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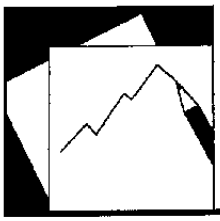

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Is Foreign Debt Portfolio Management Efficient in Emerging Economies?

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Is Foreign Debt Portfolio Management Efficient in Emerging Economies?

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Abstract

This paper develops a simple model of foreign debt portfolio management. The model suggests that, under mild conditions, the currency composition of a country's foreign debt portfolio is responsive to exchange rate movements. Empirical evidence is provided for a panel of 14 emerging economies in the period 1970-98. Attention is focused on the stocks of foreign liabilities denominated in U.S. dollars, deutsche marks (DM), Japanese yen, and Swiss francs. The results of the empirical analysis show that foreign debt portfolio management has been sub-optimal in the countries under examination. In these countries, the currency composition of foreign debt has not reflected a substitution effect away from the currencies that have appreciated over time vis-à-vis the U.S. dollar.

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I. Introduction

Foreign debt portfolio management has become a key issue for fiscal policymakers in developing countries since the Debt Crisis of the 1980s. More recently, financial crises in Asia, Latin America, and Russia have shown that governments should manage their foreign exposure so as to prevent a mismatch in the currency composition of their foreign assets and liabilities. There is now ample evidence (Corsetti and others, 1998; IMF, 1998a) that the Asian crisis was precipitated by volatility in the U.S. dollar-yen exchange rate, which drove a wedge between the region's export earnings, denominated primarily in U.S. dollars, and its financial liabilities, increasingly denominated in Japanese yen. According to the IMF (1998b), foreign debt burdens increased significantly in Asia due to the region's large and unhedged exposure to the yen, which appreciated steeply against the U.S. dollar in the first half of the 1990s. Likewise, it can be argued that the currency composition and maturity profile of Mexico's foreign debt, rather than its volume, were key elements undermining investor confidence and leading to the peso meltdown of 1994–95 (Sachs and others, 1995).

Asset pricing and portfolio selection models offer interesting insights into optimal foreign debt portfolio management. These models suggest that the portfolio share of an asset should be proportional to its rate of return, and that investors should hedge against risk and unexpected movements in liability values. An interesting application of these models is foreign debt portfolio management, given the currency composition of a country's foreign debt, and the rate of return of assets denominated in different currencies. Intuitively, other things equal, changes in the share of a country's foreign debt denominated in a given currency are associated with that currency's exchange rate movements relative to other currencies in which the country's foreign debt may be denominated. If, for instance, a country's foreign debt portfolio is denominated in U.S. dollars and Japanese yen, an appreciation of the yen vis-à-vis the U.S. dollar leads to an increase in the U.S. dollar-value of the share of debt denominated in yen. Optimal portfolio management dictates in this case that the *volume* of yen-denominated debt should be reduced so as to offset the impact of the appreciation of the yen on the U.S. dollar-*value* of the yen-denominated debt. As a result of this substitution effect, the U.S. dollar-value of the share of debt denominated in yen should not be affected by an appreciation of the yen relative to the U.S. dollar.

Obviously, the overall currency composition of a country's foreign debt portfolio may exhibit some rigidity, which reduces the efficiency of portfolio management instruments and leads to sub-optimal debt portfolio management. Rigidity may be attributed to such factors as the country's foreign trade and financial transactions, the overall supply of international credit, and capital flows. For instance, an indebted country that trades predominantly with the United States is likely to have most of its financial transactions with the United States, to peg its currency to the U.S. dollar, and hence, to have a sizeable share of its foreign debt denominated in U.S. dollars. In this case, this country may be unable, or unwilling, to alter the currency composition of its external debt according to cross-currency movements. The proportion of official lending in total foreign liabilities, often denominated in the lender's currency, is another reason why portfolio managers may be unable to alter the currency composition of foreign debt in response to adverse exchange rate movements.

In this paper, we test whether the currency composition of different countries' foreign liabilities has offset adverse exchange rate movements. We use the recently developed panel cointegration techniques that allow for heterogeneous cross-country dynamics and endogenous regressors. We apply the panel unit root test developed by Im, Pesaran, and Shin (1997; hereinafter referred to as IPS), to test the null hypothesis of no cointegration in panel data using the procedure suggested by Pedroni (1999) and Kao (1999). We also estimate the long-run relationships between each of the debt shares and the corresponding exchange rates using the panel dynamic OLS estimator proposed by Kao and Chiang (1998) and the dynamic fixed-effects estimator (see Pesaran and others, 1999), and then test whether exchange rate movements Granger-cause changes in debt shares. Our panel includes 14 emerging economies over the period 1970–98. The intuition is that, if changes in debt shares cointegrate with exchange rate movements, a stable long-run relationship is expected to exist between changes in the currency composition of a country's foreign debt portfolio and exchange rate movements. Debt portfolio management can be deemed optimal if movements in exchange rates do not Granger-cause changes in debt shares denominated in the corresponding currencies. In other words, foreign debt portfolios are managed optimally if, for instance, the U.S. dollar-value of the share of debt denominated in yen is not affected, in the Granger-causality sense, by adverse U.S. dollar-yen exchange rate movements. Preliminary tests are also carried out by comparing the correlation coefficients between changes in debt shares and exchange rate movements and their relative volatility.

The paper is structured as follows. Section II provides a brief review of the foreign debt management literature. Section III develops a simple foreign debt management model from which the basic testable hypotheses can be derived. Section IV discusses the econometric methodology for testing for unit roots and cointegration in dynamic panels. Section V presents the data and preliminary findings. Section VI reports the results of the dynamic panel analysis. Section VII discusses the empirical findings, and Section VIII concludes.

