text{and} \tag{11}$$

$$H_{01} : \beta_1 = 0 \mid \beta_2 = \beta_3 = 0 \tag{12}$$

If the model is a LSTECM, then $\beta_3 \neq 0$ since the transition function $F(z_{t-d}; r, c)$ is an odd function. If the model is an ESTECM, then $\beta_3 = 0$ but $\beta_2 \neq 0$ since $F(z_{t-d}; r, c)$ is an even function. Therefore, a rejection of the null hypothesis H_{03} confirms the model as a LSTECM. Likewise, one would select an ESTECM if the test results accept H_{03}

[6] An application of this method appears in Teräsvirta and Anderson (1992).

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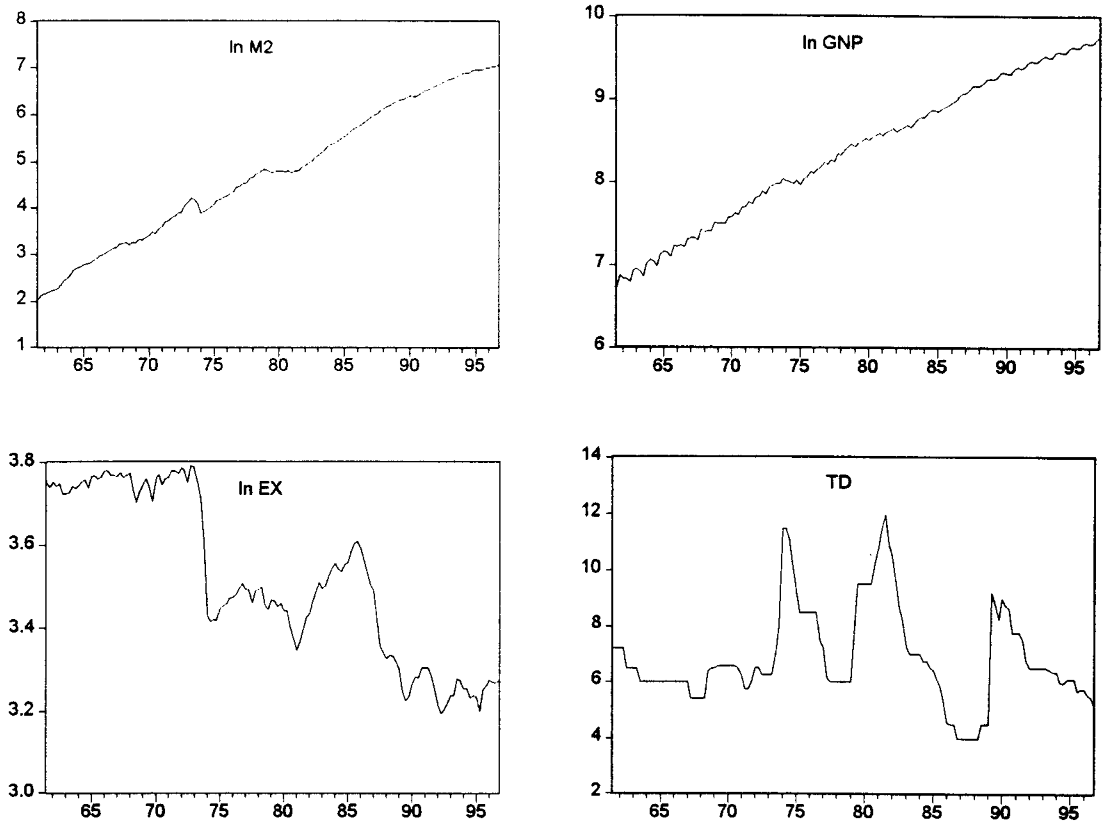

Fig. 1. *Logarithmic values of real M2, GNP, real exchange rate (EX), and time deposit rate (TD)*

and reject H_{02}. When the test results accept both H_{03} and H_{02} but reject H_{01}, a LSTECM is indicated.

