2 Sims (1987) has argued that Lucas's critique of econometric policy evaluation – at least in it usual interpretation – is logically flawed. If the parameters of the policy "rule" are subject to change, as they must be if it makes sense to evaluate changes in them, then the public must recognize this fact and have a probability distribution over the parameters of the rule. But then these parameters are themselves policy variables, taking on time series of values drawn from some probability law. Predicting how the economy will behave if the parameters of the rule are set at some value and kept there ("conditional" projections) is logically equivalent to predicting the behavior of the economy conditional on a certain path of a policy variable. Yet this is precisely the kind of exercise that Lucas claims to be meaningless.

3 Transitory shocks are better suited to the analysis of stabilization issues and have the convenient computational feature that the steady-state values of the expectational variables remain unchanged. However, interest rate and exchange rate changes have often a more permanent character, and they are modeled as such here.

4 The experiments were also performed with unanticipated shocks. As a consequence of the assumption of no wage indexation, the major difference with the results reported below is that real output effects are higher, and price effects lower, at the announcement period. Qualitatively, however, the dynamics are basically identical to those discussed in the text for the periods following implementation of policies.

5 In a different context, Lorie (1988) developed a disequilibrium macroeconomic model with fixed prices and credit rationing in which an increase in deposit interest rates (which leads, under profit maximization, to a rise in lending rates) becomes potentially expansionary because it lessens the financing constraint faced by domestic firms.

6 See, for instance, Lizondo and Montiel (1989), and Rojas-Suárez (1987).

7 Note that the parallel market premium therefore falls as a result of both the devaluation of the official exchange rate and the appreciation of the free exchange rate. This is in contrast to what is normally observed in dual exchange market models, in which the reduction of the premium is brought about strictly by changes in the official exchange rate. Our results differ due to the presence of the curb loan market. The informal interest rate is affected by the price level effects of the devaluation.

6

Epilogue

The purpose of this book has been to examine the macroeconomic implications of the coexistence of formal and informal markets – a phenomenon that has been observed in many developing countries. Specifically, we have analyzed the effects of several macroeconomic policy instruments in the presence of informal markets. Along with presenting a fairly comprehensive survey of much of the available literature on informal financial markets and of recent theoretical advances in developing country macroeconomics, we have attempted to show how these markets affect macroeconomic outcomes in developing countries. We have made two important advances in this regard. First, we developed an analytical approach to analyzing how the effects of various shocks are transmitted to the rest of the economy in the presence of informal markets. Second, we developed a detailed macroeconomic model that incorporated a number of features considered "typical" of developing countries including, most importantly, the presence of informal markets.

1 The Analytical Model

The analytical model, which is based on an open-economy, portfolio-balance framework, suggests that informal loan and foreign exchange markets play important roles in transmitting the effects of financial policy instruments to aggregate demand. In addition to interest-rate effects through the informal loan market, the effects of these markets operate through changes in household wealth and through the government budget. Financial repression implies that creditors in the official market are being taxed by the amount that the interest rate in the informal loan market differs from the officially-administered interest rate. Debtors, on the other hand, are being subsidized by this amount. Changes in monetary policy instruments, for example, often result in interest rate changes either in the formal or the informal market, and thus change the rate at which financial repression taxes household portfolios of given composition. They also lead to some switching of assets from one market to another and hence alter the composition of portfolios in ways that affect the base to which the financial repression tax applies. Therefore, such changes affect the effective degree of financial repression and thus the present value of the associated implicit taxes and subsidies and therefore lead to changes in private sector wealth, which in turn affect demand in the economy. In addition, monetary policy changes affect the profits of the central bank and, therefore, have fiscal consequences which in turn also affect economic activity.

Since the parallel market premium, the expected rate of inflation and the economy's stock of financial assets affect the state of domestic demand, the consequences of policy changes for these variables constitute important additional channels through which policy changes may exert aggregate demand effects. The premium in the free market which, as a free asset price, reacts to policy changes, affects the domestic valuation of foreign assets, and hence exerts wealth effects on demand at home. The effects of induced changes in the stock of foreign assets held in the parallel market also matter, but are felt only over time, since this stock evolves as current account surpluses are accumulated.

The analytical model was used to analyze several important monetary policy shocks. Increases in administered interest rates prove to be contractionary because the adverse wealth effects of a lower premium in the free exchange market overwhelm the positive impact emanating from a reduced loan rate. The sale of foreign exchange in the free market by the central bank is shown to be contractionary and similar to open market operations in industrial countries. An increase in the required reserve ratio is contractionary in the short run and in the long run. On the other hand, an increase in central bank credit to the banking system is expansionary in the short run and in the long run; its effect on the steady state premium is also ambiguous.

2 The Simulation Approach

An important aspect of the study has been the development of a more complete macroeconomic model which adds detail and incorporates several additional complications that arise in developing countries, such as nominal wage stickiness embodied in wage contracts, the presence of imported intermediate goods, the possibility of smuggling and unreported earnings, and the incomplete separation of the official and unofficial markets.