II. The Literature

The literature on foreign debt management in the developing world was motivated to a great extent by the Debt Crisis of the 1980s. In broad terms, the early literature focused on the debt transfer problem (Sachs, 1988). Accordingly, an indebted country could reduce its external exposure by managing its import and export flows so as to generate large enough trade surpluses to offset the costs of debt service. Limited hedging opportunities, shallow domestic capital markets, undiversified trade patterns, and adverse terms of trade movements were often highlighted as key factors explaining the difficulties in debt management facing most developing countries in the 1980s (Cooper, 1992; Fry, 1992; Dooley, 1995). The diversification of exports was the key long-term policy response advocated in this strand of literature. In the short run, international competitiveness and exports could be boosted chiefly through nominal devaluations (Edwards and Larrain, 1989; Williamson, 1990), rather than productivity gains. Against a background of unfavorable price-wage dynamics, these devaluations fueled inflationary pressures in most indebted countries in the 1980s.

Subsequently, in the late 1980s and early 1990s, the literature reflected the second phase of the Debt Crisis, in which debt rescheduling became a prominent feature of foreign liability management. Renewed access to international capital markets for most indebted countries ushered in several types of debt swap operations, alleviated important external liquidity and solvency constraints, encouraged private lending, and reduced the share of public and publicly guaranteed liabilities in total foreign debt stocks. Together with comprehensive trade and investment liberalization, these factors facilitated hedging against foreign exchange risk and gave indebted countries more leeway in foreign debt management.

More recently, the debt management literature has focused on the consolidation of market-oriented reform and macroeconomic stability in indebted countries, particularly in Latin America and Central and Eastern Europe. These countries have progressively liberalized their trade, investment and international payments regimes, phased out capital controls, and facilitated international capital movements. In this more liberal policy environment, foreign debt management has become a key element of fiscal policymaking, particularly in terms of exchange and interest rate variability, as well as volatility in capital flows (Cassard and Folkers-Landau, 1997; IMF, 1998b). In recent years, governments have played a prominent role in financial intermediation and foreign debt management. In this respect, Dooley (2000) shows that minimizing debt service costs may be inefficient for developing country governments because such a policy may increase default risks and therefore borrowing costs.

Given the importance of foreign debt portfolio management in developing countries, the dearth of empirical research on this issue is surprising. One of the few studies on the optimal currency composition of foreign debt stocks was carried out in 1992 by S. Claessens using monthly Brazilian and Mexican data over the 1970s and 1980s to estimate the relationship between total exports and the effective cost of borrowing in three currencies (U.S. dollar, Japanese yen, and the deutsche mark). The currency composition of a country's net liabilities is estimated to minimize the domestic currency variability of export earnings net of foreign debt service. The results show that Mexico and Brazil could have lowered their external exposure by continuously altering their debt portfolios. The author concludes that the low correlations between borrowing costs and export prices render the currency composition of foreign debt an imperfect hedging tool against shocks in external prices.

Demirgüç-Kunt and Detriagache (1994) provide a descriptive analysis of the currency composition of long-term foreign debt for nine highly indebted countries in the 1980s and estimate its impact on interest spreads, using panel data analysis. Their findings show that lower spreads can be explained by a large share of official lending in total foreign borrowing. The results also suggest that spreads are not significantly affected by the rising share of floating interest-bearing debt, relative to fixed interest rates that prevailed in the 1970s and 1980s.

III. The Model

Let there be n currencies such that a country's total foreign sovereign debt at time t can be denominated in any of these n currencies. Let currency n be used as a numeraire such that

exchange rates (e_k, for $k = 1,...,n-1$) can be defined as the value of currency k per unit of currency n. The policymaker's objective is to minimize the total value of the foreign liabilities in his/her portfolio, defined as $C_t(A_t; D_{kt})$, where A_t is total foreign assets and D_{kt} is foreign debt denominated in currency k. Assets and liabilities are additively separable. Let $C_k < 0$ and $C_{kk} > 0$, where $C_k = \frac{\partial C_t}{\partial D_{kt}}$ and $C_{kk} = \frac{\partial^2 C_t}{\partial D_{kt}^2}$, such that C_t admits a minimum in D_{kt}.[2] In addition, let there be exchange rate uncertainty such that the value of foreign debt, denominated in currency n, is affected by unpredictable cross-exchange rate movements. The policymaker's problem can be formalized as:[3]

(P) $\min_{D_{kt}} C_t(A_t; D_{kt})$,

s.t. $D_t = \sum_{k=1}^{n} \frac{D_{kt}}{E_t e_{kt}}$,

where E_t is the expectation operator.

Standard manipulation of the first-order conditions for cost minimization yields:

$$\lambda_t = C_1 E_t e_{1t} = \ldots = C_n. \qquad (1)$$

By equation (1), it follows that: $\frac{C_1}{C_n} = \frac{1}{E_t e_{1t}}$, ... , $\frac{C_{n-1}}{C_n} = \frac{1}{E_t e_{n-1,t}}$. In particular, if currency k is expected to appreciate with respect to currency n ($E_t e_{kt}$ falls), the impact of k-denominated debt on portfolio C rises relative to that of n. Because $C_k < 0$, the share of foreign debt denominated in k, D_{kt}, falls. As a result, the expected appreciation of a given currency k (relative to currency n) implies a fall in the share of total debt denominated in k (relative to n).

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[2] Most optimal debt management models follow the tradition of Tobin (1963) and focus on how to minimize the interest cost of domestic debt. See, for example, Boothe and Reid (1992) for a cost-minimization model where emphasis is placed on domestic optimal debt management in small open economies. See also Dooley (2000) for a sovereign debt portfolio management model with default risk.