III. DATA DESCRIPTION AND THE COINTEGRATION TEST

Quarterly data series from 1962:1 to 1996:4 are used. These data were obtained from the annual publications of the Directorate-General of Budget, Accounting, and Statistics office in Taiwan. To avoid possible distortions of the dynamic properties of the model, this paper uses seasonally unadjusted data. Variables included in the model are the natural logarithm of the real M2 money balance ($\ln M2$), the natural logarithm of real GNP ($\ln GNP$), the three months time deposit rate (TD), and the natural logarithm of the real exchange rate ($\ln EX$).

Figure 1 presents graphs of the data series, which appear to be nonstationary and exhibit rather different pattern. While the real M2 and the real GNP series show steady growth over the last three decades, the real exchange rate and time deposit interest rate displays significant variation over times. In the 1960s and early 1970s, due to the nature of government intervention, the real exchange rate and interest rate experienced less fluctuation. Significant changes of the data series occurred in 1973 and 1979, the years that Taiwan, as well as many other imported oil-dependent economies, suffered from OPEC's oil price hike first and then from the oil embargos on Iran. The Taiwanese government undertook tight monetary policy that caused the real M2 to decline, which was followed by a recession indicated in the real GNP. In the financial markets, the interest rate rose sharply and the new Taiwan dollar appreciated against the US dollar in real term. The waning interest rate in the 1990s was associated with a series of financial deregulation and liberalization that started in the late 1980s. During this period, the foreign exchange market was much less volatile compared to the two previous decades.

To test for the unit roots, the augmented Dickey–Fuller (ADF) test and the Phillips–Perron (PP) test are performed for each of the variables. The results shown in Table 1 accept the null hypothesis of unit roots for each variable at the 5% significance level and confirm that each variable is nonstationary and is integrated of order one, $I(1)$.

To model the long-run demand for money in Taiwan, the classical specification is followed as in Ericsson *et al.* (1998), and it is assumed that demand for real balance is a function of real income and the interest rate. To confirm the existence of a cointegrating relationship, $\ln M2 = f(\ln GNP, TD)$, of the $I(1)$ series, the paper applies the Johansen's (1995) maximum likelihood test. The

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Table 1. *Unit root tests*

		Levels			First-difference	
Variable	lags	ADF	PP	lags	ADF	PP
Unit root with drift						
ln *M2*	2	−1.0085	−1.2376	2	−6.6783*	−11.4991*
	4	−1.1603	−1.2541	4	−5.8704*	−11.4677*
ln *GNP*	2	−1.5424	−3.2827	2	−18.9636*	−17.8346*
	4	−1.7419	−1.3574	4	−4.2125*	−17.8186*
ln *EX*	2	−1.2454	−1.0029	2	−5.6455*	−7.8384*
	4	−1.1317	−1.0689	4	−4.9723*	−7.8255*
TD	2	−2.9420	−2.8262	2	−5.7339*	−9.5459*
	4	−2.9252	−2.8891	4	−5.2181*	−9.5195*
Unit root with drift and trend						
ln *M2*	2	−3.4446**	−3.0163	2	−6.7968*	−11.5482*
	4	−3.2112	−3.0231	4	−6.0320*	−11.4805*
ln *GNP*	2	−1.6052	−3.5035**	2	−19.5483*	−17.4256*
	4	−1.8283	−5.1044*	4	−4.5166*	−17.4141*
ln *EX*	2	−2.9129	−2.4144	2	−5.6244*	−6.3487*
	4	−2.8063	−2.5698	4	−4.9555*	−6.0599*
TD	2	−2.9354	−2.8257	2	−5.7180*	−9.5164*
	4	−2.9169	−2.8872	4	−5.2124*	−9.4875*

Notes: *, ** denote significant at 1 and 5% levels, respectively.

