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# Currency substitution and monetary autonomy: the foreign demand for US demand deposits

JEFFREY H. BERGSTRAND AND THOMAS P. BUNDT*

*University of Notre Dame, Notre Dame, IN 46556, USA*

This paper presents new evidence of currency substitution using the two-step estimator of cointegrated systems developed by Engle and Granger (1987). We estimate and find cointegration among variables suggested by the money-services model of currency substitution. Following Engle *et al.* (1989), we estimate an error correction model for each of five countries, allowing isolation of long-term from short-term influences on the foreign demand for US demand deposits. Our results suggest that currency substitution is a potentially significant phenomenon influencing long-term monetary policy independence.

World economic performance under the current international monetary system of flexible exchange rates has renewed interest in the autonomy of domestic monetary policies. One factor that potentially undermines independent monetary control is currency substitution. In the presence of currency substitution, a flexible exchange rate between two countries will not ensure domestic monetary autonomy, as had once been thought. For instance, if private agents view two currencies as substitutable inputs to the production of money services, then a rise in the opportunity cost of holding one currency can lead to a rise in both countries' demand for the other currency, a flow of that currency across borders, and an ensuing loss of monetary control; see Joines (1985) and Thomas (1985). Currency substitution has been advanced also as an explanation of exchange rate volatility; see Girton and Roper (1981) and Boyer and Kingston (1987).

This study provides empirical evidence in support of the presence of currency substitution—in particular, dollar substitution—within several major industrial countries. The paper extends previous studies in three ways. First, we utilize data on private nonbank foreign holdings of US demand deposits. Previous studies have typically used national monetary aggregates, which necessarily combine the effects of domestic and foreign demand for money and may have resulted in those studies finding little evidence in support of currency substitution. Second, we estimate an error correction model of the foreign demand for US money by

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application of the two-step estimator of cointegrated systems as developed in Engle and Granger (1987). Third, by considering both restricted and unrestricted error correction models, we distinguish between short- and long-run currency-substitution behavior. We find evidence in support of currency substitution as a long-run obstacle to independent US monetary control.

## I. Theoretical issues in the money–services approach

The theoretical framework for our analysis follows Joines (1985) and Thomas (1985). Currencies are held as inputs into the production of transactions services. Accordingly, the conditions describing optimal money demand are from production theory, not financial theory (on the latter, see Cuddington, 1983). Thomas (1985) demonstrated, in an expected utility maximizing model, that the usual opportunity cost variables influencing portfolio asset demands in general (including expected exchange rate changes) fail to influence money demands in particular. Here, the only opportunity cost variables influencing money demands are the home and foreign nominal interest rates, representing the marginal productivities of the home and foreign monetary inputs in the production of money services.

Consider two countries, home ($H$) and foreign ($F$). As in Joines (1985), the demand for real balances of currency $i$ in country $j$ is assumed to be given by:

$$\langle 1 \rangle \qquad M_{ij}^{D}/P_i = [k_{ij}(r_i, r_j)]Y_j, \qquad i,j = H, F,$$

where $M_{ij}^{D}$ is the nominal stock of currency $i$ demanded by residents of country $j$, $P_i$ is the price level in country $i$, $r_i$ ($r_j$) is the nominal interest rate in country $i$ ($j$), and $Y_j$ is real income in country $j$. If individuals minimize the cost of producing a given amount of money services, function $k_{ij}$ will be negatively related to $r_i$, the opportunity cost of holding currency $i$. Moreover, if the two currencies $H$ and $F$ are imperfect substitutes in the production of money services, $k_{ij}$ may be positively related to $r_j$, as will be shown.

Like Joines (1985), we consider the (simpler) asymmetric case where—in a two-country world—home country residents do not hold foreign money but foreign residents hold home money (*i.e.*, $k_{FH} = 0$). This asymmetry accords with our empirical investigation where the home country is the United States and the foreign country is one of potentially six other industrial countries. The historical prominence of the US dollar in international transactions is consistent with this assumption.

