行政院國家科學委員會專題研究計畫 成果報告

台灣貨幣需求之非線性動態研究

計畫類別：個別型計畫
計畫編號：
執行期間：年 月 日至 年 月 日
執行單位：國立中山大學經濟研究所

計畫主持人：吳致寧
計畫參與人員：胡育豪

報告類型：精簡報告
報告附件：出席國際會議研究心得報告及發表論文
處理方式：本計畫可公開查詢

中華民國 年 月 日
Currency Substitution and Non-Linear Error Correction in Taiwan’s Demand for Broad Money

Jyh-Lin Wu
Institute of Economics, National Sun Yat-Sen University, Kaohsiung, Taiwan
Department of Economics, National Chung-Cheng University, Chia-Yi, Taiwan

Abstract

We modify the conventional money demand function by including a real exchange rate variable to reflect the effect of currency substitution. Empirical evidence indicates that the variable is crucial to the long-run stability of Taiwan’s money demand. After finding the failure of a linear error-correction model in describing the dynamics of Taiwan’s money demand, we apply a non-linear error-correction model to examine its dynamics and support the appropriateness of the non-linear model empirically.

JEL: E41, C22

Keywords: Money Demand; Currency Substitution; Cointegration; Non-linear Error-correction;

* Corresponding author. Email: ecdjlw@ccu.edu.tw
1. Introduction

A stable money demand function is important both theoretically and empirically. From the theoretical point of view, it is a significant assumption for several models in macroeconomics (Friedman, 1956; Dornbusch, 1976; Mankiw, 1991). On the other hand, a long-run stable money demand function is also important for central banks in conducting their monetary policy. The impact of monetary expansion on income, interest rates, and prices is predictable ex-ante only when stable long-run money demand exists.

There is no consensus on the existence of a long-run stable money demand function empirically when a linear framework is applied (Hendry and Ericsson, 1991; Hoffman and Rasche, 1991; Baba et al., 1992; Stock and Watson, 1993; Haug and Lucas, 1996). The restriction of a linear cointegration, or the ECM approach, is that it assumes a constant speed of adjustment, which may not be appropriate when there exist transaction costs as suggested by Miller and Orr (1966). In this case deviations from an equilibrium money balance will move toward equilibrium only when the deviations exceed the band threshold, reflecting that the benefits of adjustment surpass the costs, or else the deviations will fluctuate randomly.

The linear ECM is also inconsistent with the buffer stock models of money demand, which recognize non-zero costs of money balance adjustments and imply that the short-run adjustments of the real balance occur only if the balance deviates from its long-run equilibrium (Cuthbertson and Taylor, 1987). These models show that the speed of adjustment for the real balances depends on the size of the deviation. Many articles have adopted a non-linear approach to examining the dynamics of money demand (Sarno et al., 2003; Sarno, 1999; Teräsvirta and Eliasson, 2001; and Huang et al., 2001), and all of these papers provide a non-linear analysis to the money
demand for a single country. Their findings point out that a non-linear model, such as a smooth transition autoregressive model (STAR), is appropriate to discuss the dynamics of money demand.

Although these previously mentioned papers apply a non-linear STAR model to discuss the dynamics of money demand, they ignore the impact of exchange rates on money demand. In fact, including exchange rates in the money demand function is important, and there are two competing forces predicting two different outcomes. First, depreciation increases the domestic price of foreign securities, generating portfolio adjustment effects, which lead to an increase in demand for the domestic currency (Arango and Nadiri, 1981). Second, currency substitution is another link between exchange rates and money demand. According to the currency substitution literature, when the domestic currency is expected to depreciate, the expected return from holding foreign money increases and the demand for the domestic currency falls. If the domestic currency’s depreciation results in the expectation of its future depreciation (appreciation), then the demand for money will decrease (increase). McKinnon (1982) points out that “national monies are substitutable to the extent of making national money demand appear quite unstable if foreign exchange considerations are ignored.” Several articles have found some success with real exchange rates to proxy for international monetary influences on domestic monetary holdings (Arize and Shwiff, 1993; Bahmani-Oskooee, 1991). We therefore examine the long-run stability of money demand in Taiwan and investigate its dynamic adjustments, taking real exchange rates into account.

The purposes of the paper are to address the following issues empirically: Is the real exchange rate variable crucial to the long-run stability of Taiwan’s money demand? Does currency substitution have significant effects on Taiwan’s real
money demand? Is a linear or non-linear error-correction model (ECM) appropriate in describing the dynamics of Taiwan’s money demand?

To examine the long-run stability of money demand, the cointegration test suggested in Johansen (1991) is conventionally applied, but Johansen’s methods may not be efficient if some variables under consideration are weakly exogenous. To test for cointegration, we thus apply the methods suggested by Pesaran et al. (2000), which are efficient when some variables under consideration are exogenous. Given the fact that variables are cointegrated and some variables are weakly exogenous, we set up a linear ECM for money demand and then examine the appropriateness of the model through several diagnostic tests on residuals. After finding evidence of a misspecification of the linear ECM, we then formulate a non-linear ECM to describe the dynamics of Taiwan’s money demand.

Our empirical results point out a significant misspecification of the money demand function when the real exchange rate is omitted from the function. We also find evidence to support the dominant effect of currency substitution on real money demand. Finally, our empirical evidence indicates that the adjustment coefficient of the error-correction term is significant only when the deviation from equilibrium money demand is large. This finding shows that a non-linear ECM is appropriate relative to a linear ECM in describing the dynamic behavior of money demand in Taiwan.