Table 2. *Johansen ML-tests on cointegrating equations* (*CE*) *with various model specifications*

Data		no trend	no trend	linear	linear	quadratic
CE		no intercept no trend	intercept no trend	intercept no trend	intercept trend	intercept trend
*H*o:	LR	50.6644**	62.8978**	29.7437*	40.4939	31.6459
$r = 0$	LV	583.2268	583.2268	599.8039	599.8039	604.2279
	AIC	−7.9887	−7.9887	−8.1869	−8.1869	−8.2077
*H*o:	LR	8.4375	14.9537	8.7237	8.9062	4.5675
$r \leqslant 1$	LV	**604.3403**	**607.1989**	**610.3139**	615.5978	617.7671
	AIC	**−8.2093**	**−8.2365**	**−8.2528**	−8.3153	−8.3178
*H*o:	LR	0.5709	4.3701	3.6685	4.1464	0.4205
$r \leqslant 2$	LV	608.2735	612.4907	612.6915	617.9776	619.8406
	AIC	−8.1792	−8.2115	−8.1999	−8.2478	−8.2604

Notes: *H*o denotes the null hypothesis tested by the likelihood ratio, LR, tests on the number of cointegrating vectors. *, ** denotes significant at 1 and 5% levels, respectively. LV is the log likelihood values and AIC is the Akaike information criterion of the corresponding model.

Johansen test for cointegration uses three seasonal dummies, S1 (first quarter), S2 (second quarter), and S3 (third quarter), as exogenous variables. Also included as exogenous are two structural change dummies, D73 and D79, which represent two unusual breaks in the data series that occurred in 1973:3 to 1974:1, and 1979:1 to 1979:3, possibly a consequence of oil shocks. The results of the Johansen tests, with various assumptions on drift, trend, and the cointegrating equations, are given in Table 2. The likelihood ratio tests reject the null hypothesis of no cointegration, $H_0 : r = 0$, but accept the hypothesis of at most one cointegration, $H_0 : r \leqslant 1$, when the co-integrating relation, $\ln M2 = f(\ln GNP, TD)$, is specified without a trend. Based on the Akaike information criterion ($\mathrm{AIC} = -8.2528$) and the log likelihood value ($\mathrm{LV} = 610.3139$), a long-run cointegrating equation is estimated as follows:

$$(\ln M2)_t = -8.0385 + \underset{(0.0293)}{1.5960}\,(\ln GNP)_t - \underset{(0.0163)}{0.0821}\,(TD)_t + z_t \tag{13}$$

The standard errors are given in parentheses. The results indicate that both slope coefficients are statistically significant at one percent level. The positive income

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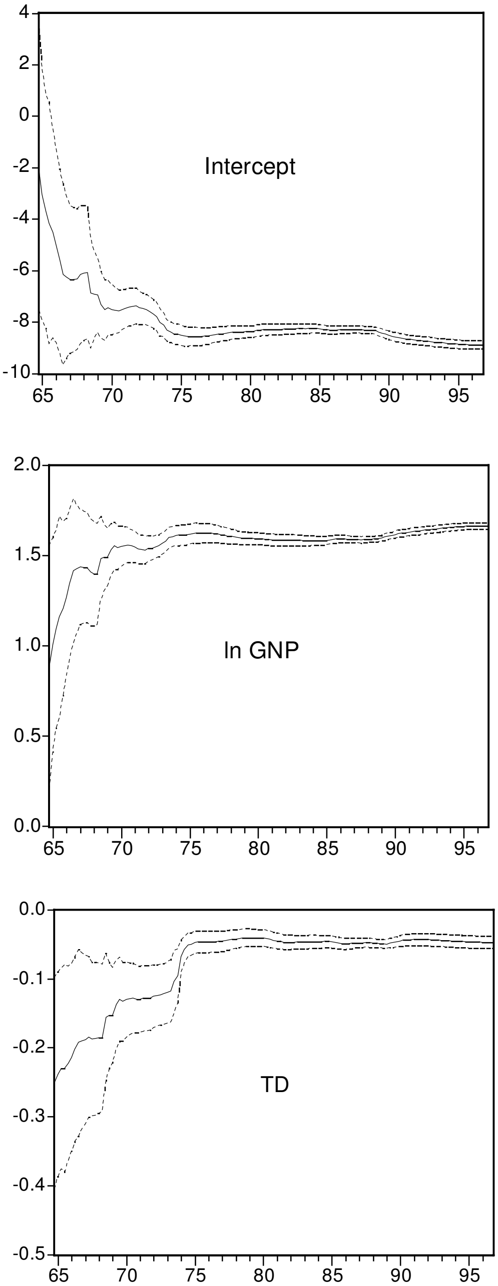