The model was developed for use in simulation experiments for the study of the effects of macroeconomic policy and other exogenous shocks in a general-equilibrium, dynamic setting. Such simulations serve to complement the existing theoretical model not only by permitting added complexity, but also by generating full dynamic solutions. The model was calibrated using a set of plausible parameters derived from existing econometric estimates for developing countries, since it was intended to be used for studying likely outcomes in developing countries in general and not in any specific country. The simulation model, therefore, provides a consistent, albeit complex, framework for a quantitative analysis of the macroeconomic implications of the features described above in the context of stabilization and other policies.

We analyzed several important shocks involving policy instruments that have featured prominently in stablization efforts, under the assumption of consistent forward-looking expectations. Specifically, we have considered the dynamic macroeconomic effects of changes in government expenditure on home goods, in central bank credit to commercial banks, in the administered interest rate on bank loans, and in the official exchange rate. The simulation results highlight the role that informal markets play in an economy subject to a variety of government restrictions. They particularly emphasise the fact that, in the presence of flexible financial markets and under rational expectations, shocks begin to exert macroeconomic effects when (credible) policy announcements are made, or when they first become anticipated, rather than when they actually occur. Finally, the results also illustrate how the dynamics of the model are strongly affected by the initial composition of asset holdings in informal markets. In particular, whether the parallel exchange rate depreciates or appreciates, or whether the curb market interest rate rises or falls, depends to a large extent on the size of the initial stock of foreign currency-denominated assets held by private agents in their portfolios, and on whether households are net creditors or net debtors with respect to the commercial banks.

Although only a limited number of policy simulations have been reported here, it should be emphasized that the model is well suited to address a number of substantive research or policy-oriented issues. For instance, a modification in the rationing rule in the official foreign exchange market (for instance, a rise in the proportion of private imports channeled through the official market) has substantial effects on prices and the parallel exchange rate, through stock-flow interactions built into the model. Similarly, a rise in the export tax rate is associated with a fall in export tax revenue and officially recorded inflows of foreign exchange. A higher tax rate on exports reduces the relative price of exports perceived by domestic producers, and this in turn tends to reduce supply and to increase the incentive to divert exports from the official market to the parallel market. Finally, an increase in world interest rates may prove to be expansionary, despite its impact on external debt payments. When the interest rate on foreign assets rises, private agents attempt to reallocate their portfolios, moving away from domestic assets and towards foreign assets. Since the central bank does not accomodate this desired portfolio shift, the parallel exchange rate depreciates. If the private sector is a net external creditor, the positive wealth effect associated with the exchange rate change stimulates private expenditure, as well as output of home goods.

The key implications of the simulation experiments can be summarized as follows. Anticipated expansionary fiscal and credit policies are associated with a rise in output and prices, a fall in the informal interest rate, as well as a current account deficit and a depreciation of the parallel exchange rate. An increase in lending and deposit interest rates in the formal financial sector may lead to a fall, rather than an increase, in private bank deposits, and may well be associated with a contraction in output. An anticipated once-and-for-all devaluation of the official exchange rate is associated with a fall in the parallel market premium, both in the short- and the long-run. The devaluation also has a temporary expansionary effect in the transition period following the announcement, and a contractionary effect when implemented, which is slow to dissipate. This result highlights the importance of allowing for a proper time frame in judging whether or not devaluations are contractionary. In all of those cases, informal markets for credit and foreign exchange play key roles in transmitting the influences of shocks, and the economy's dynamic response is influenced to an important degree by "leakages" into parallel markets. Our results, therefore, demonstrate that short-run macroeconomics in developing countries can ill-afford to ignore the influence of such "unofficial" transactions.

3 Reform

The analytical and simulation approaches that we have developed in this book help clarify the dyamics of the effects of various shocks in developing economies where informal markets are created as a response to financial repression and capital controls. However, in the face of growing evidence against the efficacy of non-market oriented policies as well as increasing financing constraints at home and abroad, increasingly countries have had to loosen control and such markets. In several instances, the adoption of policies that would allow completely liberalized domestic financial markets, allowing market-determined domestic interest rates and removing all exchange controls has been announced as a goal and in some cases even implemented. The complete freeing of borrowing and lending rates, however, remains rare in the developing world and, where adopted, has not always been carried out successfully or retained permanently.

The most recent of these announcements of reform have emanated from the former centrally-planned countries of Eastern Europe and the new republics that have been born out of the former Soviet Union. As these countries move in piecemeal fashion towards their avowed goal of complete liberalization, policy will have to attempt to achieve its targets in the presence of both formal and informal markets. The models and results presented here are, therefore, likely to be important for understanding the transmission effects in those economies as well.

While reform and the transition to reform is specifically beyond the scope of this book, the events in the former socialist economies imply that they are quite likely to be the subject of considerable research in the coming years. Models such as the ones presented here will be quite useful in analyzing issues such as liberalization of domestic financial markets, unification of exchange rates, and abolition of capital controls. Where data is available, parameters of the model could be estimated to allow simulations to be performed for alternative reform or policy scenarios in the case of a particular country. In such cases more precise estimates of the short- and long-run effects of envisaged reforms could be derived. Our simulations do, however, point to the likely consequences of a reform that serves to gradually eliminate distortions such as financial repression and dual exchange rates. The importance of initiating appropriate domestic demand management prior to undertaking any such reform is stressed by the simulations, which show that anticipated expansionary fiscal and credit policies are inflationary and lead to current account deficits as well as a depreciation of the parallel exchange rate. By itself, the reform is likely to have short run contractionary effects since, in the simulation experiments, both a devaluation of the official rate as well as an increase in leading and deposit interest rates in the formal financial sector lead to a decline in output and impact. Consequently, any reform effort must be accompanied by either other policy measures to mitigate the effects of such a contraction, or financing to meet the cost of adjusting to the post-reform environment, or some combination of the two.