[3] Although we consider the utility maximizer to be a policymaker, foreign debt portfolio management can in principle be carried out by the government or by private-sector borrowers. We opted for treating the optimizer as a policymaker to simplify the model in line with the data constraints discussed in the empirical section below. We nevertheless agree that different portfolio managers face different constraints and pursue different, often conflicting objectives, which will not be discussed in detail in this paper. In addition, for simplicity, assume that there is no lagged adjustment in portfolio management and optimization is instantaneous.

Letting $C_t = \frac{1}{2}\left[A_t - \sum_{k=1}^{n} \left(\frac{D_{kt}}{e_{kt}} \right)^2 \right]$, for example, and assuming $E_t e_{kt} = e_{kt}$, it follows from equation (1) that $\frac{D_{kt}}{D_{nt}} = e_{kt}$. As a result, an appreciation of k with respect to n (a fall in e_{kt}) leads to a fall in the share of k-denominated debt, relative to the debt share denominated in n. A change in the volume of debt denominated in k relative to n offsets the appreciation of k to keep the n-value of the debt portfolio constant.

IV. Econometric Methodology

A number of unit root and cointegration tests have been developed over the years in the time-series literature; nevertheless, there are few such tests for panel data. The growing interest in unit root and cointegration tests for panel data has been motivated, at least in part, by the well-known deficiencies of cointegration testing based on time series alone. It is argued that more thorough analyses of the unit root and cointegration properties of the data can be made by combining information derived from the time-series dimension of the data set and that obtained from its cross-sectional dimension, especially when the time series available for the variables under examination are not long enough. The estimation of dynamic panel data has yielded interesting results (Banerjee, 1999).

A. Testing for Unit Roots in Dynamic Panels

As in time-series analysis, the first step in the estimation of dynamic panels is to test whether the variables at hand contain unit roots. Recently, IPS (1997) have proposed unit root tests for heterogeneous panels that are more powerful than the alternative tests developed by Levin and Lin (1993) and Quah (1994).[4] The IPS tests allow for the heterogeneity of dynamics and error variances across groups in the panel. Furthermore, the IPS tests have better small sample properties since their asymptotic validity only requires N/T$\rightarrow k$ (k is any finite positive constant) when both N (cross-sectional dimension of the panel) and T (time periods) tend to infinity, relative to the more stringent condition that N/T $\rightarrow 0$, required for the Levin and Lin (1993) test. Consider the standard ADF equation in a dynamic panel framework:

$$\Delta e_{it} = \alpha_i + \beta_i e_{i,t-1} + \sum_{j=1}^{p_i} \delta_{ij} \Delta e_{i,t-j} + \upsilon_{it}, \qquad (2)$$

where α_i are the group intercepts, e_{it} is a stochastic process observed over N cross-sections and T time periods, δ_{ij} are the parameters associated with the pth-order augmentation which take into account any possible serial correlation across groups, and υ_{it} are the disturbance terms,

[4] In a Monte Carlo simulation, IPS demonstrate better finite sample performance of the $\Psi_{\bar{t}}$ in relation to the Levin-Lin test, as discussed below.

which are assumed to be independently distributed with zero mean and finite heterogeneous variance, σ_i^2. Also, $i = 1,..., N$; $j = 1, ..., p$; and $t = 1, ..., T$.

The null hypothesis of unit roots across all groups in the panel (H_0: $\beta_i = 0$ for all i) is tested against the alternative hypothesis that allows some of the individual series to have unit roots (H_1: $\beta_i < 0, i = 1,..., N, \beta_i = 0, i = N_1+1,..., N$). By allowing β_i to differ across groups, this formulation of the alternative hypothesis is more general than the homogeneous alternative hypothesis that $\beta_i < 0$ for all i. In order to test $\beta_i = 0$, the t-bar statistic across groups ($\Psi_{\bar{t}}$) is defined as:

$$\Psi_{\bar{t}} = \frac{\sqrt{N}(\bar{t}_{N,T} - E[\bar{t}_{N,T}(p,0)])}{\sqrt{Var(\bar{t}_{N,T})}}, \tag{3}$$

where $\bar{t}_{N,T} = (1/N)\sum_{i=1}^{N} t_i$, t_i is the t-statistic for the OLS estimate of β in equation (2) for the t-th unit of the cross section, and E is the expectation operator.

The mean and the variance of $[t_{N,T}(p,0)|\beta_i = 0]$ are tabulated in IPS (1997) for different time-series dimensions and lag orders, and p_i for each cross section. $\Psi_{\bar{t}}$ can be compared with critical values for a one-sided $N(0,1)$ distribution.[5]

B. Testing for Cointegration in Dynamic Panels

If the relevant variables in the panel are non-stationary, the system can be tested for cointegration.[6] The literature on dynamic panels provides two different cointegration tests. Both tests are residual-based. The first test, proposed by Pedroni (1995, 1999) and Kao (1999), uses residuals derived from the panel analogue of the traditional Engle and Granger (1987) two-step regression to construct the test statistics. The second test for cointegration in dynamic panels was developed by McCoskey and Kao (1998) and has its analogue in the time-series literature (i.e., Shin, 1994; and Kwiatowski and others, 1992). A clear distinction between the two types of panel cointegration tests is the null hypothesis: whereas Pedroni's test takes no cointegration as the null hypothesis, McCoskey and Kao's test takes the null of cointegration. In what follows, we discuss the residual-based test for cointegration in dynamic panels proposed by Pedroni (1999) and Kao (1999).

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[5] IPS proposed LM-bar tests for unit roots. The t-bar test is shown to perform better than the LM-bar in small samples. An important feature of these tests is that their power is favorably affected by a rise in T compared to an equivalent increase in N. This feature is important in our data set where T is larger than N.

