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The Transactions Demand For Money
In The Presence Of Currency Substitution:
Evidence From Vietnam

[Running title: *Currency Substitution and the Demand for Money*]

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Abstract

Currency substitution – the use of foreign money to finance transactions between domestic residents – is widespread in low income and transition economies. Traditionally, however, empirical models of the demand for money tend to concentrate on the portfolio, motive for holding foreign currency, while maintaining the assumption that the income elasticity of demand for domestic money is invariant to the degree of currency substitution.. We offer a simple re-specification of the demand for money which more accurately reflects the process of currency substitution by allowing for a variable income elasticity of demand for domestic money. This specification is estimated for Vietnam in the 1990s. Using a standard cointegration framework we find evidence for currency substitution only in the long-run but well-defined wealth effects operating in the short-run.

I. INTRODUCTION

Dollarization is widespread in developing and transition economies. In many, limits on the use of foreign currencies for transactions purposes mean that foreign currency is held exclusively for portfolio purposes, allowing wealth to be hedged against domestic inflation and financial repression. If the capital account is open, the private sector's foreign currency balances are adjusted through capital account transactions, if it is closed, adjustment will be occur gradually through the current account. Local currency, however, remains the sole means of exchange in the domestic economy. The parallel circulation of foreign currency for transactions purposes between domestic residents – often taken as the precise meaning of currency substitution[1] – is less common. In the past, pure currency substitution has tended to occur only in conditions of conflict or state collapse; in recent years, however, systematic currency substitution has emerged as an important feature in a number of low income and transition economies.

Although there is an extensive empirical literature purportedly on currency substitution (see, for example, Adam, 1999, Agénor and Montiel, 1999, and Sriram, 1999), it tends, in fact, to focus exclusively on portfolio or wealth motives and not on the precise phenomenon of currency substitution, which is directly related to the transactions motive. In this paper we offer a simple modification of the standard empirical money demand function which allows for direct currency substitution effects as well as the standard portfolio effects of dollarization. This alternative characterization implies a variable income elasticity of demand for domestic money. Given the costs of switching currencies at the margin for transactions purposes, it is probable that currency-substitution is more likely to be observed as a long-run phenomenon, while portfolio effects dominate in the short-run, implying a constant short-run income elasticity of demand. We show how this hypothesis can be tested in the context of a non-linear (stationary) error correction framework.

We illustrate our approach by estimating the demand for money in post-liberalization Vietnam which, at the end of the 1980s in the wake of its transition towards a market economy, experienced rapid dollarization and has seen the emergence of significant currency substitution. The paper is organized as follows. Section II discusses the new specification of the demand for real money in

presence of currency substitution. Section III describes the stylized facts of the dollarization process in Vietnam, and Section IV describes the data and presents our results. Section V concludes.

II. THE DEMAND FOR MONEY IN THE PRESENCE OF CURRENCY SUBSTITUTION

The essential feature of currency substitution is that both domestic and foreign monies provide liquidity services in financing the exchange of goods and services, including non-tradables, between domestic residents. Clearly, for both currencies to be used in parallel there must be some costs to switching between currencies. Otherwise, for any given differential in the return, one currency would always dominate in financing transactions.[2] We can illustrate the potential implications for the demand for domestic money using the standard perfect-foresight, money-in-the-utility function structure outlined in Obstfeld and Rogoff (1999, pp551-553). We start with the representative agent's inter-temporal utility which is defined over consumption, denoted *C,* and liquidity services, *L*, where the latter can be provided by either the domestic currency, *M*, or the foreign currency, *F*, but where these are imperfect substitutes:

$$U = \sum_{s=t}^{\infty} b^{s-t} u \left\{ C_s, L(\frac{M_s}{P_s}, \frac{E_s F_s}{P_s}) \right\}.$$ (1)

Taking the path for prices and the exchange rate as given, the representative agent accumulates real bonds (*B*) and the two monies to maximize (1) subject to the budget constraint,

$$B_t + \frac{M_t}{P_t} + \frac{E_t F_t}{P_t} = (1+r)B_{t-1} + \frac{M_{t-1}}{P_t} + \frac{E_t F_{t-1}}{P_t} + Y_t - C_t - T_t$$ (2)

where *E* and *P* denote the nominal exchange rate and the domestic price level, *r* the real rate of return on the output-indexed bond, *Y* is income and *T* lump-sum taxes. The first-order conditions with respect to B_t, M_t and F_t are given by

$$u_C(C_t) = (1+r)b u_C(C_{t+1})$$ (3)

$$\frac{1}{p_t} u_C(C_t) = \frac{1}{p_{t+1}} b u_C(C_{t+1}) + \frac{1}{p_t} u_L L_M \qquad (4)$$

$$\frac{E_t}{p_t} u_C(C_t) = \frac{E_{t+1}}{p_{t+1}} b u_C(C_{t+1}) + \frac{E_t}{p_t} u_L L_F \qquad (5)$$

where u_C denotes the marginal utility of consumption, u_L the marginal utility of liquidity services, and L_M and L_F represent the marginal contribution of each money to the aggregate liquidity, L. Condition (3) is the standard consumption Euler equation, while (4) and (5), in conjunction with (3), define the portfolio balance between the three assets.