Appendix A

Efficiency of Parallel Currency Markets in Developing Countries

This appendix examines the short- and long-term relationships between the official and parallel exchange rates for a group of developing countries, using causality and co-integration tests. Section 1 presents the results of bivariate causality tests. Section 2 presents and discusses the results of the co-integration analysis.

1 Causality Tests

To analyze the short-term relationship between official and parallel market exchange rates, two-way Granger causality tests are performed.[1] There is, a priori, no reason to exclude one of the causal directions in this case. On the one hand, the official exchange rate may Granger-cause the parallel rate if private agents have imperfect information or observe key macroeconomic variables with a lag. For instance, suppose that the monetary authorities are known to possess some "inside" information on the state of the economy (such as the level of net foreign assets, capital inflows, etc.) and to incorporate this information into the setting of the official exchange rate. In such a context, private agents will attach some weight to the behavior of the official exchange rate in forming expectations about the parallel exchange rate, since the former is assumed to embody "new" and relevant information.[2] On the other hand, parallel exchange rates may also Granger-cause official exchange rates in economies where exchange rate policy is "endogenous". In countries which operate a crawling peg regime for instance, the central bank often follows a real exchange rate rule whereby the decision to depreciate the official exchange rate is made dependent on the level of the (expected) domestic inflation rate. As discussed in chapter 1, the parallel rate may have a strong effect on domestic prices because it measures the marginal cost of foreign exchange. The authorities may therefore use (recent) changes in the parallel market rate as a predictor of inflationary trends in their reaction function for setting the official exchange rate. Causality from the parallel to the official rate might also occur if the authorities follow a "premium-based" devaluation rule, by which the official rate is depreciated in order to reduce the exchange-rate spread. A typical example in this context would be Bolivia in the early 1980s (Kharas and Pinto, 1989).

Since the Granger causality test is designed to analyze mainly bivariate weakly stationary stochastic processes, let $\tilde{e}$ ($\tilde{b}$) denote a transformed stationary value of the official (parallel) exchange rate, e (b).[3] $\tilde{e}$ is said to Granger-cause $\tilde{b}$ if lagged values of $\tilde{e}$ reduce significantly the forecasting errors obtained from an autoregression of variable $\tilde{b}$ on its lagged values:

$$\tilde{e}_t = \sum_{k=1}^{m_1} \alpha_{1k}\tilde{e}_{t-k} + \sum_{k=1}^{m_2} \alpha_{2k}\tilde{b}_{t-k} + \varepsilon_{1t} \qquad (A1a)$$

$$\tilde{b}_t = \sum_{k=1}^{n_1} \beta_{1k}\tilde{e}_{t-k} + \sum_{k=1}^{n_2} \beta_{2k}\tilde{b}_{t-k} + \varepsilon_{2t} \qquad (A1b)$$

If $\tilde{e}$ is causal to $\tilde{b}$, all coefficients α_{2k} are equal to zero and there is at least one coefficient $\beta_{1k} \neq 0$.[4] Since the original series are likely to be nonstationary, they have to be transformed into stationary ones by means of a difference filter or some nonlinear transformation, like a Box-Cox procedure. In this study, the exchange rate series are transformed by application of the widely used log-difference filter, so that $\tilde{z}_t \equiv \Delta \log z_t$, where $z = e, b$. Unfortunately, there is no proof that such a transformation does not affect the causality structure (see Geweke, 1984). As a consequence, the differenced time series do not have any information about the long-run relation between the trend components of the original series. Therefore the standard Granger causality tests applied here are taken to describe only *short-run* relations between the official and parallel exchange rates. This is due to the fact that the differenced variables, $\Delta \log z_{t-k}$, are equal to zero in equilibrium, imposing long-run neutrality between the variables from the beginning. Co-integration tests are used in the next paragraph to analyze the long-run relationship between the original exchange rate series.

Results of causality tests are typically quite sensitive with respect to the choice of lag length. Accordingly, instead of imposing an arbitrary (and identical) lag length for all variables, the lag length parameters $m_1, m_2, \ldots$ are chosen optimally by using the final prediction error (*FPE*) procedure due to Hsiao (1979). The *FPE* criterion is defined as:

$$FPE = \frac{T+k}{T-k}\frac{1}{T}SSR, \qquad (A2)$$

where T denotes the number of observations, k the number of regressors, and *SSR* the sum of squared residuals obtained from equations $(A1)$. The final prediction error can be seen as an unbiased estimator of the residual variance multiplied by the penalty factor $(T+k)/T$ for additional parameters.