The organization of the paper is as follows. In section 2 we provide the theoretical background of our specification of the money demand function. Section 3 provides the empirical investigation, in which unit-root tests, cointegration tests, and the estimation of a linear ECM and non-linear ECM are performed, respectively. Section 4 concludes the paper.
2. Theoretical background of a stable long-run money demand

The conventional money demand function describes the relationship among real money balances, real income, and nominal interest rates, which can be presented as follows:

\[ m_t = c + \alpha_1 y_t + \alpha_2 r_t + v_t, \]  

(1)

where \( m \) is the log of real money balance; \( y \) is the log of real income; \( r \) is the interest rate; \( v \) is a stationary error term; \( \alpha_1 \) is the income elasticity of money demand; and \( \alpha_2 \) is the interest semi-elasticity of money demand. According to quantity theory, money demand is assumed to be an increasing function of real income and a decreasing function to the opportunity cost of holding money - that is, \( \alpha_1 \) and \( \alpha_2 \) are expected to be positive and negative, respectively. The long-run stability of the money demand function is then examined by testing the existence of cointegration among variables in equation (1) with the expected sign on the cointegrating coefficients.

Although equation (1) may be appropriate to describe money demand in a closed economy, it is unlikely to be adequate for a small open economy with a significant foreign sector. Given the openness of these economies, the conventional specification in equation (1) ignores the effect of real exchange rates and foreign interest rates on money demand. In his seminal paper, Mundell (1963) states, “The demand for money is likely to depend on the exchange rate in addition to the interest rate and the level of income.” He argues that the sensitivity of demand for money to the exchange rate level plays an important role in determining the relative effectiveness of both fiscal and monetary policies. This argument also appears in Bahmani-Oskooee (1991) and Bahmani-Oskooee and Techaratanchai (2001).

An increase in foreign interest rates ceteris paribus raises the return of foreign
assets. Asset holders will therefore increase their holdings of foreign assets by reducing their demand towards domestic assets, including domestic money (i.e. the capital mobility effect). Therefore, the impact of foreign interest rates on real money demand is negative. The depreciation of domestic currency itself raises the domestic prices of foreign assets which in turn increase the demand for domestic assets, including domestic money (i.e. the portfolio adjustment effect). On the other hand, the depreciation of the domestic currency develops expectations for further depreciation, prompting a shift in asset holders’ portfolios away from the domestic currency into foreign assets, including foreign currency (i.e. the currency substitution effect). Hence, the impact of exchange rates on money demand can be either positive or negative.

Globalization has sped up the progress of integration in financial markets as well as the progress of foreign trade liberalization in Taiwan since 1990, making foreign financial assets available to the general population. The availability of foreign financial assets to hedge against risks does affect domestic money demand. It is therefore crucial to take exchange rates and foreign interest rates into account when analyzing the dynamics of money demand. These observations lead us to conjecture the appropriateness of the conventional money demand function for Taiwan. To account for the effects of international monetary developments on the demand for broad money, a real (rather than nominal) exchange rate is included in the money demand function. This is because several articles, such as Bahmani-Oskooee (1991) and Arize and Shwiff (1993), find that the money demand relationship is more pronounced when exchange rates are expressed in real terms.

It has been suggested that foreign interest rates can be included in the money demand function since there could be some degree of substitutability between
domestic and foreign financial assets (Cuddington, 1983). Hamburger (1977, p.31) argues that it is the domestic interest rate that determines money holdings by the public. This is especially true when global financial markets have become highly integrated. For this reason, we exclude the foreign interest rate from our specification, which is also supported empirically by the cointegration analysis, as one can see from the following section. Therefore, the modified money demand equation is given as follows:

\[ m_t = c + \alpha_1 y_t + \alpha_2 r_t + \alpha_3 q_t + v_t, \]  

(2)

where \( q \) is the real exchange rate and \( \alpha_3 \) is the real exchange rate elasticity of money demand, which can be either positive or negative.

The stability of real money demand can be investigated by examining the existence of a long-run equilibrium relationship among variables in equation (2) using cointegration analysis. Supporting the existence of a cointegrating relationship allows us to examine the dynamic adjustments of deviations from the equilibrium. We then discuss a linear or non-linear error-correction model that is appropriate for describing the dynamics of Taiwan’s real money demand.

3. **Empirical analysis**

3.1 **Data description**

The data for Taiwan’s money supply (M2), real gross national product (GNP), consumer price index, one-month time deposit rate, and nominal exchange rate between New Taiwan dollars and US dollars are obtained from the AREMOS database provided by Taiwan’s Ministry of Education. Real money balance is obtained by deflating M2 by the consumer price index. The data for the U.S. consumer price index are obtained from IMF’s international financial statistics. The sample period starts from the first quarter of 1962 and ends in the last quarter of 2003,
depending on the availability of data.