Noting that $\frac{P_{t+1}}{P_t} = (1 + p_{t+1})$ and $\frac{E_{t+1}}{E_t} = (1 + c_{t+1})$ where p and c denote the rate of inflation and exchange rate depreciation respectively, and imposing the Fisher condition that $(1 + i_{t+1}) = (1 + r)(1 + p_{t+1})$, where i denotes the nominal rate of interest, we can obtain the following implicit expression for the relative demand for domestic and foreign money

$$\frac{L_M}{L_F} = 1 + \frac{1}{L_F} \left(\frac{c_{t+1}}{1 + i_{t+1}} \right) \left(\frac{u_c(C_t)}{u_L} \right). \qquad (6)$$

The nested structure of (1) means that it is difficult to derive an explicit expression for the demand for domestic money. Nonetheless, equation (6) highlights the key argument in this paper. When $c_{t+1} = 0$ the relative demand for the two monies is constant and independent of the level of transactions (in other words both monies rise in proportion with the level of transactions so that the income elasticity is constant). By contrast, when $c_{t+1} \neq 0$ the relative demand for the two monies responds to a change in the level of consumption. To see this, consider the example suggested by Obstfeld and Rogoff (1999), in which $L(.)$ is linear in domestic money but is a negative quadratic function of foreign money, reflecting diminishing marginal opportunities for currency substitution

$$ L = \left[\frac{M_t}{P_t} + a_0 \left(\frac{E_t F_t}{P_t} \right) - \frac{a_1}{2} \left(\frac{E_t F_t}{P_t} \right)^2 \right]. \tag{7} $$

Assuming that $u(.)$ in equation (1) takes a standard CES-CRRA (*constant elasticity of substitution – constant relative risk aversion*) form, this specification implies a transactions elasticity of demand for domestic money of the form

$$ \boldsymbol{h}_C = \left(\frac{1-\boldsymbol{g}}{\boldsymbol{g}} \right) \left[\Omega \left(\frac{1+i_{t+1}}{\boldsymbol{c}_{t+1}} \right) \right]^{\boldsymbol{q}} \left(\frac{C}{m} \right)_t \tag{8} $$

where $\Omega = \left(\dfrac{1 - E_t (a_0 - a_1 f)}{P_t} \right) > 0$ for interior solutions, $\boldsymbol{q}$ is the elasticity of substitution between consumption and liquidity services, m and f are domestic and foreign currency in constant domestic prices. The key feature of this elasticity is that it is decreasing in the expected rate of depreciation, $\boldsymbol{c}_{t+1}$. The same result obtains, albeit in a more complex form, for more general representations of $L(.)$. The basic intuition, in any case, is relatively straightforward; first, increased consumption requires higher liquidity services in general; how much of these services are supplied by domestic and foreign currencies will depend on: (i) the elasticity of substitution between the two (itself a function of the structure of transactions costs), and (ii) the relative return to the two monies ($\boldsymbol{c}_{t+1}$).

Empirical Money Demand Functions

Empirical work on the demand for money and currency substitution typically starts with a specification of the form

$$ \frac{M}{P} = f(Y, x^e, \mathbf{Z}) \tag{9} $$

where M is the nominal domestic money aggregate, P the price level, Y a measure of the level of real economic activity, x^e the expected depreciation of the nominal exchange rate and Z is a vector of other

opportunity cost or shift factors (interest rates, inflation etc.)[3]. Estimation of (9) is then based on a semi-log representation with the general long-run form

$$(m - p)_t = b_0 + b_1 y_t + b_2 x_t^e + \text{?'}Z_t + e_t \tag{10}$$

where $m=log(M), p=log(P)$, and $y=log(Y)$. The fundamental argument in this paper is that the constant transactions elasticity of demand implied by (10), i.e. the parameter b_1, does not adequately capture the currency substitution phenomenon. Given this functional form, the relationship between the volume of transactions and the need for domestic money is constant, whatever the expected depreciation, so that the expected rate of exchange rate depreciation is treated only as an opportunity cost of *holding* domestic money, not a cost of *using* domestic money in transactions. As implied by (8), we argue that a given level of the economic activity will have a weaker (stronger) effect on the transactions demand for domestic money the higher (lower) the expected depreciation of the exchange rate, since at the margin individuals and enterprises will increase their holdings of foreign currency to finance transactions.

An obvious simple way of representing this competition between domestic and foreign currencies in financing transactions is to allow for the elasticity of the demand for domestic money to be a function of the expected rate of exchange rate depreciation. Hence

$$(m - p)_t = b_0 + b_1 x(x_t^e) y_t + b_2 x_t^e + \text{?'}Z_t + e_t \tag{11}$$

where now the transactions elasticity of demand, $b_1 x(x_t^e)$, is function of the expected rate of depreciation.

A natural functional form for this elasticity is a negative exponential $x(x_t^e) = e^{-d x_t^e}$ where d is expected to be positive (as in the traditional demand for money model). This functional form has the property that when agents expect no exchange rate depreciation, so that $x^e = 0$, then $e^{-d x_t^e} = 1$ and the transactions elasticity of demand is simply b_1. By contrast, when agents expect a depreciation (appreciation), so that $x^e > 0$ ($x^e < 0$), then $e^{-d x_t^e} < 1$ ($e^{-d x_t^e} > 1$) and the transactions elasticity

$b_1 x(x_t^e)$ is smaller (greater) than b_1. Moreover, if $d = 0$, which would correspond to the case where legal restrictions on currency substitution are binding, the specification again collapses to the constant elasticity case.