[6] It is not necessary that the variables of interest be non-stationary in order to estimate the long-run relationship. The pooled mean group (PMG) estimator of dynamic heterogeneous panels developed by Pesaran and others (1999) consider both the case where the regressors are stationary and the case where they follow unit root processes.

Consider the following panel regression:

$$y_{it} = \alpha_i + \beta_i x'_{it} + e_{it} , \tag{4}$$

where $\beta_i = (\beta_{1i}, \beta_{2i}, ..., \beta_{Mi})'$, $x_{it} = (x_{1it}, x_{2it}, ..., x_{Mit})'$, $t = 1, ..., T$, and $i = 1, ..., N$.

Notice that the slope coefficients β_i and parameter α_i (fixed-effects parameter) are allowed to vary across individual groups.[7] Therefore, the above formulation allows for considerable heterogeneity in the panel. The test uses the residuals from the cointegration regression given by equation (4). The remainder of the test is analogous to the IPS test with equation (2) being estimated using the estimated residuals of equation (4). To test the null hypothesis of no cointegration, the *t*-bar statistic is computed in the form of equation (3) based on the values of the mean and the variance of $[t_{N,T}(p,0)|\beta_i = 0]$, tabulated by Pedroni (1999).[8]

C. Estimating the Cointegrating Vectors in Panels

Once the null hypothesis of no cointegration is rejected, the coefficients of the long-run relationships can be estimated for the dynamic panel using several methods, such as the Pooled Mean Group estimator developed by Pesaran and others (1999) and the Fully Modified estimator developed by Pedroni (1996).[9] Two methods are used in this paper: dynamic OLS (DOLS) and dynamic fixed effects (DFE) for panel data.

The DOLS estimator, proposed by Kao and Chiang (1998), is based on the Stock and Watson (1993) estimator for time series. The DOLS procedure involves running the following regression:

$$y_{it} = \alpha_i + \beta_i x'_{it} + \sum_{j=1}^{p_1} c_{ij} \Delta x_{i,t-j} + \sum_{j=1}^{p_1} \gamma_{ij} \Delta x_{i,t+j} + e_{it} , \tag{5}$$

where $t = 1, ..., T$ and $i = 1, ..., N$.

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[7] For some applications, a deterministic time trend, which is specific to individual groups of the panel, may be included in equation (4).

[8] Pedroni (1999) proposes seven panel cointegration statistics, four based on pooling along the within-dimension and three based on pooling along the between-dimension. The one that we utilize here belongs to the latter category and is based on a parametric ADF test.

[9] In the time-series literature, the Fully Modified estimators were first proposed by Phillips and Hansen (1990).

Notice that equation (5) is an extension of equation (4) where lags and leads of $\Delta x_{i,t}$ are included in the cointegrating regressions in order to produce asymptotically unbiased estimators and to avoid the likely problem of estimating nuisance parameters. By estimating equation (5), it is possible to construct asymptotically valid test statistics and also to estimate the long-run relationship where the coefficients of x'_{it} are the cointegrating parameters. Thus, the causal relationship between any of the x'_{it} and y_{it} can be tested using an *F*-test. For example, x_{1it} Granger-causes y_{it} if the null hypothesis that $\beta_{1i} = c_{1ij} = \gamma_{1ij} = 0$ is rejected. Kao and others (1999) argue in favour of the DOLS estimator in estimating the cointegrated panel regressions. The Monte Carlo simulations presented in Kao and Chiang (1998) show that the DOLS estimator outperforms both OLS and FM estimators.[10]

The other method we use in order to estimate the long-run relationship is the dynamic fixed-effect estimator (DFE) for panels (see Pesaran and others, 1999). The DFE estimator is based on an autoregressive distributed lag (ADRL) model in time-series analysis (see Pesaran and Shin, 1999). The DFE technique involves estimating the following ADRL model:

$$\Delta y_{it} = \sigma_i d'_{it} + \lambda_i y_{i,t-1} + \beta_i x'_{i,t-1} + \sum_{j=1}^{p_1} \omega_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{p_1} \phi_{ij} \Delta x'_{i,t-j} + \upsilon_{it} , \qquad (6)$$

where d'_{it} is a vector of time-invariant regressors (intercepts and time trends, for instance).

The estimate of the long-run coefficient of $x'_{i,t}$ is given by $\hat{\theta}_{x'} = -\frac{\hat{\beta}_i}{\hat{\lambda}_i}$, where $\hat{\beta}_i$ and $\hat{\lambda}_i$ are the DFE estimators of β_i and λ_i in equation (6). As in panel DOLS, the causal relationship can be tested using a standard *F*-test. For example, x_{1it} Granger-causes y_{it} if the null hypothesis that $\beta_{1i} = \phi_{1ij} = 0$ is rejected.

V. Data and Preliminary Findings

The foreign debt data used in this paper are available from the World Bank's Global Development Finance (2000). Due to lack of data on the currency composition of foreign debt for most developing countries, our sample comprises 14 emerging market economies (Argentina, Brazil, Chile, Colombia, Egypt, India, Indonesia, Korea (South), Malaysia, Mexico, Thailand, Philippines, Turkey, and Venezuela) for which continuous annual data for 1970–98 are available. The exchange rate series are obtained from the World Bank's World Development Indicators (2000). Because for most countries the largest share of long-term foreign debt is

[10] The main difficulty of using the DOLS estimator is to choose the number of lags and leads. FM estimators suffer from more serious problems, however. For example, the correction terms depend primarily on the preliminary estimator which may be biased in finite samples.

denominated in U.S. dollars, followed by Japanese yen, deutsche mark, and Swiss francs, attention is focused here on these four currencies. The U.S. dollar is used as the numeraire currency, with respect to which debt shares and exchange rates are defined. The foreign debt data used here do not allow for distinguishing the portfolio managers between the government and private-sector borrowers and the foreign debt stock by debt instrument (sovereign bonds, corporate securities, accounts payable, and derivatives, among others). Exchange rate-indexed debt is treated as foreign debt and liabilities denominated in a country's own currency are treated as domestic debt.