3.2 Unit-root and cointegration tests

We start our empirical analysis by examining the stochastic properties of the time series. We first examine the unit-root hypothesis with the ADF tests. The model applied is one with a constant and a trend, and the lag order is selected by Schwartz’s Bayesian information criteria (SIC). Findings from the second row of Table 1 indicate that the unit-root hypothesis fails to be rejected for level variables, though it is rejected for differenced variables. It is worth noting, however, that the ADF test is not the most powerful test available. We also present results from the DF-GLS test provided by Elliott et al. (1996), using the same lag order selection procedure. Findings from the DF-GLS test are consistent with those from the ADF test, as the third row of Table 1 shows. Finally, we apply a multivariate test suggested by Johansen (1995) to examine the stationarity of variables and find evidence to reject the unit-root hypothesis for all level variables.¹

After finding a unit-root for variables, we first perform Johansen’s cointegration tests. To implement Johansen’s procedure, one needs to determine the lag order of the VAR system and establish whether a deterministic term should be included in the model. The lag order of the VAR model is set to be four, which is determined based on the Schwartz information criteria. Johansen (1995) points out that modeling deterministic components of the cointegrating models has important implications for the asymptotic distribution of the test statistics. We apply the Pantula procedures suggested in Johansen (1995) to determine the appropriateness of the model.

Three different models are considered herein: the model without deterministic

¹ Based on Johansen (1995, p.74), the null hypothesis of stationary money supply, for example, is tested by restricting the cointegrating vector - among the variables m, y, r, and q - to contain all zeros except for a unity corresponding to the money supply.
trends in the data and the cointegrating equations have intercepts (model A), the model with a linear trend in the data and the cointegration equations have intercepts (model B), and the model with linear trends in both data and the cointegrating equations (model C). After estimating the above three models, we examine the hypothesis of existing zero and one cointegrating vector for the models sequentially. Findings from Table 2 indicate that the hypothesis of existing zero cointegrating vector is rejected for all models, and the hypothesis of existing one cointegrating vector is not rejected only for model B. Therefore, model B is applied to examine the long-run relationship among variables of interest in our empirical analysis.  

Findings from Table 3 point out that the hypothesis of existing zero cointegrating vector is rejected at the 10% and 5% levels by trace and maximum-eigenvalue statistics, respectively. Figure 1 plots the estimated cointegrating residuals from the modified money demand function, revealing significant mean-reverting properties. This evidence indicates the existence of a cointegrating vector among real money balance, real income, interest rates, and real exchange rates.  

To examine the appropriateness of excluding the foreign interest rate (the US treasury bill rate) in the money demand equation, we test for the existence of cointegration among real money balances, domestic interest rates, foreign interest

---

2 The p-dimensional vector error-correction specification of model B is given as follows:

$$\Delta y_t = \gamma \mu_t + \alpha \tilde{\beta} \tilde{y}_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i},$$

where $\tilde{\beta} = (\beta, \mu_1), \tilde{y}_t = (y_t, 1).$

3 We also apply the tests provided by Gregory and Hansen (1996) to investigate whether there exists a structural change in the long-run relationship of money demand, and we conclude that the estimated cointegrating relationship is stable during the period under investigation. Results are not reported here, but are available upon request from the authors.
rates, real income, and real exchange rates. The findings from Johansen’s trace and maximum-eigenvalue statistics for the hypothesis of a zero cointegrating vector existing among variables are 52.99 and 27.99, respectively, which are less than their respective 10% critical values. This indicates that it is not appropriate to include a foreign interest rate variable in our augmented money demand equation.

If the real exchange rate variable is excluded from the system, then we find that Johansen’s trace and maximum-eigenvalue statistics for the hypothesis of a zero cointegrating vector existing among variables are 24.27 and 13.17, respectively. These statistics are less than their respective 5% critical values. Therefore, our empirical evidence indicates that there is no cointegration among the variables of real money balances, interest rates, and real income. This finding states that including a real exchange rate variable to reflect the effect of currency substitution is crucial to the long-run stability of money demand in Taiwan.

Johansen’s cointegration methods assume that variables under consideration are all endogenous, but Johansen’s methods may not be efficient if some variables under consideration are weakly exogenous. Pesaran et al. (2000) provide efficient methods for testing cointegration and estimating the VECM when some variables under consideration are exogenous. We hence test the hypothesis of weak exogeneity for variables under consideration by a likelihood ratio statistic and find that the hypothesis is rejected at the 10% level for interest rate and real money balances, respectively, as Table 4 provides. Since the weakly exogenous test fails to reject the hypothesis that real income and real exchange rates are weakly exogenous, we shall examine whether these two variables are jointly, weakly exogenous.

The last column of Table 4 fails to reject the previous joint hypothesis at the 5%

---

4 The U.S. treasury bill rates are obtained from IMF’s international financial statistics.
level of significance. Given the fact that there are two weakly exogenous variables in the four-variable system, we apply the modified trace and maximum-eigenvalue statistics, provided by Pesaran et al. (2000), to re-examine the cointegration among variables. The 5% critical values for the previously-mentioned statistics are obtained from MacKinnon et al. (1999). Findings from Table 5 indicate that there exists a long-run relationship among real money balance, real income, interest rates, and real exchange rates since both the trace and maximum-eigenvalue statistics reject the hypothesis of no cointegration at the 5% level of significance. The estimated long-run relationship is given as follows:

\[ m_t = -0.52 + 1.54y_t - 0.08r_t - 0.74q_t + \varepsilon_t. \] (3)

Following Johansen (1995), we next apply a likelihood ratio test to examine the significance of the estimated cointegrating coefficients and find that the likelihood ratio statistic for the slope coefficients are 18.83, 17.11, and 17.29, respectively, which are all larger than the 5% critical value. Therefore, we find that the estimated cointegrating coefficients are all significantly different from 0. According to the long-run relationship, we find that income elasticity is 1.54 and the interest semi-elasticity is –0.08, which are consistent with those in Arize (1994) and Huang and Lin (2001). Furthermore, the negative impact of real exchange rates on real money balances indicates the dominant effect of currency substitution.