Unfortunately, as Park and Phillips (2001) show, if y is non-stationary, as is typically the case, the limiting distributions for the non-linear regression under this class of exponential transformation is not well-defined, since the transformation is unbounded (as $x^e \to -\infty$). By contrast, the logistic function $(1/(1 + e^{dx^e}))$, which has similar local characteristics to e^{-dx^e} but which is bounded as $x^e \to -\infty$, is a member of the class of asymptotically homogenous functions for which limit distributions are well-defined. We therefore let $x(x^e) y_t = \left[\dfrac{y}{1 + e^{dx^e}} \right]_t$, which implies that the transactions elasticity of demand for domestic money, denoted h_y, is now $h_y = b_1 \left(\dfrac{1}{1 + e^{d x^e}} \right)$ which has the following properties. First, for $x=0$, $h_y = b_1 / 2$, for $x>0$ (the case of an expected depreciation) $h_y < b_1 / 2$ and *vice versa* of the case where $x<0$. However the income elasticity is now bounded below (as $x \to \infty$ $h_y \to 0$) and above (as $x \to -\infty$ $h_y \to b_1$). As d increases, the function tends to its limit more rapidly.

The exchange rate elasticity of money demand is given by

$$h_x = \left[-b_1 dy \left(\frac{e^{dx^e}}{\left(1 + e^{dx^e}\right)^2} \right) + b_2 \right] x^e .$$ This elasticity is increasing in x, and, assuming that $b_2 < 0$, is strictly negative.

Our preferred empirical specification therefore takes the general form

$$(m - p)_t = b_0 + b_1 \left[\frac{y}{1 + e^{dx^e}} \right]_t + b_2 x_t^e + ?' Z_t + e_t . \tag{12}$$

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Equation (12) provides a basis for a direct test of the currency substitution hypothesis (conditional on the presence of portfolio considerations). The restriction $d = 0$ implies a constant income elasticity model, while rejecting the restriction in favour of $d > 0$ indicates the presence of currency substitution for transaction motives. If b_2 is different from zero (and negative), then the portfolio dimension of dollarization is also present (i.e. exchange rate depreciation is an opportunity cost of holding domestic money as a store of value). In the remainder of this paper we test the implications of this argument.

III. DOLLARIZATION IN VIETNAM

Background

In Vietnam, the widespread circulation of US dollars first appeared during the war against the United States when the American armed forces occupied the south of the country. Following the reunification of Vietnam in 1975, the holding of foreign currency by residents was strictly forbidden in order to reinforce the national currency unit, the dong. Dollarization reappeared in the 1980s as a result of unsustainable inflationary pressures during the final years of the planned economy and the transition towards a market economy. The relaxation of price controls and the loose monetary stance in the face of weak domestic supply fuelled a period of high inflation episodes culminating in the hyper-inflation in 1986-88 (see Figure 1). Following the sharp depreciation of the parallel exchange rate, the official exchange rate of the dong was dramatically devalued (from 15 dong per dollar at the end-1985 to VND/USD 3,000 at the end-1988).

<Insert figure 1 here>

Further liberalization followed in 1988-89. Controls on external trade were relaxed, virtually all domestic price controls were eliminated, and the exchange rate regime unified. This was supported by the introduction of foreign currency deposit accounts for individuals and enterprises, although the dong remained the only legal tender for domestic transactions. Institutional reforms were

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accompanied by changes in the monetary policy stance. Interest rates on dong-denominated saving deposits were raised to above the inflation rate and kept positive in real terms throughout 1989, stimulating a significant rise in the demand for dong liquidity and a spectacular decline in inflation which fell from 350% in 1988 to 35% in 1989. This gain was short-lived; weak domestic credit control saw the money supply increase again and inflation increased to 67% in 1990 and 72% in 1991, real interest rates once again turned negative, and the dong depreciated to VND/USD 14000 by the end of 1991.

At the end of 1991 the Vietnamese authorities decided to implement a shock-therapy approach in order to break the inflation-depreciation spiral. After opening two foreign exchange markets, one in Hanoi and one in Ho Chi Minh City, the monetary authorities sold huge amounts of dollars on these markets, causing an appreciation of the exchange rate from VND/USD 14,000 at the end-1991 to VND/USD 10,500 in January 1993. Moreover, the central bank announced that it stood ready to satisfy any demand of gold purchase made by individuals and enterprises (Guillaumont Jeanneney, 1994a and 1994b). These measures seem to have played an important role in establishing the credibility of the authorities' stabilization policy and for the next five years they successfully pursued a policy of shadowing the US dollar (at a rate of around VND/USD 11,000). From 1992 to the end of the decade, inflation averaged less than 10 percent per annum.

A weakening trade balance in 1995-96 prompted concerns about the overvaluation of the dong and induced to a speculative demand for dollars and eventual depreciation of the dong. This process was reinforced by the Asian crisis in 1997-98 that curbed the dollar inflow into Vietnam. Despite new administrative measures, such as foreign exchange surrender requirements and import restrictions, the exchange rate came under pressure and depreciated from VND/USD 11,000 at the end of 1996 to VND/USD 13,900 at the end of 1998. From the beginning of 1999 a crawling depreciation was applied until a rate of VND/USD 15,000 at the end of 2001.