Using cross-section data analysis, where the data set was constructed by averaging the exchange rates and debt shares for each time period over the sample of countries under examination, a preliminary visual test of optimality in foreign debt portfolio management is provided in Figure 1. Debt shares and exchange rates are plotted in the vertical axes. Should portfolio management be optimal, as suggested above, an appreciation of a given currency should lead to a reduction in the *volume* of foreign debt denominated in that currency to keep its U.S. dollar *value*-constant. Because $\overline{D}_{kt} = \frac{s_{kt}}{e_{kt}}$, where $s_{kt} = \frac{D_{kt}}{D_{nt}}$, an appreciation in currency k, vis-à-vis the U.S. dollar (a fall in e_{kt}) leads to an increase in $\overline{D}_{kt}$, the U.S. dollar-value of s_{kt} (the share of k-denominated foreign debt, relative to the debt denominated in U.S. dollars), unless s_{kt} falls at the same time to compensate for the currency appreciation. In Figure 1, optimal portfolio management would require the U.S. dollar-denominated debt share schedule to be flat irrespective of the downward and upward movements in the exchange rate curve.

However, the figure suggests that an appreciation of the three currencies relative to the U.S. dollar increases the U.S. dollar-value of foreign debt denominated in these three currencies. This is particularly true in the case of the Japanese yen: a persistent appreciation of the yen vis-à-vis the U.S. dollar since 1970 has led to a persistent increase in the U.S. dollar-value of yen-denominated debt with respect to the total debt denominated in U.S. dollars, without an offsetting reduction in the volume of foreign debt denominated in yen. Figure 1 shows that, between 1970 and 1998, the rise in the U.S. dollar-value of the yen-denominated debt has been impressive: from 5 per cent to roughly 45 percent of the U.S. dollar-denominated debt. In this 28-year span, the U.S. dollar-value of the share of deutsche mark-denominated debt has varied between approximately 5 and 20 per cent of the dollar-denominated debt in the sample of countries under examination. The U.S. dollar-value of the Swiss franc-denominated debt has fluctuated considerably between roughly 15 and 35 percent of the debt denominated in U.S. dollars.

When the two lines move together, a downward (upward) slope indicates an appreciation (depreciation) of the exchange rate relative to the U.S. dollar, and a fall (rise) in the U.S. dollar-value of the debt denominated in that particular currency. In this case, a downward (upward) move implies a reduction (increase) in the volume of debt denominated in that particular currency to compensate for an increase (fall) in its U.S. dollar-value as a result of the currency

appreciation. This simultaneous move is particularly evident, in Figure 1, in the case of the Swiss franc and the deutsche mark from 1970 to 1976/77.

Using time-series analysis, additional informal tests can be carried out, consisting of comparing (i) the standard deviations of changes in the exchange rate and debt share series, and (ii) the correlations between the two series. When foreign debt management is optimal, the U.S. dollar-value of the debt share denominated in a given currency is not affected by exchange rate movements and, therefore, the correlation between changes in debt shares and exchange rates should be zero. The U.S. dollar-value of debt shares should also be less volatile than exchange rate movements. In this case, the ratio of the standard deviation of the U.S. dollar-value of debt shares to the standard deviation of the exchange rates should be less than 1.

The results of the preliminary descriptive tests are reported in Table 1. With respect to the correlation tests, the results provide mixed evidence. The null hypothesis of no correlation is rejected in five countries (Argentina, South Korea, India, Philippines, and Venezuela). In the other nine countries, there are mixed signals where the no-correlation hypothesis cannot be rejected for all currencies. On the other hand, the results of the volatility tests provide overwhelming evidence to support the assumption of optimal foreign debt management in our sample of indebted countries. In the case of Turkey, there is mixed evidence where the null hypothesis cannot be rejected for all currencies.

VI. Dynamic Panel Results

Because the time span in the panel to be estimated is short, panel data analysis offers useful ways to increase the power of statistical tests. To identify possible unit roots, the IPS (1997) test described above is performed for the levels and first differences of (the logarithm of) each of the variables of interest. The degree of augmentation (p) was determined following the general-to-specific procedure recommended by Campbell and Perron (1991). We started with four lags and estimated equation (2) with and without a time trend. In both cases, the null hypothesis is that the variable in question contains a unit root against the alternative hypothesis that the variable is stationary.

The results of the panel unit root tests are reported in Table 2. With two exceptions, the t-bar statistics suggest that the six variables are non-stationary. The hypothesis of a unit root in (the logarithm of) the three debt-share variables cannot be rejected in any case except for the debt share denominated in deutsche mark without a time trend. Similar results are reported in the case of the three exchange rates, where the t-bar statistic is greater than the critical value only in the case of the deutsche mark/U.S. dollar exchange rate ($e_{DM/US\$}$) with a time trend. By contrast, when first differences are taken, the results of the panel unit root tests strongly indicate rejection of the null hypothesis in all cases. Thus, our findings show that all six variables are I(1).

Given the results of the unit root tests, we proceeded to the panel cointegration tests. Based on the panel cointegration test developed by Pedroni (1999) and Kao (1999), we defined three

models, where each debt share is regressed on the three exchange rates. The *t*-bar statistics for the panel cointegration tests are reported in Table 3. The null hypothesis is that there is no cointegrating vector and the alternative hypothesis is that there is one cointegrating vector. The results provide strong evidence of the existence of long-run relationships between the debt shares and the corresponding exchange rates in the three models when a time trend is included in the estimating equations. In the three models, the null hypothesis is rejected at the 1 percent level of significance. Furthermore, in the other three models that exclude a time trend, the cointegration results are supportive of the hypothesis of a long-run relationship between each of the debt shares and the exchange rates, except in the case of the debt share denominated in Swiss francs, where the null hypothesis cannot be rejected.