### 3.3 A linear error-correction model

Before examining a non-linear error-correction model, we investigate the appropriateness of a linear ECM. Based on the results of cointegration and weakly exogenous tests, the ECM for money demand is given as follows:  

\[ \Delta m_t = \sum_{j=1}^{p} B_j \Delta m_{t-j} + \sum_{j=1}^{p} \beta_j y_{t-j} + \sum_{j=1}^{p} \gamma_j r_{t-j} + \sum_{j=1}^{p} \delta_j q_{t-j} + \varepsilon_t. \] (4)

5 These critical values depend on the number of variables, exogenous variables, and cointegrating vectors.

6 It is worth noting that \( \Delta r_t \) is not included in equation (4) since \( r_t \) fails to pass the weakly exogenous test.
\[
\Delta m_t = a_0 + a_1S_1 + a_2S_2 + a_3S_3 + b_1z_{t-1} + \sum_{i=0}^{3} b_{ii} \Delta m_{t-i} + \sum_{i=0}^{3} b_{si} \Delta y_{t-i} + \sum_{i=0}^{3} b_{Rt-i} + \sum_{i=0}^{3} b_{d} \Delta q_{t-i} + \nu_t
\]

Here, \( z_t \) is the equilibrium error normalized on \( m_t \), \( b_1 \) is the adjustment coefficient of the equilibrium error, which is expected to be negative, and \( S_i \)'s are centered seasonal dummies.\(^7\)

Not all of the coefficients in the previous equation may be statistically significant in practice, and greater efficiency may actually be gained by removing the insignificant coefficients. We exclude those insignificant variables as long as their elimination does not produce evidence of serial correlation based on a Q-statistic at 8 and 12 lags. Using this procedure, variables are included, even when they are insignificant, if their deletion has resulted in serial correlation. The following parsimonious equation is estimated.

\[
\Delta m_t = a_0 + a_1S_1 + a_2S_2 + a_3S_3 + b_1z_{t-1} + b_{11} \Delta m_{t-1} + b_{12} \Delta m_{t-2} + b_{13} \Delta m_{t-3} + b_{21} \Delta y_{t-1} + b_{22} \Delta y_{t-2} + b_{31} \Delta q_t + b_{32} \Delta q_{t-1} + b_{33} \Delta q_{t-2} + b_{d} \Delta q_{t-3} + \nu_t.
\]

Table 6 reports the estimation results of equation (5). As for diagnostic checking, the Ljung-Box’s Q statistic fails to reject the hypothesis of no autocorrelation in residuals and squared residuals. In addition, the ARCH statistic of Engle (1982) fails to reject the hypothesis of no autoregressive conditional heteroscedasticity in residuals. However, the RESET statistic of Ramsey’s regression specification error test rejects the hypothesis of no misspecification for the linear ECM at the 5% level of significance. The Jarque-Bera (JB) statistic rejects the hypothesis of normality in residuals. Findings from Table 6 indicate that a linear ECM may not be appropriate to describe the dynamics of Taiwan’s money demand.

---

\(^7\) Seasonal dummies are centered in order to ensure that they sum to zero over time.

---

\(^7\) In other words, the current change of a variable is included in an error-correction model if the variable is weakly exogenous (Mckinnon, 1999).
Since the linear ECM is not supported empirically, we conjecture that a non-linear model may be appropriate to describe the dynamics of money demand. To find out if there are non-linearities in the adjustment process of money demand, we include the squared equilibrium error (\(z_{t-2}^2\)) in equation (5) and then estimate the model with non-linear least squares. We discover that the coefficient estimates on the residual (\(z_{t-1}\)) and squared residual (\(z_{t-2}^2\)) are all significant at the 5% level, and there is no evidence for serial correlation and heteroscedasticity in estimated residuals.\(^8\) Our findings hence provide some support for the hypothesis of non-linear adjustment.

### 3.4 A non-linear error-correction model

After taking into account the existence of transaction costs, we consider a smooth transition autoregressive (STAR) error-correction model to analyze the dynamics of money demand in Taiwan. The model is described as follows:

\[
\Delta m_t = (\kappa_1 + \Phi_1 W_t)(I - F(z_{t-4}; \gamma, c)) + (\kappa_2 + \Phi_2 W_t)F(z_{t-4}; \gamma, c) + \varepsilon_t, \tag{6}
\]

where \(W_t = (S_t, S_{t-1}, z_{t-4}, \Delta m_{t-1}, \Delta m_{t-2}, \Delta m_{t-3}, \Delta y_{t-1}, \Delta y_{t-2}, \Delta r_{t-1}, \Delta q_{t-1}, \Delta q_{t-2})'\) is the vector of significant explanatory variables in equation (5); \(F\) is a transition function, which by convention is bounded by zero and one; and \(\varepsilon_t\) is assumed to be an identical, independent and normal distribution with mean zero and a constant variance. Parameters \(\kappa_1\) and \(\kappa_2\) are scalar, and \(\Phi_1\) and \(\Phi_2\) are row vectors of coefficients. Parameter \(\gamma\) in the transition function reflects the speed of transition between regimes, and parameter \(c\) represents the threshold around which the model’s dynamics change.