The Scale of Dollarization

Foreign currency deposits and US dollar banknotes represent a substantial proportion of the total money supply in Vietnam. The exact volume of US banknotes in circulation is hard to determine

precisely, but one source estimated it to be around VND 42 trillion (USD3billion at a rate of VND/USD 14,000) in 2000, approximately 10 percent of GDP (Unteroberdoerster, 2002)[4] . Foreign currency deposits in the banking system are more easily tracked. As Figure 2 indicates, these rose very rapidly during the early 1990s in line with inflation and the sharp depreciation of the exchange rate. Since the mid-1990s, however, foreign currency deposits have grown steadily, from just under 20 percent of broad money in 1994 to just under 35 percent in 2001, although these accounts cannot be used directly for domestic transactions settlements. By the end of 2000, total foreign currency accounted for 42 percent of the total money in Vietnam (Table 1).

<Insert table 1 here>

<Insert figure 2 here>

IV. DATA AND ESTIMATION

We focus on the demand for the narrow money aggregate, M1, which consists of dong in circulation outside banks and dong-denominated demand deposits in the banking system. The price index is the consumer price index, the only consistently reported price index in Vietnam. Data on these variables are available on a monthly frequency. We face a greater problem in choosing a measure for the level of real economic activity. Ideally we would use a measure of gross domestic product or gross domestic expenditure but the series on these data are incomplete and only available at annual or quarterly frequency. Instead we use the index of monthly industrial output. Although this measures less than one third of constant-price GDP in Vietnam over the 1990s,[5] industrial activity is highly correlated with total GDP and has the advantage that it reflects the sector of the economy in which currency substitution is, arguably, most likely to be observed.

The lack of direct measures of expectations means that we must proxy the expected rate of exchange rate depreciation. Ideally we would use a measure based on the forward exchange rate, but in the absence of such markets empirical work on developing countries tends to use a variety of

proxies, either a rational expectations structure in which the actual current or future rate of depreciation is used to proxy the expected rate in which lagged values of the exchange rate depreciation and other regressors are used to instrument the proxy (e.g. Adam, 1999; Bush, 2001; Perera, 1993; Chowdhury, 1997; Weliwita, 1998; Arize, 1994; Choudhry, 1998; Tan, 1997), or an adaptive structure based directly on lags of the rate of depreciation (Bahmani-Oskooee, 1991; Arize, 1992).

We have chosen to compute the expected rate of depreciation using a moving average of actual and lagged values of exchange rate depreciation,[6] defined as $x_t^e = \frac{1}{4}\sum_{s=0}^{3} x_{t-s}$ where $x_t = (E_t - E_{t-1})/E_{t-1}$ and E is the parallel exchange rate (i.e. the Hanoi black market rate). Given the degree of dollarization in the Vietnamese economy, the parallel exchange rate provides a reliable indicator of the marginal opportunity cost of holding domestic as opposed to foreign currency.

All the data have been obtained from the State Bank of Vietnam and cover the period from January 1991 to June 1999, giving 102 data points for estimation. As indicated in Appendix Table 1 there is strong evidence that real money balance and real economic activity contain a unit root, and slightly weaker evidence that the expected depreciation does. We therefore employ a cointegration framework for analysis.

Since equation (12) is non-linear in parameters, these cannot be estimated using conventional (linear) cointegration methods.[7] We therefore structure our estimation procedure as follows. First, to examine the possible long-run equilibrium structure of the model we impose the restriction that $d = 2$ so that (12) is rendered linear in (free) long-run parameters thus [8, 9]

$$(m-p)_t = b_0 + b_1 \left[\frac{2y}{1+e^{2x^e}} \right]_t + b_2 x_t^e + e_t . \qquad (12')$$

Our analysis is therefore based on a version of (12'), which allows us to test the number of cointegrating vectors and the weak exogeneity status of the regressors (which is required to define a

single equation dynamic representation of (12')). Then, in a second step we explore the (linear) dynamic representation of 12', always with $\delta=2$. Finally, we re-estimate the (stationary) dynamic error correction representation of (12) using non-linear methods to directly estimate the coefficient δ. We will show that this last estimation accepts the restriction that $\delta=2$. Moreover, the similarities between the linear and the non linear estimates allows us to consider the results of the (linear) cointegration analysis, i.e. the number of cointegrating vectors and the weak exogeneity, as a valid basis for the non-linear dynamic error correction representation.

Cointegration analysis

We estimate a vector error-correction model (VECM) of the form

$$\Delta Y_t = \boldsymbol{a}\boldsymbol{b}\,'Y_{t-k} + \sum_{i=1}^{k-1}\Gamma_i\Delta Y_{t-i} + D_t + \boldsymbol{e}_t \tag{13}$$

where $Y_t = \{(m\text{-}p)_t\,,\left(\dfrac{2y}{1+e^{2x^e}}\right)_t,\ x_t^e\,\}$, $\boldsymbol{b}\,'$ denotes the matrix of parameters of the cointegrating vectors

($\boldsymbol{b}'Y_{t-k}$ are long-run relationships), and $\boldsymbol{a}$ the matrix of equilibrium-correction or feedback effects, Γ_i is the matrix of short-run parameters, D_t denotes deterministic components (the constant and monthly seasonal dummy variables)[10] and $\boldsymbol{e}_t$ is the vector of error-term.

A lag length *k=6* fully captures the dynamics between the variables of the vector *Y* and renders $\boldsymbol{e}_t$ approximately Gaussian. Table 2 reports the VECM residual diagnostic statistics and Table 3 summarizes the principal features of the cointegration analysis.