In testing if $\tilde{e}$ is non-causal to $\tilde{b}$ for instance, the test procedure is as follows. First, regress $\tilde{b}$ on lagged $\tilde{b}$ with a "large" univariate maximum lag n_2. Second, reduce the lag length as long as the *FPE* is reduced. This gives the optimal univariate lag length n_2^*. Third, try to reduce the final prediction error at this minimum lag length by adding lagged $\tilde{e}$ variables. If there is no reduction in the *FPE* by including $\tilde{e}$, $\tilde{e}$ is non-causal to $\tilde{b}$. That means H_0 is rejected.[5]

The maximum lag length of both variables is initially chosen to be 8. Table A.1 presents the empirical results obtained by applying the Hsiao procedure described above to 21 developing countries: Argentina, Bangladesh, Bolivia, Brazil, Colombia, Ecuador, Greece, India, Indonesia, Korea, Malaysia, Malawi, Mexico, Morocco, Nigeria, Pakistan, the Philippines, Saudi Arabia, Singapore, Tunisia, and Zambia.[6] Both the official and the parallel exchange rates are end-of-period domestic currency/US dollar rate and are taken, respectively, from the *International Financial Statistics*, and the *World Currency Yearbook* (formerly *Pick's Currency Yearbook*). Data are quarterly and cover the period 1972:1–1989:4. The table presents the F-test on block significance of the lagged independent variables.

Table A.1 Bivariate causality tests between official and parallel exchange rates, 1972:1–1989:4

Country	*Official → Parallel*[a]	*Parallel → Official*[b]
Argentina	0.848	1.243
Bangladesh	0.317	1.102
Bolivia	4.463**	17.864**
Brazil	5.116**	1.587
Colombia	2.004	2.793**
Ecuador	2.116*	3.134**
Greece	2.536*	2.973*
India	0.346	2.297*
Indonesia	0.589	1.106
Korea	1.743	0.576
Malaysia	1.819	0.489
Malawi	0.497	0.605
Mexico	2.159*	1.963
Morocco	2.708*	0.572
Nigeria	1.407	0.687
Pakistan	1.609	3.464**
Philippines	3.107*	2.884*
Saudi Arabia	4.158**	1.506
Singapore	1.474	1.164
Tunisia	1.540	2.216*
Zambia	5.673**	5.048**

[a] The null hypothesis is that the official exchange rate does not cause the parallel rate.
[b] The null hypothesis is that the parallel exchange rate does not cause the official rate.
→ Direction of causation test.
* Significant at 0.05 level.
** significant at 0.01 level.

The results indicate that for four countries in the sample (Brazil, Mexico, Morocco, and Saudi Arabia) causality runs only from the official to the parallel exchange rate.[7] For four countries (Colombia, India, Pakistan and Tunisia), the parallel rate seems to lead the official exchange rate while for five countries (Bolivia, Ecuador, Greece, Philippines and Zambia), causation seems to run both ways. In all other cases, there is not much evidence of a causal relationship between exchange rates.

It is important, however, to keep in mind that the causality

tests reported here may be dependent on the data frequency used for performing the tests. The contemporaneous relationship found in a few cases may result, for instance, from the fact that in some countries the speed of adjustment to new information in foreign exchange markets is relatively fast. Using lower frequency data - monthly data or even daily data as done by Akgiray *et al.* (1990) for instance in their study of the parallel market in Turkey - could generate different results.

2 Co-integration Tests

As shown in chapter 2, an important class of analytical models dealing with the macroeconomic implications of parallel currency markets in developing countries is the portfolio-currency substitution approach. One of the key implications of this type of model is the prediction that, in the long run, the parallel market premium is constant, implying that the parallel exchange rate depreciates in the same proportion as the official exchange rate.

The long-run proportional relationship between the rate of depreciation of the official and parallel exchange rates suggested by stock-flow models of parallel currency markets can formally be expressed in regression form as:

$$\log b_t = \alpha \log e_t + u_t, \qquad (A3)$$

where u_t denotes an error term. Accordingly, theory predicts that $\alpha \cong 1$, implying a stable long-run relationship between (the logarithms of) the official and parallel exchange rates.

A testable definition of long-run equilibrium existing among nonstationary time series is provided in the recently developed theory of co-integration.[8] Simply put, if two time series, z_1 and z_2, are nonstationary (as is the case with many economic magnitudes which typically trend through time), but some linear combination of them is a stationary process, then z_1 and z_2 are said to be co-integrated.

Statistically, a time series is said to be (weakly) stationary if its mean, variance, and covariances are all invariant with respect to time. Such a series is denoted $I(0)$, meaning "integrated of order zero". A time series requiring first-order differencing to achieve stationarity is said to be $I(1)$. Generally, any linear combination of two $I(1)$ time series will also be an $I(1)$ series. However, if there exists some linear combination of the two series which itself is $I(0)$, then co-integration exists. Hence, if z_1 and z_2 are each $I(1)$ except for some constant α, the sequence, x_t, defined as $x_t = z_{1t} - \alpha z_{2t}$, is $I(0)$, and z_1 and z_2 are co-integrated. The co-integrating parameter, α, is unique when it exists. More generally, a variable z is said to be integrated of order d, denoted $z_t \sim I(d)$, if the dth difference of z is stationary. Let Z_t be an $n \times 1$ vector of time series. The vector Z_t is said to be co-integrated of order d, p, denoted $Z_t \sim CI(d,p)$ if (i) each component of Z_t is integrated of order d; and (ii) there exists a non-zero vector α such that the linear combination $\alpha' Z_t$ is integrated of order $d-p$, for $p>0$.