Overall, since the cointegration results are supportive of the hypothesis of a stable long-run relationship between each of the debt shares and the corresponding exchange rates, we can estimate the cointegrating vectors using the DOLS and DFE procedures described above, and then test for causality. In principle, optimal foreign debt portfolio management requires that the volume of foreign debt be adjusted in response to adverse exchange rate movements to keep the U.S. dollar-value of the debt portfolio constant. As a result, the exchange rate movements should not Granger-cause changes in debt shares.

The estimated long-run relationships using the DOLS procedure are reported in Table 4. There is strong evidence that foreign debt management is sub-optimal with respect to the three different debt shares. The first row of Table 4 shows that the Japanese yen/U.S. dollar exchange rate ($e_{Yen/US\$}$) has a significant impact on the debt share denominated in Japanese yen ($Debt_Y$). The *F*-test of the exclusion of the dynamic terms and the level of $e_{Yen/US\$}$ in the $Debt_Y$ model shows that $e_{Yen/US\$}$ Granger-causes $Debt_Y$. The same results are reported in the other two cases where the debt share denominated in DM ($Debt_M$) is affected by the deutsche mark/U.S. dollar exchange rate ($e_{DM/US\$}$), and the effect of the Swiss franc/U.S. dollar exchange rate ($e_{SFr/US\$}$) on the debt share dominated in Swiss francs ($Debt_{SF}$) is significant at the 1 percent level of significance. The *F*-statistics also show that $e_{DM/US\$}$ causes $Debt_M$, whereas $Debt_{SF}$ is caused by $e_{SFr/US\$}$.

To test the robustness of the results reported above, we re-estimated the cointegrating vectors using the DFE estimator and tested for temporal causality. The results, reported in Table 5, reinforce the previous findings that foreign debt portfolio management is sub-optimal in the indebted economies in the panel. Based on standard *F*-tests, the null hypothesis of no-causation is strongly rejected in two cases out of three. The *F*-statistics show that $e_{Yen/US\$}$ Granger-causes $Debt_Y$, whereas $e_{DM/US\$}$ Granger-causes $Debt_M$ at the 1 percent level of significance. Only in the case of the debt share denominated in Swiss francs do the results based on the DFE estimator contradict the previous finding that $e_{SFr/US\$}$ Granger-causes $Debt_{SF}$.

VII. Discussion

The findings reported above show that exchange rate movements Granger-cause changes in debt shares. This means that foreign debt portfolio management is sub-optimal. Sub-optimality may result from some rigidity in the currency composition of a country's foreign debt stock. This may be attributed to a variety of factors. For instance, limited access to, or availability of, instruments to hedge against exchange rate risk may not allow for offsetting changes in the composition of debt portfolios in response to adverse changes in exchange rates. With undiversified trade and financial transaction patterns, a considerable share of the country's foreign assets and liabilities may be denominated in the same currency. Countries may therefore be unable, or unwilling, to alter the composition of their foreign debt portfolios in response to adverse movements in exchange rates.

Sub-optimality in foreign debt management can also be attributed to the term-structure of international lending contracts. A high share of foreign assistance, such as grants and concessional lending, in a country's total foreign borrowing may lead to sub-optimal foreign debt management (Li, 1992). In addition, a high share of public or publicly-guaranteed liabilities in total foreign debt may discourage efficient debt management by private sector borrowers. Access to foreign exchange by private sector firms may also be limited, thereby preventing changes in the composition of private sector debt portfolios.

An important policy question in efficient foreign debt portfolio management is whether the government in developing countries should act as a debt portfolio manager and hence participate actively in financial intermediation. This is because, among other things, emerging markets often lack the financial markets needed for efficient private sector-led portfolio management, such as a liquid market for long-term, fixed-interest government debt to be used as a hedge against exchange rate risk, for instance. Although there are welfare-improving roles for a developing country's government as a debt portfolio manager, its participation in financial markets may increase corporate and sovereign default risks and therefore portfolio management costs (Dooley, 2000). In the context of this paper, the costs and benefits of foreign debt portfolio management need to be weighed against the default risks that may change in different currencies in which a country's foreign debt may be denominated.

VIII. Conclusion

The Debt Crisis of the 1980s highlighted the need for capital-importing countries to manage their foreign liabilities efficiently so as to preserve their intertemporal solvency. A crucial aspect of optimal foreign debt portfolio management is the currency composition of debt portfolios. This is because exchange rate movements, which are exogenous to the debt portfolio manager, tend to alter the relative value of foreign liabilities and the shares of the country's foreign debt denominated in different currencies. The recent experience of several Asian countries suggest that these exchange rate swings may exacerbate international solvency problems even without significant debt accumulation. In this case, although many causes of balance of payments crises may be deemed to be home-grown, developing and emerging economies have also been affected adversely by volatility in exchange and interest rates.

This paper developed a simple model of foreign debt portfolio management and used the recently developed dynamic panel data analysis to test for the existence of a stable long-run relationship between the currency composition of a country's foreign debt portfolio and exchange rate movements. Attention is focused on a sample of 14 emerging market economies in the period 1970–98. The results reported here suggest that foreign debt portfolios have been managed sub-optimally in the countries under examination in the sense that adverse exchange rate movements have not been offset by a reduction in the volume of debt denominated in the appreciating currency. This is particularly true in the case of the yen-denominated foreign debt. The persistent appreciation of the yen vis-à-vis the U.S. dollar has not been translated into a reduction in the volume of yen-denominated debt, so as to prevent the increase in the U.S. dollar-value of the debt denominated in Japanese yen.