<insert table 2 here>

<insert table 3 here>

The residual-based tests reported in Table 2 indicate the absence of serious statistical misspecification of the underlying vector autoregression while the first block of Table 3 suggests the presence of two

cointegrating vectors. Neither is immediately interpretable, although the first vector (the column denoted β1) would appear to be broadly consistent with the long-run demand for money function of the form:

$$(m-p)_t = 3.74 + 1.05\left[\frac{2y}{1+e^{2x^e}}\right]_t + 1.94\,x_t^e \qquad (14)$$

in which the residuals enters with a feedback coefficient of $\alpha 1 = -0.07$ in the short-run demand for money equation $\boldsymbol{D}(m\text{-}p)_t$. The problem with this interpretation is that the coefficient on the expected rate of depreciation is positive when theory would suggest it be negative. However, as the third block of Table 3 suggests it is possible to reject this variable but not the other two from the (first) cointegrating vector (the Likelihood ratio test has a value of $\chi^2(1) = 0.37$ [0.54]), suggesting that in the long run at least, and conditional on the presence of currency substitution, portfolio-based dollarization effects do not appear to operate on the demand for M1. This result is consistent with the evidence from the second vector, β2, which indicates that x^e appears to be stationary (as suggested by the univariate tests reported in Appendix Table 1).

These two features point to the following restrictions on the vector $\Pi = \boldsymbol{ab}'$ conditional on the cointegrating rank being $r=2$.

$$\Pi_r = \boldsymbol{a}_r \boldsymbol{b}_r' = \begin{pmatrix} \boldsymbol{a}_{11} & 0 \\ 0 & \boldsymbol{a}_{22} \\ 0 & \boldsymbol{a}_{23} \end{pmatrix} \begin{pmatrix} 1 & \boldsymbol{b}_{12} & 0 & c_1 \\ 0 & 0 & 1 & c_2 \end{pmatrix}. \qquad (15)$$

The restrictions on the $\boldsymbol{b}$ component of the Π matrix imply that no long-run portfolio effect enters the first cointegrating vector representing the long-run demand for money, while those on the $\boldsymbol{a}$ component imply that the (short-run) demand for money corrects to deviations from the cointegrating vector only, and that the first cointegrating vectors only feeds back onto the short-run demand for money. If these restrictions are accepted then it is possible to move from the VECM to a single-equation equilibrium-correction representation of the short-run demand for money where the variables

$\left[\dfrac{2y}{1+e^{2x^e}}\right]$ and x^e are weakly exogenous (Johansen, 1992). As we discuss below, this provides a basis

for investigating the non-linearity of our preferred specification more thoroughly.

<Insert table 4 here>

Table 4 reports the results from estimating the model under the restrictions described in (15). The likelihood ratio test indicates that the restrictions are accepted by the data leading to the restricted long-run demand function taking the form

$$(m-p)_t = 4.41 + 0.95\left[\frac{2y}{1+e^{2x^e}}\right]_t \tag{16}$$

with a feedback coefficient on Δ(m-p) of –0.10 (i.e. 10 percent per month).

So far we have only rejected a long-run *portfolio dimension* to dollarization in Vietnam. Since we have imposed the restriction $d=2$, our characterization of the currency substitution hypothesis remains incomplete, although not inconsistent with the data under this restriction. We therefore turn to alternative single equation representations of the data to attempt to cast more light on the relationship.

Single equation error-correction model.

The general specification of the dynamic model for $\mathbf{D}$*(m-p)*$_t$ can be written as:

$$?(m\text{-}p)_t = a\,[\,(m\text{-}p) - \mathbf{b}_l\left[\frac{2y}{1+e^{dx^e}}\right] - c\,]_{t-k} + \sum_{i=0}^{k-1} ?_i\,?\,Y_{t-i} + \mathbf{n}_t \tag{17}$$

where Δ is the monthly difference operator, the term in square brackets the deviation of (m-p) from its long-run equilibrium level), a is the speed-of-adjustment coefficient, and $\mathbf{g}$ the vector of parameters describing the short-run dynamic behavior of the demand for money. Our results are reported in summary form in Table 5[11]. We estimate three alternative representations of (17). First we estimate the model using a two-step method which embeds the cointegrating vector (16) as the representation

of the long-run demand for money, and second we re-estimate (17) as single-equation equilibrium correction model. In both cases we maintain the restriction that $d=2$. Finally, we estimate a non-linear equilibrium-correction model in which we directly estimate d, allowing us to directly test the restriction we exploited to implement the (linear) cointegration analysis[12]. The dynamics of (17) are data-determined; we start by allowing for both currency substitution and portfolio effects of dollarization to shape the demand for money in the short-run, and follow a standard reduction strategy to eliminate insignificant regressors. We also include a full set of seasonal dummy variables which, in all cases, are jointly significant.

<Insert table 5 here>

All three representations appear to be broadly coherent with the data and consistent with theory. There is no evidence of serious dynamic misspecification nor of significant parameter instability[13]. The equilibrium-correction structure is validated and the feedback of a plausible magnitude of around 11 percent per month across all three models.

In terms of the principal argument of the paper, Table 5 highlights two key results. The first concerns the nature of the long run demand for money. Columns [1] and [2] suggest that the two-step and one-step equilibrium correction models in which we impose the restriction $d=2$ generate virtually identical results, both in terms of their statistical properties and the point estimates they generate. Column [3], which reports the results of using a non-linear ECM estimator, generates an estimate of $d=1.832$. This is strongly statistically different from zero implying that we can reject a constant long-run income elasticity of the demand for money in favor of the currency-substitution hypothesis.