Fig. 2. *Recursive estimates and two-standard error band of the cointegrating vector*

elasticity is consistent with the theory of money demand. The coefficient estimate, 1.5960, indicates that money in Taiwan has no economies of scale, a contradiction to findings of other research on industrial countries.[7] The coefficient of TD confirms the traditional view of a negative relationship between the real balance and the interest rate.

The cointegrating relation in Equation 13 implicitly assumes that the parameters of income and interest rate are constant. Testing the assumption of parameter constancy is indeed an important step in checking the adequacy of a stable long-run demand for money. Figure 2 graphs the recursive estimates of the parameters in Equation 13 and the two-standard error bands.[8] The recursive estimates appear visibly constant, except in the first few years where the number of observations is small.

IV. TESTING FOR THE LINEARITY AND ESTIMATION OF NONLINEAR STECM

Given the estimate of the long-run demand for money in Equation 13, the Akaike information and Schwarz criteria were used to select the short-run dynamic model and the orders of autoregression. The linear ECM with AR(2) is specified as follows with the estimates given in Table 3:

$$\begin{aligned}\Delta(\ln M2)_t = {} & \pi_0 + \pi_1 z_{t-1} + \pi_2 \Delta(\ln M2)_{t-1} \\ & + \pi_3 \Delta(\ln M2)_{t-2} + \pi_4 \Delta(\ln GNP)_{t-2} \\ & + \pi_5 \Delta(\ln EX)_t + \pi_6 \Delta(\ln EX)_{t-1} + \pi_7 (S1)_t \\ & + \pi_8 (S2)_t + \pi_9 (S3)_t \\ & + \pi_{10} (D73)_t + \pi_{11} (D79)_t + u_t \end{aligned} \tag{14}$$

The short-run ECM includes the changes in real exchange rate $\Delta(\ln EX)$. To compare the policy effectiveness for an open economy under different exchange rate systems, Mundell (1963) first proposed that the demand for money is likely to depend on the exchange rate, in addition to the income and the interest rate. He argued that government actions of sterilization and regulation in the foreign exchange markets would change the amount of foreign reserve and therefore the high-power money. His point is most important for an economy such as Taiwan, where the sum of exports and imports accounts for nearly 85% of its national gross product and has been consistently on the list of the top three countries holding the largest amount of foreign reserves.[9] Even though economists disagree on the specific mechanism through which the exchange rate could influence the demand for money, evidence from previous

[7] See Goldfeld (1976) and Hendry and Ericsson (1991) for a review.

[8] The recursive estimates are based on Engle–Granger least-square regression rather than the Johansen's maximum likelihood vector autoregression. In principle, the recursive estimates could be obtained from the Johansen's method. However, the computation and the statistical properties are neither recursive nor known. The Engle–Granger recursive estimates and Fig. 2 are intended for a first approximation only.

[9] *Taiwan Statistical Data Book* (various issues) published by the Council for Economic Planning and Development, Taiwan.

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Table 3. *Estimates of linear ECM and STECM*