In this asymmetric case, equilibrium in the home country money market requires:

$$\langle 2 \rangle \qquad M_H/P_H = [k_{HH}(r_H)]Y_H + [k_{HF}(r_H, r_F)]Y_F,$$

where $M_H$ is the nominal stock of home money. Consider now an exogenous increase in the foreign interest rate ($r_F$), the opportunity cost of holding foreign money. A rise in $r_F$ lowers foreign demand for real foreign money ($k_{FF}$) and raises foreign demand for real home money ($k_{HF}$), since the relative opportunity cost of using home money as a monetary input is now lower. Because foreign money is not a substitutable input into the home money–services production function, the rise in $r_F$ has no impact on home money demands (*i.e.*, $k_{FH} = 0$ and $k_{HH}$ is unchanged). The consequent excess worldwide demand for real home money causes the home currency to appreciate, the home price level ($P_H$) to fall, and the real stock of home money to increase. Since the home residents' demand for real home money is unchanged, home currency must flow to foreign residents. This flow, initiated by changes in the foreign interest rate, is termed 'currency substitution'.

This positive relationship between $k_{HF}$ and $r_F$ can potentially erode the home country's monetary autonomy. For instance, monetary policy's effectiveness is enhanced the greater is the stability and predictability of the velocity of money. A typical definition of the home country's velocity ($V_H$), in the context of this model, is $(P_H Y_H)/M_H$. If $r_F$ significantly alters $k_{HF}$, and consequently $P_H$, home country velocity will be destabilized from abroad, eroding home monetary autonomy.

Note, if the two currencies are not substitutes in production, neither $k_{HH}$ nor $k_{HF}$ will be related to $r_F$. However, the demand for real home currency is still related to real foreign, as well as domestic, income; real home money demand is only independent of the foreign interest rate.

## II. Data limitations

Our focus is to estimate foreign demand for real home money functions, where the United States is taken as the home country and potentially six major industrial countries—Canada, Italy, Japan, the United Kingdom, Switzerland, and (West) Germany—are the foreign countries. Foreign country selection was based upon data availability and the absence of exchange controls and capital-outflow restrictions.

Time-series data on foreign holdings of US 'currency' are unavailable, of course. However, the US *Treasury Bulletin* (Table CM-I-4) provides quarterly data reported by US banks on demand deposit liabilities (in US dollars) to private nonbank foreign residents starting in 1978:Q2. To our knowledge, such detailed data on foreign holdings of US money have never been used to study currency substitution.[1]

Before estimation, each country was examined for the presence of exchange restrictions and capital controls, which might have hindered optimal money-demand behavior. A detailed record of the system of controls can be found in the International Monetary Fund's *Annual Report on Exchange Arrangements and Exchange Restrictions*. Annual reports from 1978 through 1988 were examined to determine which countries had controls in place that might have effectively constrained currency-substitution behavior.

The reports revealed that, for our purposes, no effective controls over exchange transactions and outward capital mobility prevailed in Canada, Germany, and Switzerland over the sample period. Exchange controls were abolished in the United Kingdom in October of 1979, with the exception of transactions with Zimbabwe. Several measures were announced in Japan in the spring of 1978 liberalizing regulations concerning the acquisition of foreign currency deposits by Japanese residents. In Italy, capital outflows by residents had to be accompanied by lire deposits in non-interest-bearing Italian accounts; while reducing the effective return on foreign investment, the restrictions did not prohibit such investment.

## III. Econometric issues

The simple static econometric analogue to the foreign component of equation ⟨2⟩

is:

$$\langle 3 \rangle \qquad \ln(M_{HF}/P_H) = \beta_0 - \beta_1 r_H + \beta_2 r_F + \beta_3 \ln Y_F + u,$$

where $u$ is an error term and the $\beta$s are defined to be positive. However, estimation of equation ⟨3⟩ as is raises several econometric issues. First, the time-series properties of the data should be investigated. For instance, are the variables stationary in levels or first-differences? Estimation of equation ⟨3⟩ using nonstationary data may lead to unreliable $t$-statistics as the underlying time series would have theoretically infinite variances. Even if each individual time series is level nonstationary, is the linear combination of these series as suggested by equation ⟨3⟩ level stationary? Second, the dynamic adjustment of the demand for money is neglected. Notably, is the adjustment of money demand to changes in opportunity cost and income variables instantaneous? If not, how should the dynamic adjustment be modeled?