We cannot, however, reject the restriction $d=2$; the LR test of the restriction has a value of $?^2(1)=0.055[0.814]$. The implication, as implied by the comparison across the columns of Table 5, is that the simple specification for the income variable, $\dfrac{2y}{1+e^{2x^e}}$ offers a good approximation of the data.

The second result is that the model does not admit a role for currency substitution in the short-run, even though it is present in the long-run. By contrast there is strong evidence of a more conventional portfolio effect at work. An increase in the expected rate of exchange rate depreciation

induces a shift out of domestic money, with this portfolio effect being felt with a three to four month lag. This is consistent with the idea that because of the costs of changing the transactions technology agents need time to adjust their behavior in transactions: sellers and buyers have to learn how to use the new currency and to approve the adoption of the new currency as the means of payment, short-run dollarization is likely to dominated by portfolio rather than currency substitution considerations.

V. CONCLUSION

The evidence presented in this paper suggests that traditional linear specifications of the demand for money may be mis-specified for economies in which currency substitution is an important phenomenon. We have shown that for one such country, Vietnam, the data for the 1990s suggest a characterization of the dollarization process in which currency-substitution effects alter the economy's transactions technology and hence the (traditionally specified) long-run income elasticity of the demand for money whereas more traditional portfolio or hedging considerations are relevant only in the short-run. Our specification implies a variable long-run income elasticity of demand. In industrialized countries monetary targeting has tended to be abandoned for interest rate policies, but in developing countries the relative thinness of financial markets does not allow monetary authorities to rely only on interest rate and a monetary aggregate is often required as an intermediate target of monetary policy. In such circumstances, the failure to estimate correctly the currency substitution effect will lead to systematic mis-prediction of the demand for money in circumstances where there is a tendency for the nominal exchange rate to move over time (e.g. in high inflation contexts).

An important feature of the results for Vietnam is that a simple representation for the income elasticity (i.e. where $d=2$) could not be rejected against a more general specification. If this result were true more generally it would imply a very simple respecification of the demand for money. A natural next step in the analysis is therefore to examine the properties of the demand for money across a wider range of low income and transition economies.

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APPENDIX

<center><Insert table A.1 here></center>

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[1] Calvo and Végh (1992) use the term *currency substitution* to describe the use of a foreign currency as a means of exchange and the term *dollarization* denotes the use of a foreign currency in any of its three functions: unit of account; means of exchange; or store of value. We maintain this distinction here.

[2] In Matsuyama *et al* (1993), for example, these costs arise as a result of the random matching of agents. Alternative mechanisms might include menu costs or other costs associated with maintaining parallel payment technologies.

[3] For example Adam (1992, 1999), Arize (1992, 1994), Bahmani-Oskooee and Malixi (1991), Buch (2001), Choudhry (1998), Chowdhury (1997), Dekle and Pradhan (1997), Perera (1993), Tan (1997), Weliwita (1998) and the survey paper by Sriram (2001).

[4] This is consistent with the anecdotal evidence suggesting that private sellers of goods or services rarely refuse to conduct trade in US dollar instead of Vietnamese dong, and that curb markets in foreign currencies are widespread in all the cities. In both cases, these activities are officially forbidden but very widely tolerated.

[5] In 1990, the structure of GDP was agriculture 38.7%, industry 22.7% and services 38.6% and in 1999 agriculture 25.4%, industry 34.5% and services 40.1%. Source GSO (2000), table 17 p.28. The sample correlation between industrial output and total constant price GDP, based on quarterly data from CIEM (2000), is 0.983.

[6] Other specifications of the expected depreciation do not reject our model, but preliminary results have indicated that this type of specification leads to the best goodness-of-fit of the model.

[7] The literature on 'non-linear cointegration' is relatively new and incomplete, see Park and Phillips (2001). Empirical work to date tends to focus on non-linearities in the adjustment process. See for example Enders and Granger (1998).

[8] The initial parameter value $d = 2$ was chosen using a simple grid-search method using the maximum of the first eigenvalue within the Johansen cointegration analysis as the criterion. Having chosen this initial value we then scaled the logistic function by a factor of 2. This does not alter the model but allows a direct interpretation of the coefficients of the model.

[9] Given the relatively under-developed financial system in Vietnam, we find that the demand for money can be fully described in terms of the income and the rate of exchange rate depreciation; we have therefore dropped the vector **Z** from our empirical specification.

[10] Preliminary results suggest that the deterministic components included in the dynamics should be a constant restricted to the long-run relationships, unrestricted (centred) seasonal dummies but no deterministic trend.

[11] Full details of all the results are available from the corresponding author.

[12] The non linear least squares estimation is performed using PcGive 10.0 ; see Hendry and Doornik (2001) vol. 1, chapter 17, pp. 247-248.