The concept of co-integration provides a useful statistical definition of long-run equilibrium existing between two (or more) non-stationary time series. When a linear combination of z_1 and z_2 is a stationary stochastic process, the variables may be said to be in a state of "statistical equilibrium". If this condition is not met, then the two series will not display any systematic tendency to evolve in time in proximity of each other. Here, as shown by equation ($A3$), the long-run equilibrium predicted by theory between changes in official and parallel exchange rates suggests that these variables should be co-integrated with a co-integrating parameter, α, close to one. If this condition is met, any short-run divergences will tend to be eliminated by equilibrating economic forces.

The tests for co-integration reported here follow the two-step procedure outlined in Engle and Granger (1987). Following the definition of co-integration, the order of integration of each (non-stationary) time series is determined using a standard augmented Dickey-Fuller procedure (see Dickey and Fuller, 1981).[9] To test the null hypothesis that each element z_{kt} of Z_t is $I(1)$, the following equation is estimated using ordinary least squares:

$$\Delta z_{kt} = c_0 + c_1 z_{kt-1} + \sum_{j=1}^{q} c_{j+1} \Delta z_{kt-j} + u_t \qquad (A4)$$

The null hypothesis is H_0: z_{kt} is $I(1)$, which is rejected in favor of $I(0)$ if $\hat{c}_1$ is found to be negative and statistically significant. The conventionally calculated t-statistic for the estimated coefficient $\hat{e}_1$ can be used, although it does not follow the t-distribution

under H_0. Critical values are taken from the tables produced by MacKinnon (1990). The parameter q is chosen to render the disturbance term u_t approximately white, namely, to obtain serially uncorrelated residuals. For $q=0$, the test procedure is referred to as the Dickey-Fuller (DF) test, while for $q>1$ it is referred to as the Augmented Dickey-Fuller (ADF) procedure.

Table A.2 presents ordinary least squares estimates of ($A4$), with

Table A.2 Unit root tests of quarterly log of the official and parallel exchange rates, 1972:1–1989:4[a]

Country	*q*	*DF/ADF test statistic*
Argentina		
Official	3	−1.293
Parallel	2	−0.744
Bangladesh		
Official	0	−2.071
Parallel	1	−2.375
Bolivia		
Official	3	−1.953
Parallel	2	−2.121
Brazil		
Official	2	1.819
Parallel	1	2.156
Colombia		
Official	2	−0.580
Parallel	4	−0.552
Ecuador		
Official	1	0.994
Parallel	4	−0.854
Greece		
Official	1	−1.840
Parallel	4	−2.614
India		
Official	2	−0.399
Parallel	2	−2.687
Indonesia		
Official	0	−2.571
Parallel	0	−2.557
Korea		
Official	0	0.051

Table A.2 *cont.*

Country	*q*	*DF/ADF test statistic*
Parallel	1	−1.034
Malaysia		
Official	0	−2.509
Parallel	3	−2.247
Malawi		
Official	4	−0.827
Parallel	0	−3.852
Mexico		
Official	3	−1.991
Parallel	2	−1.945
Morocco		
Official	4	−2.059
Parallel	4	−2.407
Nigeria		
Official	1	−0.469
Parallel	0	−1.828
Pakistan		
Official	0	−4.891
Parallel	1	−3.070
Philippines		
Official	2	−2.124
Parallel	0	−1.875
Saudi Arabia		
Official	4	−3.192
Parallel	4	−2.345
Singapore		
Official	0	−3.375
Parallel	2	−1.963
Tunisia		
Official	2	−2.129
Parallel	2	−2.279
Zambia		
Official	4	−1.862
Parallel	4	−3.995

[a] Sample size = 72. The approximate 1, 5, and 10 percent critical values for these tests are −4.09, −3.47, and −3.16, respectively (MacKinnon, 1990).

a trend term added, for the 21 countries considered above. In each case, the "optimal" value of q (which has been selected in the range 0 to 4) is also reported. The results indicate that, apart from the parallel rate in Malawi and the official rate in Saudi Arabia, and Singapore, the null hypothesis that the log-levels of the official and parallel exchange rates have a unit root cannot be rejected, implying therefore nonstationarity in levels.[10]

As stated above, a necessary condition for the exchange rate series to be co-integrated is that both series must be integrated of the same order. Since for all countries except Brazil, Malawi, Saudi Arabia and Singapore the *DF/ADF* test establishes that both the official and parallel exchange rates can reasonably be taken to be $I(1)$, the co-integrating parameter α can be estimated by running the "co-integrating regression":

$$\log b_t = \alpha_0 + \alpha \log e_t + v_t, \qquad (A5)$$

and testing whether the residual series $\hat{v}_t$ is stationary. Stock (1987) has shown that $\hat{\alpha}$ will be "superconsistent" in that it will have a finite sample bias of the order of $1/T$, T being the sample size.[11] If the residual series rejects the null of $I(1)$, then the series used in ($A5$) are co-integrated. In what follows, and as suggested by Engle and Granger (1987), the *DF/ADF* procedure (equation $A4$) is used here to conduct the co-integration test.