Variable	Coefficient	ECM Estimate		LSTECM Estimate	
constant	π_0	0.0426	(0.0049)*	0.0300	(0.0052)*
z_{t-1}	π_1	-0.0480	(0.0170)*	-0.1175	(0.0370)*
$\Delta(\ln M2)_{t-1}$	π_2	0.1750	(0.0748)*	0.1175	(0.0724)
$\Delta(\ln M2)_{t-2}$	π_3	0.2007	(0.0560)*	0.4473	(0.0679)*
$\Delta(\ln GNP)_{t-2}$	π_4	0.1311	(0.0567)*	0.1093	(0.0566)**
$\Delta(\ln EX)_t$	π_5	0.7294	(0.0901)*	0.6554	(0.0798)*
$\Delta(\ln EX)_{t-1}$	π_6	-0.2899	(0.1002)*	-0.0126	(0.1053)
$(S1)_t$	π_7	-0.0319	(0.0065)*	-0.0290	(0.0060)*
$(S2)_t$	π_8	-0.0161	(0.0058)*	-0.0205	(0.0052)*
$(S3)_t$	π_9	-0.0294	(0.0062)*	-0.0295	(0.0056)*
$(D73)_t$	π_{10}	-0.0744	(0.0164)*	-0.0693	(0.0148)*
$(D79)_t$	π_{11}	-0.0559	(0.0136)*	-0.0565	(0.0118)*
z_{t-1}	δ_1			0.0135	(0.0616)
$\Delta(\ln M2)_{t-1}$	δ_2			0.7523	(0.3941)**
$\Delta(\ln M2)_{t-2}$	δ_3			-0.5129	(0.2217)*
$\Delta(\ln GNP)_{t-2}$	δ_4			0.4579	(0.3059)
$\Delta(TD)_{t-3}$	δ_5			-0.0151	(0.0068)*
$\Delta(\ln CPI)_{t-1}$	δ_6			2.0342	(0.8817)*
$F(z_{t-d}; r, c)$	r			3.4171	(1.7321)**
	c			0.1308	(0.0313)*
Adjusted R^2		0.7206		0.8063	
SE		0.0223		0.0193	
AIC		-4.6780		-4.9217	
JB		5.5919	[0.0611]	8.5672	[0.0138]
BG(2)		14.4850	[0.0007]	0.2248	[0.8937]
ARCH(1)		0.7855	[0.3755]	0.0037	[0.9513]
ARCH(4)		2.8192	[0.5885]	1.0914	[0.8956]
RESET(2)		7.7435	[0.0208]	1.1242	[0.5700]

Notes: Asymptotic standard errors are in parentheses and * denotes significance at the 1% level and ** at the 5% level. The standard error of the regression is given in the SE. AIC is the Akaike information criterion value. Statistic JB is the Jarque–Bera test of normality for the error; BG(2) is the Breusch–Godfrey serial correlation LM $\chi^2(2)$ test; ARCH(1) and ARCH(4) are the autoregressive conditional heteroscedasticity χ^2 test of order one and four, respectively; RESET(2) is the Ramsey specification χ^2 test of second order. The p-value of the test are given in brackets.

research supports the inclusion of the exchange rate in the specification of a money demand function.[10]

The linear ECM estimates shown in Table 3 appear to be reasonable with expected signs and are significant at 1% level. The negative coefficient of the error-correction term reconfirms that the short-run adjustment moves the demand for money towards the long-run equilibrium. However, there are a few problems with this specification. First, the Breusch–Godfrey serial correlation LM test, $\text{BG}(2) = 14.4850$, points to the existence of a strong serial correlation. Second, the results of Ramsey's RESET test, $\text{RESET}(2) = 7.7435$, rejects the null hypothesis of no misspecification in the linear ECM model. It is reasonable to suspect that nonlinearity in the error-correction term could have caused these results.

Because results from the linear ECM model are inadequate, it is important to test the linearity of the error-correction model using the STECM specification in Equation 3. The delay parameter, d, in the STECM model is determined by the minimum p-value rule suggested by Teräsvirta (1994). If $p(d)$ is the p-value of the LM-type test on the linearity hypothesis, $H_0 : \beta_1 = \beta_2 = \beta_3 = 0$, then d is chosen so that $p(d)$ is the minimum. Table 4 shows the p-values of the LM-type test for various delay parameter values. The minimum p-value rule yields $d = 1$ with the p-value of 0.0031. The p-values

[10] For instance, while Logue and Willet (1974) looked at how the exchange rate influences capital flow, Arango and Nadiri (1981) examined how it affects investors' portfolio. Bahmani–Oskooee and Pourheydarian (1990) found significant effect of the real exchange rate on the money demand in Canada, Japan, and the United States.