The recent literature on cointegrated systems suggests some statistical procedures for addressing these issues. As noted in Granger (1981), several time series, which individually may be stationary only after differencing, may have linear combinations which are stationary *without* differencing. A group of time series with this property is termed 'cointegrated.' The necessary amount of differencing depends on the linear representation being considered, not the univariate properties of the data, *i.e.*, estimation of equation ⟨3⟩ with cointegrated data would be misspecified if the data were first-differenced and would omit important constraints if the data were used in level form. Moreover, if equation ⟨3⟩ is estimated using cointegrated data with a conventional method of correction for serial correlation, the parameter estimates will be biased.

Engle and Granger (1987) established how a cointegrated system can be represented in an error correction model (ECM), providing a framework for estimation of cointegrated systems. Short- and long-run dynamics are handled by imposing long-run equilibrium constraints on the short-run ECM. Hence, cointegration allows long-run equilibrium responses to be isolated from short-run dynamic behavior using a suitable linear combination of levels of variables (see Aoki, 1988).

Engle and Granger (1987) developed tests of the underlying assumption of cointegration and proposed a two-step estimator of cointegrated systems. First, the 'cointegrating regression'—such as equation ⟨3⟩—is estimated by ordinary least squares without the imposition of any dynamic structure. The residual term from this regression reflects the cointegrating linear relationship. Second, the ECM is estimated imposing the constraint of long-term equilibrium among levels of variables by including the cointegrating regression residuals. Engle and Yoo (1987) showed that forecasting gains exist by utilizing the Engle and Granger (1987) two-step estimator in the presence of cointegrated data.

To illustrate the error correction procedure, consider the following simplified model. Suppose time-series $x_t$ and $y_t$ are each first-difference stationary, that is, integrated of order one, I(1). Yet suppose the linear combination of these time series $u_t = y_t - \alpha x_t$ is level stationary, *i.e.*, I(0). Then time-series $y_t$ can be represented by the ECM:

$$\langle 4 \rangle \qquad \Delta y_t = \delta + \theta w_t + \gamma u_{t-1} + v_t,$$

where $v_t$ is white noise and $w_t$ is a vector of I(0) explanatory variables which likely includes *lagged changes* in $x_t$ and $y_t$. Variable $u_{t-1}$ represents deviations from the long-run equilibrium relationship; consequently, $\gamma$ is expected to be negative. This formulation has been referred to as the *restricted* ECM.

Hall (1986) suggests an alternative *unrestricted* ECM as well. In this approach, the restriction implied by the cointegrating regression (*i.e.*, $u_t = y_t - \alpha x_t$) is relaxed. Consequently, the error-correction term in equation ⟨4⟩, $u_{t-1}$, is replaced by the elements of the long-term relationship, $y_{t-1}$ and $x_{t-1}$:

$$\langle 5\rangle \qquad \Delta y_t = \delta + \theta w_t + \lambda y_{t-1} + \phi x_{t-1} + v_t.$$

Estimates of $\phi$ reveal the long-run effect of $x$ on $y$, where the $t$-statistics corresponding to $\phi$ are asymptotically valid; see also Engle *et al.* (1989).

## IV. Results

A necessary condition for a cointegrating relationship is that all time series must be of the same order of integration. Each individual time series (*i.e.*, the log of foreign-held US demand deposits in 1985 prices, home and foreign short-term interest rates, and the log of foreign income in 1985 prices) was tested for first-order integration using univariate Dickey–Fuller (DF) and Augmented Dickey–Fuller (ADF) tests. Stationarity in the first-differences of each time series could not be rejected at the 1 per cent significance level. Thus, estimation of equation ⟨3⟩ as the cointegrating relationship is a consistent model, as defined in Granger (1981).

Equation ⟨3⟩ was estimated as the cointegrating regression for each country with quarterly seasonal dummies and a linear time trend included. This removed deterministic components from the cointegrating residuals, which then represent the cointegrating linear relationship for use in the error correction structure.

Test statistics for cointegration, as developed by Engle and Granger (1987), are the Cointegrating Regression Durbin–Watson (CRDW), Dickey–Fuller (DF), and Augmented Dickey–Fuller (ADF). Since it was indeterminate *a priori* whether the cointegrating regression residuals were white noise, the ADF test included four lagged differences allowing for the possibility of fourth-order autocorrelation (see Wallis, 1972). Not surprisingly, most of the lagged differences were statistically insignificant suggesting that the CRDW and DF test statistics, assuming first-order autoregressive processes, were the appropriate ones, instead of the ADF statistics; nevertheless, we also reported the ADF statistics.