[13] The evidence from Hansen's instability tests is supported by recursive estimation results, which are available on request from the corresponding author.
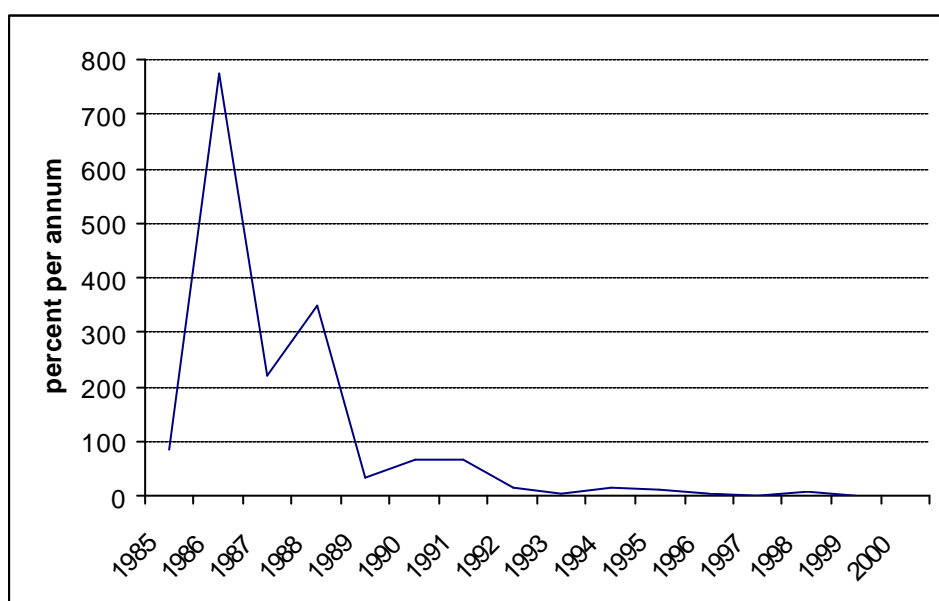
Notes: Rate of change of retail price index of goods and consumer services, in %. Data before 1991 measure the retail price index of goods, excluding consumer services. *Data source*: GSO (2000).

Fig. 1. *Inflation 1985-2000 (December to December)*
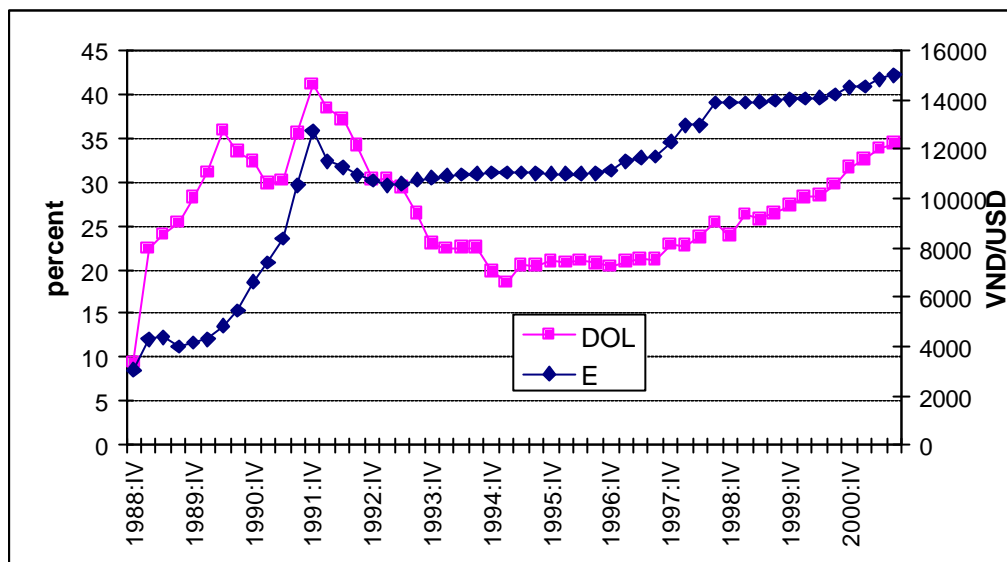
Notes: DOL, left-scaled, ratio of foreign currency deposits to broad money (including foreign currency deposits). E, right-scaled, is the banking exchange rate of the Vietnam dong to the US dollar. *Data source:* State Bank of Vietnam.

Fig. 2. *Dollarization (DOL) and the VND/USD exchange rate (E). 1988-2001.*

Table 1.

The composition of broad money including foreign currency in circulation (Dec. 2000)

	trillions of dong	in %
Domestic currency in circulation outside banks	52.2	20
Foreign currency in circulation outside banks (estimates)*	*42.0*	*16*
Domestic currency demand deposits	58.4	22
Foreign currency demand deposits	*16.8*	*6*
Other domestic currency deposits	41.9	16
Other foreign currency deposits	*53.6*	*20*
Total	264.9	100
of which foreign currency	*112.4*	*42*

Notes: Data Sources: IMF (2002) and * from Unteroberdoerster (2002). Calculations from the authors.

Table 2. *VECM residuals diagnostic statistics for Y_t.*
Sample 1991(8) – 1999(6). T=95.

	AR(1)	JB	ARCH(1)	H
Y	0.81 [0.61]	15.5 [0.02]		0.83 [0.89]
$m\text{-}p$	0.01 [0.93]	7.67 [0.02]	1.71 [0.19]	0.32 [0.99]
$\left(\dfrac{2y}{1+e^{2x^e}}\right)$	3.33 [0.07]	1.44 [0.48]	0.72 [0.40]	1.40 [0.17]
x^e	1.23 [0.27]	3.07 [0.21]	11.4 [0.00]	3.81 [0.00]

Notes: AR(1) is an LM test for first-order autocorrelation. JB is the Jarque-Bera test for normality. ARCH is an LM test for conditional heteroscedasticity. H is White's (1980) test for heteroscedasticity. Marginal significance levels are in parentheses.