There are two different transitional functions in general. One is the logistic

\(^8\) Results are not reported here, but are available upon request from the authors.
function, \( F(z_{t-d}; \gamma, c) = \left[ 1 + \exp(-\gamma(z_{t-d} - c)) \right]^{-1} \), and the other one is the exponential function, \( F(z_{t-d}; \gamma, c) = \left[ 1 - \exp(-\gamma(z_{t-d} - c)^2) \right] \). The STAR model can thus be distinguished in the logistic STAR (LSTAR) model and the exponential STAR (ESTAR) model. Since the transition function in the LSTAR model is not symmetric about \( c \), it is useful to describe an economy with an asymmetric transition. On the other hand, an ESTAR model is helpful to model an economy with a symmetric transition since its transition function is symmetric about \( c \). However, there is no theoretical guidance on the choice of an ESTAR or a LSTAR model. As such, we apply a purely statistical approach to select the transition function in our empirical analysis.

Following Granger and Teräsvirta (1993) and Teräsvirta (1998), we first perform linearity tests based on the following auxiliary regression:

\[
\hat{\nu}_t = \zeta_0 + \zeta_1 W_t + \zeta_2 W_t \hat{z}_{t-d} + \zeta_3 W_t \hat{z}_{t-d}^2 + \zeta_4 W_t \hat{z}_{t-d}^3 + \eta_t,
\]

where \( \hat{\nu}_t \) is the estimated residual from a linear ECM. The linearity test examines the hypothesis of \( H_0: \zeta_2 = \zeta_3 = \zeta_4 = 0 \). The Wald test is used to test the hypothesis of linearity (\( H_0 \)) against a STAR-type non-linearity for various values of \( d \). If \( H_0 \) is rejected for more than one value of \( d \), then \( d \) is set to the value of \( \hat{d} \) that minimizes the p-value of the linearity test.

After rejecting \( H_0 \), the choice between LSTAR and ESTAR models is based on testing the following sequential hypothesis:

\[
\begin{align*}
H_{04} & : \zeta_4 = 0 \\
H_{03} & : \zeta_3 = 0 \mid \zeta_4 = 0 \\
H_{02} & : \zeta_2 = 0 \mid \zeta_4 = \zeta_3 = 0 
\end{align*}
\]

If the p-value of the test corresponding to \( H_{03} \) is the smallest, then an ESTAR model
is selected; otherwise, a LSTAR model is selected. Table 7 reports results of the linearity tests. Based on the previously-mentioned procedure, we set \( d = 2 \) and choose an ESTAR model for our empirical analysis.

The ESTAR model can be estimated by non-linear least squares. In practice, the estimation of the smoothing parameter, \( \gamma \), is problematic, because the convergence of \( \gamma \) may be slow in the neighborhood of \( c \). One solution suggested by Granger and Teräsvirta (1993) and Teräsvirta (1994) is to scale the smoothing parameter, \( \gamma \), by the standard deviation of the transition variable. Thus, the transition function is replaced by:

\[
F(z_{t-2}; \gamma, c) = 1 - \exp\left[\frac{\gamma(z_{t-2} - c)^2}{\sigma^2(z_{t-2})}\right].
\]

We estimate the ESTAR model using a non-linear least square method and apply a general-to-specific rule to exclude insignificant explanatory variables. The final parsimonious form of the non-linear ECM is given as follows:

\[
\Delta m_t = [a_0 + a_1S_3 + b_{11}\Delta m_{t-1} + b_{12}\Delta m_{t-2} + b_{13}\Delta y_t + b_{31}\Delta q_1 + b_{32}\Delta q_{t-2} + b_{33}\Delta r_{t-1}] (1 - F(z_{t-2}; \gamma, c)) + [a'_0 + a'S_1 + a'_1 S_2 + a'_2 S_3] + b'_1 z_{t-1} + b'_2 \Delta m_{t-1} + b'_3 \Delta y_{t-2} + b'_4 \Delta q_1 F(z_{t-2}; \gamma, c) + \nu_t.
\]

Table 8 reports the estimation results from the non-linear ECM, revealing several interesting findings. First, all estimates including those in the transition function are significant. The estimate of \( \gamma \) is large and significant, implying a large adjustment speed between regimes. Second, residual diagnostic checks point out that there is no evidence for the ARCH effect in residuals and no evidence for a misspecification of the model. The JB statistic also fails to reject the normality of residuals. Third, if the deviation from equilibrium money demand, \( z_{t-2} \), is extremely large and hence \( F = 1 \), then the dynamic equation for real money demand degenerates to a linear model with a negative adjustment coefficient of the
error-correction term. The adjustment coefficient from dis-equilibrium errors is –0.06 and it is significant. On the other hand, if \( z_{t-2} \) is close to \( c \) and hence \( F=0 \), then the dynamic equation of money demand degenerates to an alternative linear equation with an insignificant adjustment coefficient of dis-equilibrium errors. After comparing findings from Table 8 with those from Table 6, we claim that regime switching does exist in the dynamic error-correction model. It is hence not appropriate to investigate the dynamics of real money demand with a linear model.

To provide formal and rigorous evidence for the predominance of the non-linear ECM (equation 7) relative to the linear ECM (equation 5), we apply a non-nested testing procedure provided by Davidson and Mackinnon (1981).\(^9\) The J-statistic for the hypothesis of the linear ECM against the non-linear ECM is 7.93, which rejects the linear ECM at the 5% level. Moreover, the J-statistic for the hypothesis of the non-linear ECM against the linear ECM is 0.85, and hence it fails to reject the non-linear ECM at the conventional significance level. The above findings show that the adjustment of money demand in Taiwan is predominately non-linear.