Table 1 reports that the null hypothesis of non-cointegration is rejected for each country with the exception of Japan based upon the CRDW and DF test statistics. The apparent presence of cointegration for five of the six countries implies a significant long-run equilibrium relationship governing the behavior of foreign-held dollar deposits as predicted by the money–services model of currency substitution for those five countries.

Estimates of the restricted ECM for each country, except Japan, using the seemingly unrelated regression (SUR) technique are reported in Table 2. Short-run I(0) explanatory variables (*i.e.*, $w_t$ in equation ⟨4⟩) included lagged changes in US and foreign interest rates, the lagged change in foreign real dollar income, and the lagged change in the real US demand deposits held by foreign residents. The cointegrating regression's long-run equilibrium behavior is imposed on each model by including the cointegrating regression residual in the restricted ECM.

TABLE 1. Cointegration test results.

Cointegrating regression:

$$\ln(M^D_{HF}/P_H) = \beta_0 - \beta_1 r_H + \beta_2 r_F + \beta_3 \ln Y_F + u$$

| | Canada | Italy | Japan | Switzerland | UK | W. Germany |
|---|---|---|---|---|---|---|
| CRDW | 1.390*** | 1.138*** | 0.467 | 1.645*** | 1.954*** | 1.859*** |
| DF | −4.698** | −4.051* | −2.326 | −5.295*** | −6.292*** | −5.987*** |
| ADF | −2.603 | −2.704 | −2.134 | −2.405 | −4.453*** | −1.378 |
| | (0) | (1) | (0) | (0) | (3) | (1) |

*Notes:* ***, **, and * denote statistical significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively. Appropriate critical values for our sample size and the $N=4$ case were obtained from Engle and Yoo (1987). The number of significant lagged differences are reported in parentheses under the reported ADF $t$-statistic.

TABLE 2. Restricted error correction model.

| | Canada | Italy | Switzerland | UK | W. Germany |
|---|---|---|---|---|---|
| Constant | −0.016 | 0.021 | −0.034 | 0.027 | 0.010 |
| | (−0.55) | (1.02) | (−0.88) | (0.80) | (0.28) |
| $\Delta r_H(-1)$ | 0.025 | 0.013 | 0.001 | 0.001 | −0.034 |
| | (0.85) | (0.82) | (0.03) | (0.03) | (−1.08) |
| $\Delta r_F(-1)$ | 0.003 | 0.005 | −0.036 | −0.023 | 0.078* |
| | (0.15) | (0.25) | (−0.82) | (−0.86) | (1.52) |
| $\Delta Y_F(-1)$ | 1.305 | −0.493 | 4.828* | −2.378*** | −0.939 |
| | (0.63) | (−0.55) | (1.30) | (−2.66) | (−1.18) |
| $\Delta M(-1)$ | −0.257** | −0.201* | −0.086 | −0.294** | 0.006 |
| | (−1.90) | (−1.32) | (−0.48) | (−2.27) | (0.04) |
| $u(-1)$ | −0.568*** | −0.319** | −0.748*** | −0.848*** | −0.844*** |
| | (−3.03) | (−1.80) | (−3.28) | (−3.86) | (−3.72) |
| SEE | 0.171 | 0.126 | 0.227 | 0.236 | 0.249 |
| DW | 2.10 | 2.17 | 2.04 | 2.14 | 2.13 |
| Degrees of freedom | 35 | 35 | 35 | 35 | 35 |

*Notes:* ***, **, and * denote statistical significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively, in one-tailed $t$-tests. (−1) denotes a one-period lag. Estimation is from 1978:IV to 1988:IV.

Since the error term represents deviations from the long-run equilibrium, the coefficient estimate of the error term is expected to be negative; when foreign deposits are above their long-run equilibrium level, downward pressure on deposits next period is expected. As reported in Table 2, estimated coefficients of the error terms have the anticipated sign and are statistically significant for all countries. Specifically, deviations from the long-term equilibrium relationship governing the foreign demand for dollar deposits is shown to affect the dynamic short-run foreign demand for home money.

The money–services model predicts a negative (positive) short-run response of foreign demand for home money to increases in home (foreign) interest rates. With the exception of Germany, Table 2 shows virtually no dynamic short-run currency-substitution behavior, as explained by lagged first-differences of the explanatory variables. In the context of our model, currency substitution appears to be fairly absent in the short term.