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Table 4. *p-values of the LM-type tests for linearity and sequential STECM hypotheses*

Null	$d=1$	$d=2$	$d=3$	$d=4$	$d=5$
H_0	0.0031	0.1148	0.5728	0.6567	0.5646
H_{03}	0.8404	0.1815	0.4634	0.7021	0.6821
H_{02}	0.1984	0.2195	0.2763	0.5328	0.6101
H_{01}	0.0001	0.1086	0.5766	0.2379	0.1288

on testing $H_{03}: \beta_3 = 0$, $H_{02}: \beta_2 = 0 \mid \beta_3 = 0$, and $H_{01}: \beta_1 = 0 \mid \beta_2 = \beta_3 = 0$ are also tabulated in Table 4. The minimum p-value rule accepts the hypotheses H_{03} and H_{02}, but rejects the hypothesis H_{01} for the delay parameter value $d = 1$. These results imply that the linear ECM model should be rejected, and the STECM model is of logistic type; in other words, it is a LSTECM model.

The next step is an estimation of the LSTECM model using the nonlinear least-squares method. Following the suggestion of Teräsvirta (1994), the exponent of $F(z_{t-1}; r, c)$ is standardized by dividing it by the sample standard deviation $\sigma_z = 0.1269$. Two variables, $\Delta(\text{TD})_{t-3}$ and $\Delta(\ln \text{CPI})_{t-1}$, have been added to the equation to catch the effect of the nonlinearity. The government in Taiwan has closely monitored and regulated the changes of the 3-month time deposit rate and the consumer price. Since the government may take into account the value of z_{t-1} and the economic condition when setting target for changes in the real balance or taking actions to intervene the market, $\Delta(TD)$ could have only sporadic and asymmetric effects on $\Delta(\ln M2)_t$. When $F(z; r, c) = 0$, i.e., in the lower regime, neither $\Delta(TD)$ nor $\Delta(\ln CPI)$ enters the short-run demand equation. On the other hand, when $F(z; r, c) = 1$, i.e., in the upper regime, $\Delta(TD)$ and $\Delta(\ln CPI)$ have the largest impact on $\Delta(\ln M2)$. The LSTECM model is specified as follows:

$$\begin{aligned}
\Delta(\ln M2)_t = {} & \pi_0 + \pi_1 z_{t-1} + \pi_2 \Delta(\ln M2)_{t-1} + \pi_3 \Delta(\ln M2)_{t-2} \\
& + \pi_4 \Delta(\ln GNP)_{t-2} + \pi_5 \Delta(\ln EX)t \\
& + \pi_6 \Delta(\ln EX)_{t-1} + \pi_7 (S1)_t + \pi_8 (S2)_t \\
& + \pi_9 (S3)_t + \pi_{10} (D73)_t + \pi_{11} (D79)_t \\
& + \{\delta_1 z_{t-1} + \delta_2 \Delta(\ln M2)_{t-1} + \delta_3 \Delta(\ln M2)_{t-2} \\
& + \delta_4 \Delta(\ln GNP)_{t-2} + \delta_5 \Delta(TD)_{t-3} \\
& + \delta_6 \Delta(\ln CPI)_{t-1}\} \\
& \times \{1 + \exp[-r(z_{t-1} - c)/\sigma_z]\}^{-1} + u_t \qquad (15)
\end{aligned}$$

Table 3 shows the estimates of the LSTECM model. The results from the diagnostic tests on serial correlation, the ARCH effect, and the RESET model specification all support the LSTECM model. Figure 3 graphs the series $\Delta(\ln M2)_t$ and the predicted residuals u_t from the linear ECM model and the LSTECM model. It appears that the LSTECM model outperforms the linear ECM model in predicting the short-run fluctuation of real M2, especially during the oil shock periods in 1973 and 1979. The adjusted-R^2 further indicates that the LSTECM model explains about 8% more of the overall variation of the real M2 than the linear ECM model.