Figure 2 plots the estimated transition function against the transition variable, \( \hat{z}_{t-2} \), which can be used to examine the strongly non-linear behavior implied by our non-linear model. The estimated exponential transition function is bounded between 0 and 1 and is symmetrically inverse-bell-shaped around –0.035. The figure shows that the value of the transition function is close to one when \( \hat{z}_{t-2} \) is large. This implies a fast speed of transition between regimes for large deviations from the long-run equilibrium.

---

\(^9\) The J-statistic provided by Davidson and Mackinnon (1981) is asymptotically distributed as \( N(0,1) \) under the null hypothesis. Moosa (1994) applies the J-statistic to test non-linearities in purchasing power parity.
We finally obtain the estimated changes in real money balance from the non-linear ECM model and then plot the estimated changes with those of actual changes in Figure 3. We find that the estimated changes fit very well to actual changes. In addition, the estimated changes match most of the turning points of actual changes.

4. Conclusion

Globalization and technology speed up the integration of economies, making foreign assets available to domestic agents, blurring the distinction between national and international uses of money. Currency substitution therefore opens a channel through which domestic money markets are exposed to shocks occurring abroad. It also raises the unpredictability of money demand and reduces the effectiveness of monetary policy under flexible exchange rates.

The purposes of this paper are to examine the long-run stability of Taiwan’s money demand function and to investigate its dynamics. A modified money demand function, motivated by the literature of currency substitution, is applied in our empirical analysis in which a real exchange rate variable is included in the function. We also pay attention to deciding whether a deterministic trend should be included in the models. In the cointegration analysis we apply the method by Pesaran et al. (2000), which is more efficient than that of Johansen (1995), when some variables in the system are weakly exogenous. Although the presence of real exchange rates is crucial to the long-run stability of Taiwan’s money demand, we find that a linear ECM is misspecified in describing its dynamics. A non-linear ECM is therefore applied to re-examine the dynamics of money demand. We find that the speed of adjustment between regimes in an ESTAR model is large and significant. Moreover, the adjustment coefficient of the error-correction term is significant only when the
deviation from equilibrium money demand is large. Our results lend support to the
dominant effect of currency substitution and to the specification of a non-linear
error-correction in Taiwan’s money demand.

In terms of policy implications, our findings point out that real exchange rates
exert a significant effect on the stability of domestic money demand in Taiwan.
Thus, neglecting this can produce biased results. If adjustments in money demand
induced by real exchange rates are ignored, then monetary policy actions can generate,
at best, only uncertain results.
Table 1. Unit-root tests

<table>
<thead>
<tr>
<th></th>
<th>(m_t)</th>
<th>(y_t)</th>
<th>(r_t)</th>
<th>(q_t)</th>
<th>(\Delta m_t)</th>
<th>(\Delta y_t)</th>
<th>(\Delta r_t)</th>
<th>(\Delta q_t)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\tau_t)</td>
<td>-0.15</td>
<td>-0.08</td>
<td>-1.93</td>
<td>-0.74</td>
<td>-4.63*</td>
<td>-5.09*</td>
<td>-11.81*</td>
<td>-11.55*</td>
</tr>
<tr>
<td>DF-GLS</td>
<td>-1.30</td>
<td>-0.29</td>
<td>-1.98</td>
<td>-0.99</td>
<td>-3.09*</td>
<td>-4.08*</td>
<td>-11.87*</td>
<td>-11.57*</td>
</tr>
<tr>
<td>JOH</td>
<td>16.45*</td>
<td>22.64*</td>
<td>16.56*</td>
<td>23.21*</td>
<td>--</td>
<td>--</td>
<td>--</td>
<td>--</td>
</tr>
</tbody>
</table>

Note: \(\tau_t\) and DF-GLS are unit-root statistics provided by Dicky-Fuller and Elliot et al. (1996). JOH indicates the multivariate test for stationarity of variables provided by Johansen (1995), which has a chi-square distribution with the degree of freedom being 4. Term * indicates significance at the 5% level, and term ‘--’ indicates a statistic is not computed.

Table 2. Model selection for Johansen’s cointegration tests

<table>
<thead>
<tr>
<th></th>
<th>Model A</th>
<th>Model B</th>
<th>Model C</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\omega = 0)</td>
<td>(\text{Tr})</td>
<td>(\lambda_{\text{Max}})</td>
<td>(\text{Tr})</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>63.31*</td>
<td>31.84*</td>
<td>45.36#</td>
</tr>
<tr>
<td></td>
<td>(49.65)</td>
<td>(25.56)</td>
<td>(43.95)</td>
</tr>
<tr>
<td>(\omega = 1)</td>
<td>31.47</td>
<td>23.31*</td>
<td>17.64</td>
</tr>
<tr>
<td></td>
<td>(32.00)</td>
<td>(19.77)</td>
<td>(26.79)</td>
</tr>
</tbody>
</table>

Note: \(\text{Tr}\) and \(\lambda_{\text{Max}}\) are trace and maximum eigenvalue statistics, respectively. Term \(\omega\) is the number of cointegrating vectors. Terms * and # indicate significance at the 5% and 10% levels, respectively. Model A refers to the model without deterministic trends in data, and the cointegrating equations have intercepts. Model B refers to the model with linear trends in data and the cointegration equations have intercepts. Finally, model C refers to the model with linear trends in both data and the cointegrating equations. A number in the parenthesis indicates the 10% critical value.
Table 3. Johansen’s cointegration tests