Following Hall (1986) and Engle *et al.* (1989), we estimated the error-correction model free from the cross-equation cointegrating regression restriction using SUR. While this unrestricted version of the error correction model removes the cointegrating regression's restriction concerning long-run behavior, it reveals currency substitution behavior relative to the restricted (former) model. Under the money–services approach we expect a negative (positive) coefficient estimate for the lagged *levels* of home (foreign) interest rates.

Table 3 reveals little evidence of dynamic *short-run* currency substitution,

TABLE 3. Unrestricted error correction model.

| | Canada | Italy | Switzerland | UK | W. Germany |
|---|---|---|---|---|---|
| Constant | 11.46*** | −0.472 | 5.634** | 4.199 | 5.944 |
| | (5.28) | (−0.31) | (2.27) | (1.27) | (1.31) |
| $\Delta r_H(-1)$ | 0.077*** | 0.036** | 0.003 | 0.030 | 0.061* |
| | (2.55) | (2.34) | (0.11) | (0.95) | (1.42) |
| $\Delta r_F(-1)$ | 0.011 | 0.011 | −0.066* | −0.013 | 0.076* |
| | (0.48) | (0.54) | (−1.51) | (−0.46) | (1.52) |
| $\Delta Y_F(-1)$ | 1.505 | −0.235 | 6.834** | −0.756 | −0.065 |
| | (0.76) | (−0.28) | (1.88) | (−0.78) | (−0.08) |
| $\Delta M(-1)$ | −0.262** | −0.364** | −0.171 | −0.258** | −0.060 |
| | (−2.23) | (−2.33) | (−1.06) | (−1.73) | (−0.38) |
| $r_H(-1)$ | −0.077** | −0.047*** | −0.023* | −0.062*** | −0.076* |
| | (−2.41) | (−3.64) | (−1.49) | (−2.64) | (−1.46) |
| $r_F(-1)$ | 0.038* | 0.017** | −0.007 | 0.042* | 0.066 |
| | (1.40) | (1.72) | (−0.21) | (1.64) | (1.16) |
| $Y_F(-1)$ | −1.25*** | 0.360 | −0.814 | 0.139 | −0.277 |
| | (−4.06) | (1.04) | (−1.07) | (0.19) | (−0.39) |
| $M(-1)$ | −0.674*** | −0.348** | −0.571*** | −0.776*** | −0.761*** |
| | (−4.93) | (−2.23) | (−3.41) | (−3.53) | (−3.70) |
| SEE | 0.142 | 0.115 | 0.234 | 0.235 | 0.232 |
| DW | 1.69 | 2.28 | 2.06 | 1.84 | 2.24 |
| Degrees of freedom | 32 | 32 | 32 | 32 | 32 |

*Notes:* ***, **, and * denote statistical significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively, in one-tailed *t*-tests. (−1) denotes a one-period lag. Estimation is from 1978:IV to 1988:IV.

consistent with Table 2. Only one of the coefficient estimates of the lagged change in the foreign interest rate was significant and of the anticipated positive sign (Germany). Several of the short-run lagged changes had coefficient estimates of the unanticipated sign and statistical significance.

Yet Table 3 reveals robust evidence of *long-run* currency-substitution behavior as predicted by the money–services model. Specifically, the estimated coefficient on the lagged level of the home country interest rate has the anticipated negative sign and is significant for all countries. The lagged foreign interest rate level coefficient estimate has the anticipated positive sign for four of five countries and is significant for three countries.

We also examined the presence of currency substitution in the long-run by joint tests of restrictions excluding the lagged home and foreign interest rate levels. The joint null of a zero coefficient on the lagged home interest rate level was rejected at the 1 per cent marginal significance level. The joint null of a zero coefficient on the lagged foreign interest rate level was rejected at the 12 per cent marginal significance level. The joint null of a zero coefficient on the lagged home *and* foreign interest rate levels was rejected at the 1 per cent marginal significance level.

Finally, we estimated the system jointly restricting the coefficients of the lagged home interest rate levels to be equal across countries *and* the coefficients of the lagged foreign interest rate levels to be equal across countries. As reported in Table 4, both the home and foreign interest rate lagged levels' coefficient estimates had the anticipated negative and positive signs, respectively, and were statistically different from zero at the 1 per cent and 5 per cent marginal significance levels, respectively. These results are consistent with the presence of long-run currency substitution affecting the foreign demand for US demand deposits, as suggested by the money–services model.