Table $A.3$ gives the *DF/ADF* test statistic for all the countries in the sample, together with the selected value of q, except for those countries for which the assumption of a unit root in log levels was rejected by the *DF/ADF* test for either one of the series. The null hypothesis of no co-integration is rejected for nine countries: India, Indonesia, Malaysia, Mexico, Morocco, the Philippines, Tunisia and – to a lesser extent – Brazil and Greece. Moreover, the co-integrating parameter is relatively close to unity in all countries pertaining to this group, except India and Tunisia, indicating therefore a long-run proportional relationship between the official and parallel exchange rates. For all other countries, the *DF/ADF* test statistic does not reject the null of a unit root. The results suggest therefore mixed evidence in favor of a co-integrating relationship among official and parallel exchange rates in developing countries.

Finally, it should be noted that there is a crucial relationship

Table A.3 Estimation results for the co-integrating regression, 1972:1–1989:4[a]

Country	*q*	*DF/ADF test statistic*	$\hat{\alpha}$[b]
Argentina	1	−2.514	0.927
Brazil	4	−3.431	0.893
Bangladesh	0	−2.089	0.013
Bolivia	4	−1.299	1.052
Colombia	4	−2.887	0.827
Ecuador	3	−2.660	1.007
Greece	4	−3.379	0.945
India	0	−4.769	0.685
Indonesia	0	−5.914	0.821
Korea	4	−2.752	0.806
Malaysia	4	−3.787	0.927
Malawi	–	–	–
Mexico	0	−4.152	0.991
Morocco	0	−5.298	1.035
Nigeria	2	−2.299	0.967
Pakistan	2	−1.892	0.731
Philippines	1	−4.001	0.927
Saudi Arabia	–	–	–
Singapore	4	−2.511	0.901
Tunisia	2	−3.953	0.731
Zambia	–	–	–

[a] Sample size = 72. The null hypothesis is that the official and parallel exchange rates are not co-integrated, or that the residual series in equation (*A5*) is *I*(1). The approximate critical values for the *DF/ADF* statistic at the 1, 5 and 10 percent level are, respectively, −4.56, −3.92, and −3.60 (MacKinnon, 1990).
[b] $\hat{\alpha}$ is the co-integrating parameter given in equation (*A5*).

between co-integration, error-correction models, Granger causality and tests of market efficiency. If $\alpha' Z_t \sim CI(1,1)$, then by the Granger Representation Theorem (Engle and Granger, 1987), here exists an error-correction representation of the form:

$$A(L)\Delta Z_t = -\gamma\alpha' Z_{t-1} + \beta(L)\varepsilon_t, \qquad (A6)$$

where $A(L)$ is a matrix polynomial in the lag operator L with $A(0) = I_n$, γ is a $n \times 1$ vector of constants, $\beta(L)$ is a scalar polynomial in L, and ε_t is a white noise disturbance term.

From (*A6*) it is obvious that since at least one lagged value of the vector Z_t enters this system with a non-zero coefficient, then it immediately follows that the knowledge of Z_{t-1} can be used to help forecast the current level of Z_t. In other words, at least one exchange rate Granger-causes the other. However, in a (weakly) efficient exchange market, current prices are assumed to reflect all available information, in such a way that the best predictor of a spot rate is its own lagged value. As a result, the existence of a co-integration relationship is inconsistent with this definition of market efficiency.[12] Since the results given above provide mixed evidence regarding co-integration between official and parallel exchange rates, no general conclusion emerges regarding the efficiency of unofficial foreign currency markets in developing countries.

Notes

1 In simple terms, a time series x is said to Granger-cause a series y if the prediction of y is improved by including lagged values of x in addition to lagged values of y in a regression where y is treated as the dependent variable. For a more detailed discussion, see Harvey (1990).

2 As a result, evidence of a causal relationship from the official to the parallel exchange rate is often viewed as an indication of market efficiency (see, for instance, Gupta, 1981).

3 The stationary assumption is needed because the Granger test of causality does not assume strict exogeneity of lagged $\tilde{e}$ in the regression for $\tilde{b}$ and vice versa.

4 Alternatively, this hypothesis can be formulated in terms of Granger non-causality. $\tilde{b}$ is *not* Granger-causal to $\tilde{e}$ if H_0: $\alpha_{2k} = 0$ (for all k) and $\tilde{e}$ is not Granger-causal to $\tilde{b}$ if H_0: $\beta_{1k} = 0$ (for all k).

5 Other criterion functions commonly used for selecting the optimal lag length in Granger causality tests are the Akaike's information criterion (*IC*) and the Hannan-Quinn (*HQ*) criterion. Teräsvirta and Mellin (1986) have shown that all three criteria can be seen as classical *F*-tests with different critical values depending on the number of observations as well as on the number of parameters.

6 Although this group of countries was chosen primarily on the basis of data availability, it includes both low- and middle-income developing economies, manufacturing- and primary-exporters, as well as service

and remittance countries, and heavily-indebted countries. This diversity makes the sample reasonably representative of developing countries in general.

7 The fact that the F statistic is not significant for Korea may be because we consider only the parallel market rate relative to the US dollar. This may not be appropriate, since during the 1970s and early 1980s the parallel market in the Japanese yen has developed substantially. A similar problem was noted Akgiray *et al.* (1990) with respect to the West German mark market in Turkey.