In principle, the currency composition of foreign debt stocks can be deemed to be an effective tool for foreign debt portfolio management. However, the results reported here are suggestive of the inability of foreign debt portfolio managers to adjust the currency composition of foreign debt portfolios in line with exchange rate movements. These findings may be attributed to some rigidity in the currency composition of foreign debt, given such factors as a country's trade and investment patterns, as well as the currency composition of foreign borrowing and capital inflows, which impose constraints on portfolio diversification. Additional constraints may be due to the term-structure of international lending contracts, the share of aid and concessional lending in a country's total foreign borrowing, and the share of public or publicly-guaranteed liabilities in total foreign debt. Domestic credit market imperfections, institutional factors and the depth of markets for longer-term hedging may also lead to sub-optimal portfolio management. In this respect, the currency composition of foreign debt portfolios may be determined more directly by supply factors, given the structure of international lending and capital movements, rather than portfolio diversification mechanisms.

Our empirical results are not without policy implications. Policies aimed at capital market development, including adequate prudential regulations in the financial sector, would tend to improve risk management in the private sector, and encourage the development of more sophisticated hedging instruments and private sector-led foreign debt portfolio management. Also, the removal of restrictions to residents' holding of foreign exchange, as well as the liberalization of capital movements in general, could reduce transaction costs in foreign exchange operations and therefore encourage efficient foreign debt management by private sector firms. When governments are the key foreign debt portfolio managers, they should avoid asset and liability structures that are likely to trigger default and subsequently increase portfolio management costs, as well as discourage the development of hedging instruments and markets in the private sector. When the government is an active foreign debt portfolio manager, the implicit assets and liabilities associated with exchange rate and lender of last resort commitments are of particular importance. If such commitments are in place, the government should regulate the behavior of the private sector in order to control the growth of implicit liabilities.

Figure 1: Foreign Currency-Denominated Debt Shares and Exchange Rates, 1970–98
(DM, Japanese Yen, and Swiss Franc)