Table 3. *I(1) cointegration analysis*

Reduced-Rank Statistics

Eigenvalues	H0: rank =	Trace test
0.252	0	55.811 [0.000]
0.199	1	27.966 [0.003]
0.067	2	6.6902 [0.148]

Standardised eigenvectors (scaled on diagonal) and adjustment coefficients.

b	**b1**	**b2**	**V3**	**a**	**a1**	**a2**	**V3**
$m\text{-}p$	1.00	-1.31	-0.12	$\mathbf{D}(m\text{-}p)$	-0.07	0.02	0.16
$\left[\dfrac{2y}{1+e^{2x^e}}\right]$	-1.05	1.00	0.09	$\mathbf{D}\left[\dfrac{2y}{1+e^{2x^e}}\right]$	-0.01	-0.07	0.55
x^e	-1.94	-7.33	1.00	$\mathbf{D}\,x^e$	-0.00	0.00	-0.08
C	-3.74	7.26	0.71				

Individual significance tests

	$m\text{-}p$	$\left[\dfrac{2y}{1+e^{2x^e}}\right]$	x^e	Constant
$\chi^2(1)$	4.08 [0.04]	4.81 [0.03]	0.37 [0.54]	2.77 [0.09]

Notes: Marginal significance levels are in square parentheses. Individual significance tests conducted under the assumption of one cointegrating vector.

Table 4. *Restricted eigenvectors and adjustment coefficients*.

b	$b1$	$b2$	a	$a1$	$a2$
$m\text{-}p$	1	0	$\boldsymbol{D}(m\text{-}p)$	-0.10 [0.02]	0
$\left[\dfrac{2y}{1+e^{2x^e}}\right]$	-0.95 [0.08]	0	$\boldsymbol{D}\left[\dfrac{2y}{1+e^{2x^e}}\right]$	0	-0.59 [0.19]
x^e	0	1	$\boldsymbol{D}\,x^e$	0	-0.03 [0.02]
c	-4.41 [0.61]	0.05 [0.02]			

Notes: Standard errors of unrestricted coefficients estimates reported in [..]. LR test of restrictions: Chi^2(4)= 4.64 [0.326]

Table 5. Estimations of alternative ECM representations of (17).

Dependent variable is $\Delta(m\text{-}p)$. 1991(1)-1999(6). T=96.

	[1] 2-steps OLS	[2] 1-step OLS	[3] NLS
Long-run coef			
d	2	2	**1.832**
			(2.57)
β_1	0.953	**0.909**	**0.924**
		(4.95)	(3.13)
constant	4.415	**4.745**	**4.929**
		(4.32)	(4.41)
Dynamic coef			
a	**-0.102**	**-0.105**	**-0.109**
	(-5.19)	(-5,17)	(-3.77)
? for :			
$?(m-p)_{t-1}$	**-0.303**	**-0.306**	**-0.307**
	(-2.86)	(-2.88)	(-2.87)
$? x^e_{t-3}$	**-0.620**	**-0.592**	**-0.578**
	(-1.82)	(-1.72)	(-1.63)
$? x^e_{t-4}$	**-0.847**	**-0.832**	**-0.819**
	(-2.51)	(-2.45)	(-2.35)
R^2	0.546	0.548	
LL	225.2	225.4	225.4
Eqn s.e.	0.025	0.025	0.026
AR(6)	[0.857]	[0.864]	[0.856]
ARCH(6)	[0.339]	[0.325]	[0.337]
H	[0.186]	[0.269]	[0.419]
J-B	[0.008]	[0.010]	[0.009]
Inst-var	0.155	0.169	
Inst-joint	3.793	3.934	

Notes: Coefficient of seasonal dummies variables not reported. t-value in brackets. LL is the log-likelihood. s.e. is the standard error. AR(6) and ARCH(6) are tests against the null of autocorrelation and autoregressive conditional heteroscedasticity of order 6. H and J-B are tests against the null of homoscedastic and normally-distributed errors. Inst-var and Inst-joint are Hansen's tests for variance and joint parameter stability. Figures in square brackets [..] are tests statistics marginal significance levels. See for details the Pc Give 10.0 Manual by Hendry and Doornick (2001).

Table A.1. *ADF test for monthly series with 6 initial lags. 1991 (8) to 1999 (6). T=95.*

	m-p		y		x^e		$\frac{2y}{1+e^{2x^e}}$	
	Test	ADF-t	Test	ADF-t	Test	ADF-t	Test	ADF-t
level	ADF(2)	-1.18	ADF(5)	0.33	ADF(5)	-2.46	ADF(4)	-1.685
1st diff.	ADF(1)	-5.12**	ADF(4)	-5.96**	ADF(4)	-5.05**	ADF(3)	-6.91**

Note: Critical values used in ADF test: 5%=-2.89, 1%=-3.50

ADF (lags) with constant, **deterministic trend**, and centred seasonal dummies.

	m-p		y		x^e		$\frac{2y}{1+e^{2x^e}}$	
	Test	ADF-t	Test	ADF-t	Test	ADF-t	Test	ADF-t
level	ADF(2)	-1.66	ADF(6)	-3.43	ADF(5)	-2.51	ADF(4)	-2.77
1st diff.	ADF(1)	-5.17**	ADF(4)	-5.95**	ADF(4)	-5.11**	ADF(3)	-6.96**

Note: Critical values used in ADF test: 5%=-3.45, 1%=-4.05