## V. Conclusion

This paper investigated short- and long-run currency-substitution behavior using the two-step estimator of cointegrated systems as developed by Engle and Granger (1987). The money–services model of currency substitution was employed in defining the null hypotheses concerning currency substitution. This model provides an optimizing framework for currency substitution based upon production theory.

An error correction model (ECM) was used to discern evidence of short- and long-run currency substitution. The restricted ECM suggests a long-run equilibrium relationship among the variables predicted by the money–services approach and that deviations from this relationship significantly affect the short-run foreign demand for US demand deposits. Evidence was presented in support of long-run currency substitution behavior by examining results from an unrestricted ECM. Specifically, increases in foreign interest rates tend to increase foreign demand for US deposits in the long run, as the opportunity cost to foreigners of using the transactions services of their local currency increases. The results are strongest for Italy, which is appealing since one might expect dollar substitution to be strongest in a country that has faced intermittent episodes of quite high inflation.

Our results provide little support for currency substitution as a significant short-term dynamic phenomenon. Hence, the short-run impact of currency

TABLE 4. Unrestricted error correction model with joint equality restrictions.

| | Canada | Italy | Switzerland | UK | W. Germany |
|---|---|---|---|---|---|
| Constant | 9.826*** | −0.754 | 7.634*** | 3.468* | 6.059* |
| | (5.10) | (−0.56) | (3.68) | (1.54) | (1.58) |
| $\Delta r_H(-1)$ | 0.059** | 0.034** | 0.014 | 0.013 | 0.035 |
| | (2.24) | (2.28) | (0.48) | (0.46) | (1.21) |
| $\Delta r_F(-1)$ | 0.014 | 0.007 | −0.072** | −0.004 | 0.074* |
| | (0.66) | (0.35) | (−1.74) | (−0.16) | (1.51) |
| $\Delta Y_F(-1)$ | 1.965 | −0.217 | 7.101** | −0.915 | −0.097 |
| | (1.05) | (−0.26) | (1.91) | (−0.96) | (−0.12) |
| $\Delta M(-1)$ | −0.206** | −0.356*** | −0.166 | −0.298** | −0.045 |
| | (−1.75) | (−2.48) | (−1.00) | (−2.12) | (−0.29) |
| $r_H(-1)$ | −0.041*** | −0.041*** | −0.041*** | −0.041*** | −0.041*** |
| | (−4.79) | (−4.79) | (−4.79) | (−4.79) | (−4.79) |
| $r_F(-1)$ | 0.015** | 0.015** | 0.015** | 0.015** | 0.015** |
| | (2.07) | (2.07) | (2.07) | (2.07) | (2.07) |
| $Y_F(-1)$ | −0.955*** | 0.390* | −1.42** | 0.148 | −0.275 |
| | (−3.82) | (1.34) | (−2.31) | (0.28) | (−0.47) |
| $M(-1)$ | −0.698*** | −0.330*** | −0.571*** | −0.646*** | −0.791*** |
| | (−5.10) | (−2.77) | (−3.48) | (−3.45) | (−3.96) |
| SEE | 0.146 | 0.116 | 0.237 | 0.239 | 0.237 |
| DW | 1.75 | 2.28 | 2.00 | 1.95 | 2.15 |
| Degrees of freedom | 32 | 32 | 32 | 32 | 32 |

*Notes:* ***, **, and * denote statistical significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively, in one-tailed $t$-tests. $(-1)$ denotes a one-period lag. Estimation is from 1978:IV to 1988:IV.

substitution on monetary policy is likely to be of little importance, which is a plausible finding. But the evidence presented here suggests that currency substitution may be an important factor influencing long-term monetary policy independence.

## Notes

1. Quarterly interest rates and real incomes are from the IMF's *International Financial Statistics* (tape), 1978:II–1988:IV. The short-term interest rate used was the three-month Treasury bill rate (series '6 0c') when available; in lieu of that, the three-month call money rate (series '6 0b') was used. Real income was measured in 1985 prices using real gross national product (series '9r9a') when available; in lieu of that, real gross domestic product (series '9p9b') was used.

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