<table>
<thead>
<tr>
<th></th>
<th>Tr</th>
<th>CV5%</th>
<th>CV10%</th>
<th>$\lambda_{\text{Max}}$</th>
<th>CV5%</th>
<th>CV10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\omega=0$</td>
<td>45.36#</td>
<td>47.21</td>
<td>43.95</td>
<td>27.72*</td>
<td>27.07</td>
<td>24.73</td>
</tr>
<tr>
<td>$\omega=1$</td>
<td>17.64</td>
<td>29.68</td>
<td>26.79</td>
<td>10.10</td>
<td>20.97</td>
<td>18.60</td>
</tr>
<tr>
<td>$\omega=2$</td>
<td>7.54</td>
<td>15.41</td>
<td>13.33</td>
<td>5.65</td>
<td>14.07</td>
<td>12.07</td>
</tr>
<tr>
<td>$\omega=3$</td>
<td>1.89</td>
<td>3.76</td>
<td>2.69</td>
<td>1.89</td>
<td>3.76</td>
<td>2.69</td>
</tr>
</tbody>
</table>

Note: CV5% and CV10% are critical values at the 5% and 10% levels, respectively. Same as in Table 2.

Table 4. Weakly exogenous tests

<table>
<thead>
<tr>
<th></th>
<th>$m_t$</th>
<th>$y_t$</th>
<th>$r_t$</th>
<th>$q_t$</th>
<th>$(y_t, q_t)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>LR</td>
<td>3.16#</td>
<td>0.30</td>
<td>9.45*</td>
<td>0.69</td>
<td>0.70</td>
</tr>
<tr>
<td>p-value</td>
<td>0.075</td>
<td>0.586</td>
<td>0.002</td>
<td>0.406</td>
<td>0.706</td>
</tr>
</tbody>
</table>

Note: LR indicates a likelihood ratio statistic. Same as in Table 2.

Table 5. Pesaran’s cointegration tests

<table>
<thead>
<tr>
<th></th>
<th>Tr</th>
<th>CV5%</th>
<th>$\lambda_{\text{Max}}$</th>
<th>CV5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\omega=0$</td>
<td>35.20*</td>
<td>28.88</td>
<td>27.03*</td>
<td>21.52</td>
</tr>
<tr>
<td>$\omega=1$</td>
<td>8.17</td>
<td>14.39</td>
<td>8.17</td>
<td>14.39</td>
</tr>
</tbody>
</table>

Note: The 5% critical values (CV5%) of Tr and $\lambda_{\text{Max}}$ are obtained from MacKinnon et al., (1999). Same as in Table 2.
Table 6. Results of Linear ECM

\[ \Delta m_t = a_0 + a_1 S_t + a_2 S_{t-1} + a_3 S_{t-2} + b_1 \Delta m_{t-1} + b_2 \Delta m_{t-2} + b_3 \Delta m_{t-3} + b_4 \Delta y_t + b_5 \Delta y_{t-2} + b_6 \Delta q_t + b_7 \Delta q_{t-1} + b_8 \Delta q_{t-2} + b_9 \Delta r_{t-1} + \nu_t \]

<table>
<thead>
<tr>
<th></th>
<th>( b_1 )</th>
<th>( b_{11} )</th>
<th>( b_{12} )</th>
<th>( b_{13} )</th>
<th>( b_{21} )</th>
<th>( b_{22} )</th>
<th>( b_{23} )</th>
<th>( b_{31} )</th>
<th>( b_{32} )</th>
<th>( b_{33} )</th>
<th>( b_{41} )</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>-0.064</td>
<td>0.259</td>
<td>0.181</td>
<td>-0.088</td>
<td>0.210</td>
<td>0.126</td>
<td>0.639</td>
<td>-0.198</td>
<td>-0.142</td>
<td>-0.010</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-2.45)</td>
<td>(3.57)</td>
<td>(2.26)</td>
<td>(-1.612)</td>
<td>(3.78)</td>
<td>(2.52)</td>
<td>(9.01)</td>
<td>(-2.17)</td>
<td>(-1.630)</td>
<td>(-3.30)</td>
<td></td>
</tr>
</tbody>
</table>

ARCH(1)= 1.97 [0.16]; ARCH(4)=7.84 [0.10];
Q(8)=8.44 [0.39]; Q(12)=13.51 [0.33], Q(16)=16.10 [0.45]
Q^2(8) = 8.87 [0.35]; Q^2(12) = 12.95 [0.37]; Q^2(16) = 16.36 [0.43];
JB=15.13 [0.00]  adj \( R^2 \) = 0.72; RESET(1)=4.26 [0.04];

Note: The number in the parenthesis under an estimate is the t statistic. A number in [ ] is a p-value. Q\( (j) \) and Q^2\( (j) \) are Ljung-Box autocorrelation tests for residuals and squared residuals, respectively, which have \( \chi^2 \) distribution with \( j \) degrees of freedom. ARCH\( (p) \) is the autoregressive conditional heteroscedasticity test of Engle (1982) and has \( \chi^2 \) distribution with \( p \) degrees of freedom. RESET is Ramsey’s regression specification error test. It tests for the functional form misspecification and has \( \chi^2 \) distribution with 1 degree of freedom. JB is the Jarque-Bera normality test, which has \( \chi^2 \) distribution with 2 degrees of freedom. \( S_1, S_2, \) and \( S_3 \) in the regression equation are centered seasonal dummies. Term \( z_t \) is the error-correction term, which is normalized with \( m_t \).