The transition coefficient $r = 3.4171$ is statistically significant at the 5% level, which reconfirms a nonlinear STECM model for the money demand function in Taiwan. The following estimated logistic transitional function is illustrated in Fig. 4:

$$F(z_{t-1}; r, c) = \{1 + \exp[-3.4171(z_{t-1} - 0.1308)/0.1269]\}^{-1}$$

The function shows that the transition from one regime, $F(z_t; r, c) = 0$, to another, $F(z_t; r, c) = 1$, is almost instantaneous at the thresholds of $z_t = 0$ and 0.2, i.e., at no growth and 20% growth in the real M2. The short-run dynamic is asymmetric and the dynamic switching occurs mainly during the expansion period, when the real M2 is about 13.08% $(c = 0.1308)$ above its long-run equilibrium level. The short-run dynamic demand for M2 reaches the lower regime when $F(z_t; r, c) = 0$:

$$\begin{aligned}
\Delta(\ln M2)_t = {} & 0.0300 - 0.1175 z_{t-1} + 0.1175 \Delta(\ln M2)_{t-1} \\
& + 0.4473 \Delta(\ln M2)_{t-2} \\
& + 0.1093 \Delta(\ln GNP)_{t-2} + 0.6554 \Delta(\ln EX) \\
& - 0.0126 \Delta(\ln EX)_{t-1} \\
& - 0.0290(S1)_t - 0.0205(S2)_t - 0.0295(S3)_t \\
& - 0.0693(D73)_t - 0.0565(D79)_t
\end{aligned}$$

and it reaches the upper regime when $F(z_t; r, c) = 1$:

$$\begin{aligned}
\Delta(\ln M2)_t = {} & 0.0300 - 0.1040 z_{t-1} + 0.8698 \Delta(\ln M2)_{t-1} \\
& - 0.0656 \Delta(\ln M2)_{t-2} + 0.5672 \Delta(\ln GNP)_{t-2} \\
& + 0.6554 \Delta(\ln EX) - 0.0126 \Delta(\ln EX)_{t-1} \\
& - 0.0151 \Delta(TD)_{t-3} + 2.0342 \Delta(\ln CPI)_{t-1} \\
& - 0.0290(S1)_t - 0.0205(S2)_t - 0.0295(S3)_t \\
& - 0.0693(D73)_t - 0.0565(D79)t
\end{aligned}$$

The coefficients of z_{t-1}, $(-0.1175$ and $-0.1040)$, indicate that the short-run error correction is slightly milder at the upper regime than at the lower regime.

The dominant root of the characteristic polynomial for the lower region is real with the value 0.7301, and is 0.7864

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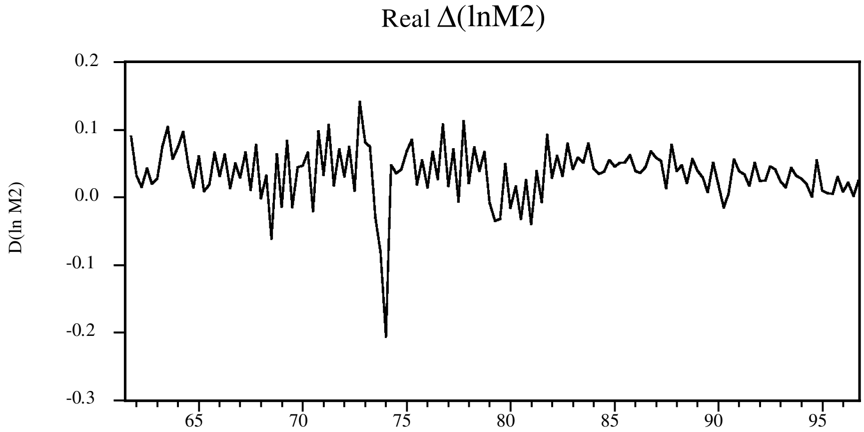