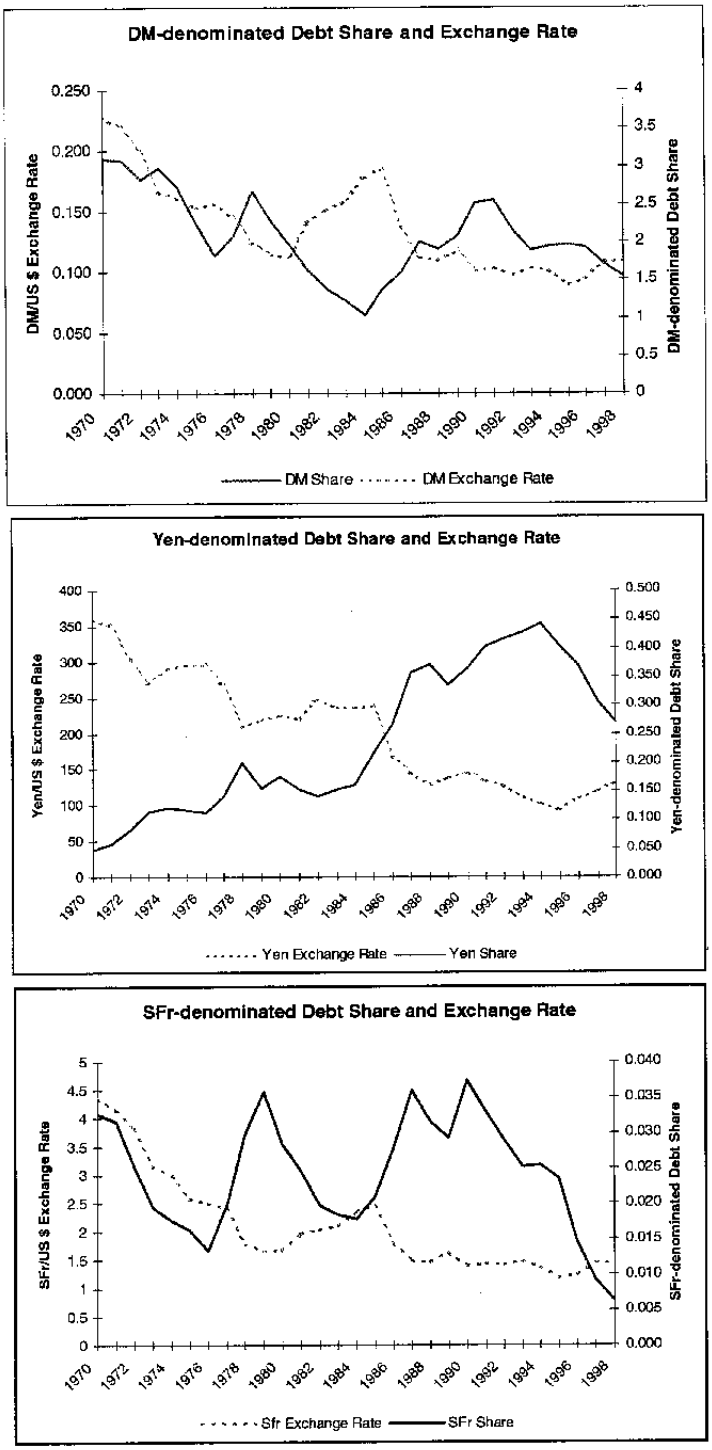

Table 1: Preliminary Tests

	Correlations			Volatility Ratios		
	Ycn	DM	SFr	Yen	DM	SFr
Argentina	-0.39**	-0.55*	-0.67*	15.47*	7.10*	10.34*
	(-2.09)	(-3.15)	(-4.33)			
Brazil	-0.23	-0.21***	-0.12	197.09*	6.41*	13.72*
	(-1.13)	(-1.35)	(-0.58)			
Chile	0.08	-0.40**	0.13	19.47*	3.80*	18.06*
	(0.38)	(-2.09)	(0.63)			
Colombia	0.16	-0.31***	...	26.34*	4.16*	...
	(0.78)	(-1.57)				
Egypt	-0.13***	-0.03	-0.06	6.94*	8.87*	10.29*
	(-1.99)	(-0.45)	(-0.95)			
India	-0.44**	-0.49*	-0.40**	3.39*	2.09***	14.91*
	(-2.34)	(-2.70)	(-2.09)			
Indonesia	-0.35***	-0.22	-0.14	4.10*	3.81*	6.37*
	(-1.79)	(-1.08)	(-0.69)			
Korea (South)	-0.24	-0.17*	-0.13*	32.06*	12.02*	21.77*
	(-4.74)*	(-3.39)	(-3.68)			
Malaysia	-0.28	-0.14	-0.33***	9.93*	25.58*	11.39*
	(-1.40)	(-0.69)	(-1.68)			
Mexico	-0.41**	-0.22	-0.11	13.64*	7.89*	11.21*
	(-2.15)	(-1.08)	(-0.56)			
Philippines	-0.42**	-0.50*	...	9.08*	6.37*	...
	(-2.22)	(-2.76)				
Thailand	-0.22	-0.41**	...	5.97*	4.90*	...
	(-1.08)	(-2.15)				
Turkey	0.08	-0.40**	-0.38**	19.94*	1.68	4.89*
	(0.39)	(-2.09)	(-1.97)			
Venezuela	-0.38**	-0.34**	...	20.79*	23.69*	...
	(-1.97)	(-1.74)				

Sources: IFS data; and IMF staff estimations.

The numbers in parentheses are *t*-statistics. Statistical significance at the 1, 5, and 10 percent levels is denoted by, respectively, (*), (**), and (***). In the case of the bivariate correlations, the null hypothesis is that the correlation between changes in debt shares and exchange rate movements is equal to zero. In the case of volatility, the null hypothesis is that the ratio of the variance of the changes in debt shares to the variance of changes in exchange rates is equal to one.

Table 2: Unit Root Tests in Heterogeneous Panels

	Level		First Difference	
	Without trend	With trend	Without trend	With trend
$Debt_Y$	-1.30	-0.74	-17.21**	-19.32**
$Debt_M$	-2.63*	0.04	-19.15**	-16.55**
$Debt_{SF}$	0.40	0.36	-8.26**	-8.82**
$e_{Yen/US\$}$	2.06	1.89	-3.77**	-3.80**
$e_{DM/US\$}$	-0.21	-3.96**	-3.74**	-3.81**
$e_{SFr/US\$}$	-1.30	0.81	-4.06**	-4.27**

Sources: IFS data; and IMF staff estimations.

Notes: Statistical significance at the 1, 5, and 10 percent levels is denoted by, respectively, (*), (**), and (***). $Debt_Y$ is the debt share denominated in Japanese yen, $Debt_M$ is the debt share denominated in DM, $Debt_{SF}$ is the debt share denominated in Swiss Francs, $e_{Yen/US\$}$ is the Japanese yen/US dollar exchange rate, $e_{DM/US\$}$ is the DM/US dollar exchange rate, and $e_{SFr/US\$}$ is the Swiss franc/US dollar exchange rate.

Table 3: Panel Cointegration Tests

	$Debt_Y$	$Debt_M$	$Debt_{SF}$
Without trend	-2.19**	-1.72***	-0.42
With trend	-3.27*	-4.02*	-4.85*

Sources: IFS data; and IMF staff estimations.

Notes: Statistical significance at the 1, 5, and 10 percent levels is denoted by, respectively, (*), (**), and (***).

Table 4: Panel Cointegration Results
(DOLS Estimator)

	$e_{Yen/US\$}$	$e_{DM/US\$}$	$e_{SFr/US\$}$	R^2	F-statistic
$Debt_Y$	-1.16* (0.23)	0.75 (0.54)	-1.34* (0.41)	0.86	5.59*
$Debt_M$	0.31 (0.21)	-4.56* (0.47)	3.10* (0.33)	0.66	23.03*
$Debt_{SF}$	0.15 (0.31)	-2.71* (0.72)	2.01* (0.52)	0.68	10.17*

Sources: IFS data; and IMF staff estimations.

Notes: Estimations are based on the pooled data for 1970–98 and 14 countries with one lead and two lags of first differenced explanatory variables. All regressions include (unreported) country-specific constants. Standard errors are in parentheses. Statistical significance at the 1, 5, and 10 percent levels is denoted by, respectively, (*), (**), and (***).

Table 5: Panel Cointegration Results
(DFE Estimator)

	$e_{Yen/US\$}$	$e_{DM/US\$}$	$e_{SFr/US\$}$	F-statistic
$Debt_Y$	-0.23*** (0.13)	-0.11 (0.25)	0.22 (0.23)	5.21*
$Debt_M$	0.11 (0.13)	-0.59** (0.28)	0.33 (0.23)	5.78*
$Debt_{SF}$	0.41** (0.22)	-0.06 (0.45)	-0.39 (0.39)	1.32

Sources: IFS data; and IMF staff estimations.

Notes: Estimations are based on the pooled data for 1970–98 and 14 countries with two lags of first differenced explanatory variables. All regressions include (unreported) country-specific constants. Standard errors are in parentheses. Statistical significance at the 1, 5, and 10 percent levels is denoted by, respectively, (*), (**), and (***).

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