8 A comprehensive treatment of the statistical theory of co-integration is giving in Campbell and Perron (1991).

9 One important shortcoming of Dickey-Fuller procedures is the assumption of independent and identically distributed errors. Other methods which allow for weakly dependent and heterogeneously distributed error processes have recently been developed; see Pagan and Wickens (1989).

10 For Brazil, the *DF/ADF* test provides statistics that have the wrong sign for the series to be level stationary. For Pakistan, The assumption that the parallel exchange is not $I(1)$ is only marginally rejected at the 5 percent significance level.

11 For $I(0)$ variates, an estimator is considered consistent if it has finite sample bias of the order or $\sqrt{T}$.

12 Koveos and Seifert (1986) examine the issue of market efficiency within the framework of the purchasing power parity theory. Culbertson (1989) presents econometric evidence (tests of random walk behavior), using monthly data for ten developing countries, in favor of the assumption that parallel currency markets are weakly efficient. See also Gupta (1981) and Booth and Mustafa (1991).

Appendix B

Informal Credit Markets and Capital Mobility in Developing Countries

The data presented in chapter 1 suggest that interest rates in informal credit markets tend to be relatively high by the standards of industrial countries as well as by those of interest rates in the formal financial sectors of developing countries. While the microeconomic literature considers whether such rates may arise from monopoly power or from higher costs of operating in these markets, in fact there may be good macroeconomic reasons for market-determined nominal interest rates in developing countries to be high by the standards mentioned above. Not only can *real* interest rates be expected to be high in the capital-scarce context of these countries, but also ongoing inflation in many such countries would imply that nominal rates must persist at high levels to preserve positive real interest rates. On the other hand, domestic market-determined rates may be determined by arbitrage relationships with rates of return on holding foreign financial assets. These may also be high, in the presence of political and exchange rate risk.

In this appendix we describe the results of a series of tests conducted by Haque and Montiel (1990, 1991) that aim at determining whether market-determined interest rates in developing countries' informal loan markets respond primarily to domestic phenomena or to external arbitrage conditions. In effect, this is a test of capital mobility, since external factors would dominate only if domestic and foreign assets are close substitutes in households' portfolios and if portfolio equilibrium holds continuously. The difficulty in performing such tests is that the observable interest rate series in developing countries pertain to controlled rates in the formal sector, not to the relevant market-determined informal rates. In the absence of such information, arbitrage conditions cannot be tested directly. Haque and Montiel (*op. cit.*) therefore developed an alternative indirect test, which is described below.

1 Macroeconomic Models and Capital Mobility

The extent to which domestic market-determined interest rates respond primarily to external arbitrage opportunities or to domestic factors has an important bearing on the short-run effects of stabilization policies - including monetary, fiscal, and exchange rate policies - in developing countries. It affects, for example, the extent to which expansionary fiscal policy crowds out private investment. It also determines the degree to which a nominal devaluation may be contractionary as well as the ability of monetary policy to affect aggregate demand. Perhaps because of the absence of time-series data on interest rates in the informal sector, despite a considerable interest in stabilization policies in many developing countries no clear consensus has emerged on the degree of capital mobility that can be taken as characteristic of such economies in general or of any groups of developing countries in particular. Policy initiatives have often been undertaken without any clear notion of the extent to which domestic market-determined interest rates were likely to be affected over any relevant period of time. Frequently, in empirical and theoretical models, developing countries are modeled as having closed capital accounts, except for public-sector external borrowing and amortization. Perhaps such an assumption is often maintained because, as indicated in chapter 1, the vast majority of developing countries have maintained significant legal restrictions over capital movements - both inflows and outflows - for balance of payments reasons as well as to facilitate monetary control.[1]

Despite the widespread maintenance of restrictions on the books among developing countries, however, there are a number of factors that suggest that capital controls have not been completely effective. Massive episodes of capital flight have been observed in many countries in spite of restrictions on capital movements.[2] Illegal means (over and under-invoicing, smuggling, bribery, etc.) have frequently been used to circumvent capital controls. As a result, the stocks of flight capital held by citizens of developing countries have at times been a significant fraction of the total stock of their countries' gross external debt. The accumulation of holdings of foreign capital by domestic residents (e.g., as a result of past capital flight or the accumulated earnings of migrants) reinforces the ability of residents to evade controls. When domestic residents own substantial assets abroad, remittance decisions may effectively perform the financial arbitrage function in a manner that cannot easily be controlled by the domestic authorities.