Table 7. Linearity test

\[ \hat{\nu}_t = \alpha_0 + \alpha_1 W_t + \alpha_2 W_t z_{t-d} + \alpha_3 W_t z_{t-d}^2 + \alpha_4 W_t z_{t-d}^3 + \eta_t \]

<table>
<thead>
<tr>
<th></th>
<th>( H_0 )</th>
<th>( H_{04} )</th>
<th>( H_{03} )</th>
<th>( H_{02} )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( d=1 )</td>
<td>38.243</td>
<td>4.536</td>
<td>16.400</td>
<td>19.837</td>
</tr>
<tr>
<td></td>
<td>(0.64)</td>
<td>(0.99)</td>
<td>(0.29)</td>
<td>(0.14)</td>
</tr>
<tr>
<td>( d=2 )</td>
<td>55.631#</td>
<td>12.013</td>
<td>27.345*</td>
<td>15.462</td>
</tr>
<tr>
<td></td>
<td>(0.08)</td>
<td>(0.61)</td>
<td>(0.02)</td>
<td>(0.35)</td>
</tr>
<tr>
<td>( d=3 )</td>
<td>47.862</td>
<td>18.781</td>
<td>21.905#</td>
<td>5.702</td>
</tr>
<tr>
<td></td>
<td>(0.25)</td>
<td>(0.17)</td>
<td>(0.08)</td>
<td>(0.97)</td>
</tr>
<tr>
<td>( d=4 )</td>
<td>48.232</td>
<td>17.850</td>
<td>11.982</td>
<td>17.705</td>
</tr>
<tr>
<td></td>
<td>(0.24)</td>
<td>(0.21)</td>
<td>(0.61)</td>
<td>(0.22)</td>
</tr>
</tbody>
</table>

Notes: \( H_0 : \alpha_2 = \alpha_3 = \alpha_4 = 0 \); \( H_{02} : \alpha_2 = 0 | \alpha_4 = \alpha_3 = 0 \); \( H_{03} : \alpha_3 = 0 | \alpha_4 = 0 \); \( H_{04} : \alpha_4 = 0 \). Same as in Table 2.
Table 8. Results of non-linear ECM

\[ \Delta m_t = [a_0 + a_1 s_{t-1} + b_{12} m_{t-1} + b_{13} m_{t-2} + b_{21}^2 s_{t-3} + b_{21}^3 s_{t-4} + b_{31} \Delta q_t + b_{33} \Delta q_{t-1} + b_{33}^2 \Delta q_{t-2} + b_{41}^3 \Delta r_{t-1}](1 - F(\bullet)) + [a'_0 + a'_1 s_{t-1} + a'_2 s_{t-2} + a'_3 s_{t-3} + b'_1 z_{t-1} + b'_2 \Delta m_{t-1} + b'_3 \Delta y_{t-2} + b'_4 \Delta q_t]F(\bullet) + \nu_t \]

\[ F(\bullet) = 1 - \exp\{[\gamma(z_{t-2} - c)^2] \]

<table>
<thead>
<tr>
<th>( b_{11} )</th>
<th>( b_{12} )</th>
<th>( b_{13} )</th>
<th>( b_{21} )</th>
<th>( b_{31} )</th>
<th>( b_{32} )</th>
<th>( b_{33} )</th>
<th>( b_{41} )</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.261</td>
<td>0.429</td>
<td>-0.272</td>
<td>0.321</td>
<td>0.754</td>
<td>-0.351</td>
<td>-0.235</td>
<td>-0.015</td>
</tr>
<tr>
<td>(2.34)</td>
<td>(4.136)</td>
<td>(-2.74)</td>
<td>(4.73)</td>
<td>(7.23)</td>
<td>(-2.41)</td>
<td>(-1.66)</td>
<td>(-2.99)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>( b'_{11} )</th>
<th>( b'_{21} )</th>
<th>( b'_{31} )</th>
<th>( b'_1 )</th>
<th>( \gamma )</th>
<th>( c )</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.330</td>
<td>0.227</td>
<td>0.306</td>
<td>-0.06</td>
<td>-20.542</td>
<td>-0.035</td>
</tr>
<tr>
<td>(3.53)</td>
<td>(3.81)</td>
<td>(2.69)</td>
<td>(-2.55)</td>
<td>(-3.79)</td>
<td>(-6.38)</td>
</tr>
</tbody>
</table>

ARCH(1) = 1.08 [0.30]; ARCH(4) = 3.15 [0.53];
Q(8) = 6.05 [0.64]; Q(12) = 9.30 [0.68]; Q(16) = 12.62 [0.70];
Q^2 (8) = 3.95 [0.86]; Q^2 (12) = 9.80 [0.63]; Q^2 (16) = 13.78 [0.62];
JB = 2.94 [0.2300]; adj \( R^2 = 0.79 \); RESET(1) = 0.18 [0.67]; RESET(2) = 0.27 [0.87];

Note: Same as in Table 6.
Figure 1. Cointegrating residuals

Figure 2. The transition function against $\hat{z}_{1-2}$
Figure 3. Actual and fitted values of the change in real money demand
References