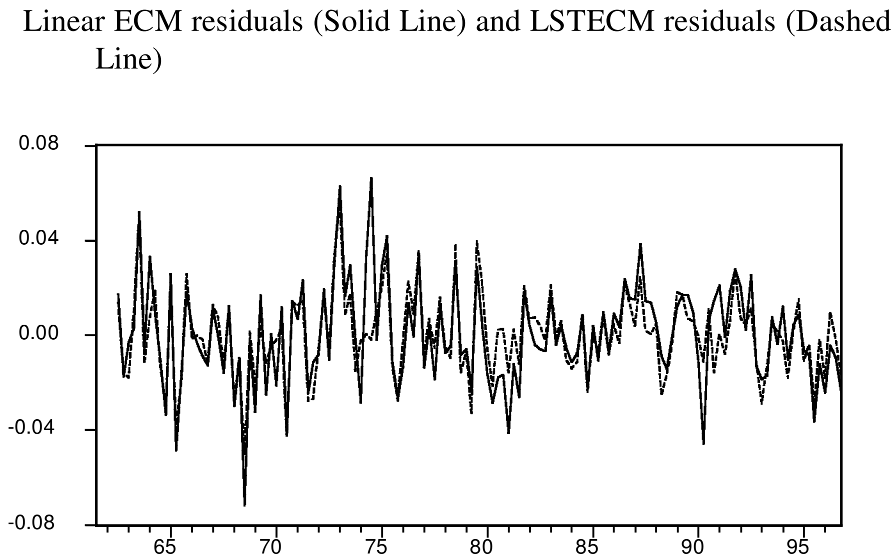

Fig. 3. *The real* Δ*(lnM2) and the predicted residuals from ECM and LSTECM*

for the upper region.[11] Since the roots of the lower and upper regions are less than unity, the short-run dynamic demand for M2 in Taiwan is stable.

V. CONCLUSION

The results of this paper support the idea that a nonlinear error-correction mechanism is needed, even for a newly industrialized country such as Taiwan, when modelling the demand for money. Using data from the first quarter of 1962 to the fourth quarter of 1996, this paper found that the short-run dynamics towards the long-run equilibrium of the money demand in Taiwan follow a nonlinear path. The nonlinear path has been characterized by a logistic smooth transition process proposed by Granger and Teräsvirta (1993) and Teräsvirta (1994). The transition from one state to another is smooth but asymmetric. It implies that, within the context of money demand, households and the government may respond differently when the economy is in a different regime. Although the linear ECM specification gives similar estimation on parameters such as income elasticity and interest rate sensitivity, the nonlinear specification outperforms the linear ECM when judged by such diagnostic tests as serial correlation, the ARCH effect, Ramsey's test, and the overall adjusted-R^2. The nonlinear ECM is especially superior to the linear ECM when the country experiences a volatile economic condition as Taiwan experienced during the two oil shocks.

[11] The characteristic roots for the lower regime are $(-0.6126, 0.7301)$ and are $(0.0834, 0.7864)$ for the upper regime.

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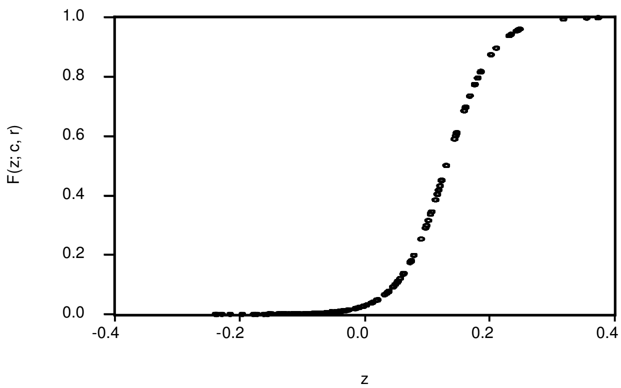

Fig. 4. *The estimated logistic transitional function F(z; r, c)*

This paper has added to the literature a study on money demand in Taiwan and calls for the need to revise the traditional linear specification on dynamics. This is certainly an important issue in monetary economics. While the stable long-run relationship between the real demand for money and other macroeconomic variables serves as the guideline for macro policy, the short-run dynamics is crucial to policy-makers in determining the timing and extent of intervention. Since this paper is the first attempt to model a nonlinear STECM model for the demand for money in Taiwan, further study could extend the framework to include other monetary assets and to explore other forms of nonlinearity.

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