2 Empirical Evidence on Capital Mobility

Empirical research on capital mobility in developing countries has been meager. Until recently, few attempts had been made to estimate the degree of capital mobility for large groups of developing countries. The early empirical tests of the monetary approach to the balance of payments estimated reserve-flow equations to ascertain the magnitude of the "offset coefficient" applicable to changes in domestic credit. These studies typically found that attempts by domestic monetary authorities to control the money supply were thwarted by the capital outflows that were triggered in response. This limited scope for monetary autonomy suggested that a high degree of capital mobility prevailed in developing countries during the years that the tests were conducted.[3] More recent evidence provided by Cumby and Obstfeld (1983) for the Mexican experience as well as by Takagi (1986) for a number of Central American countries, is also consistent with the view that there is little if any scope for monetary autonomy in developing countries. Because of the proximity of these countries to North America, however, capital may be more mobile than other developing countries. To that extent, these countries may constitute special cases. By contrast, Phylaktis (1988), finds that capital controls are able to account for a substantial portion of the internal-external nominal interest rate differential in Argentina. Edwards and Khan (1985) on the other hand, find that capital mobility has been very high (effectively infinite, in fact) for Singapore, but much less so in Colombia. Recently, some studies have treated the estimation of capital mobility in a more systematic manner and across a broader group of developing countries. Haque, Lahiri and Montiel (1990) embedded the degree of capital mobility in a macro-econometric model for a large number of developing countries. The results showed that on average a high degree of capital mobility prevailed in developing countries in general.

3 Estimation of the Degree of Capital Mobility

The empirical tests of interest arbitrage to be described below are based on estimation of a single parameter that places individual countries on a continuum between completely open financially and completely closed. Consider first, the case of a completely open economy. In equilibrium, the market-determined interest rate in the domestic economy has to be equal to the interest rate that prevails abroad plus the expected rate of depreciation of the exchange rate – i.e, uncovered interest parity would obtain as follows:

$$i^* = i^w + \mathrm{E}(\hat{e}), \qquad (B1)$$

where i^* is the market-determined interest rate that would prevail in the domestic informal market if the capital account were completely open, i^w is the world interest rate, E is the expectation operator, and $\hat{e}$ is the rate of change of the exchange rate. If uncovered interest parity obtains, people would be indifferent between holding domestic interest-bearing assets such as loans in the informal credit market and foreign interest-bearing assets.

Next, consider the polar case of a completely closed economy where capital controls are perfect. Since capital flows are now zero by assumption, the domestic market-determined interest rate has to adjust to clear the money market. In that case, the supply of domestic currency would be willingly held. Finally, consider the intermediate case where capital controls are less than perfect, but not completely ineffective. Domestic market-determined interest rates would now be influenced by both interest arbitrage considerations and by domestic money market conditions. In this way, the interest rate in the informal financial sector can be viewed as a weighted average of the two polar extremes, corresponding to completely closed and completely open economies. Thus the domestic market-clearing interest rate (i) can be expressed as a weighted average of the uncovered interest parity interest rate (i^*) and the domestic market-clearing interest rate that would be observed if the private capital account were completely closed (i').[4] Hence:

$$i = \varphi i^* + (1 - \varphi)i'; \qquad 0 \leqslant \varphi \leqslant 1. \qquad (B2)$$

The parameter φ is taken to be a structural feature of the economy, and as such amenable to empirical estimation. This parameter serves as an index of capital mobility, or of effective substitutability between domestic and foreign interest-bearing assets. If $\varphi = 1$, the domestic market-clearing interest rate is equal to its uncovered parity value. In this case, external financial influences overwhelm domestic monetary factors in the determination of domestic market-clearing interest rates – i.e., capital mobility is perfect. On the other hand, if $\varphi = 0$, external financial factors play no role in the determination of the domestic interest rate, a situation which could arise only if the private capital account were effectively closed. Intermediate values of φ arise when foreign and domestic financial factors interact in the determination of domestic interest rates, and correspond to intermediate degrees of capital mobility. As φ moves from zero to unity, of course, the effective degree of capital mobility increases.[5]

Expressing domestic interest rate determination as in equation ($B2$) has the advantage of treating the degree of interest arbitrage as an estimable parameter. The expression can be motivated in a number of different ways. Perhaps the simplest is to regard it as a product of "artificial nesting" in which the competing hypotheses of completely perfect and zero capital mobility are combined into a single equation and the data are allowed to discriminate between them. Alternatively, an equation like ($B2$) could represent the reduced form of a two-equation system which simultaneously explains the domestic interest rate and the magnitude of capital flows as a function of foreign interest rates, the arguments in domestic asset-demand functions, and components of the domestic money supply other than capital flows (i.e., lagged reserves, domestic credit, and the current account).[6] On this interpretation, the interest rate i' would capture the effects of factors which affect the domestic money market other than foreign interest rates and capital flows.

Equation (*B*2) could also emerge as the outcome of the behavior of the authorities that administer the system of capital controls. A slight rearranging of (*B*2) yields:

$$i - i^* = (1 - \varphi)(i' - i^*), \tag{B3}$$

that is, deviations from uncovered parity are proportional to the divergence between the interest rate that would correspond to a closed private capital account and the external (uncovered parity) rate. This may arise in a situation in which capital controls are partially relaxed *de jure* when financial pressures (either for outflows or inflows) are strong, or when such circumstances make controls difficult to enforce, so that the effectiveness of controls varies with the severity of such pressures de facto.

If the three interest rates – the open economy rate (i^*), the closed economy rate (i'), and the rate that actually prevails (i) – were observable, then the weighted relationship (*B*2), could be estimated directly. The estimate of the weighting factor φ, would then give us an estimate of the degree of interest arbitrage. However, as already indicated, the relevant market-determined interest rate in many developing countries is the informal market rate, for which time series data are rarely available, and the interest rate that would prevail if capital markets (i') were closed is only a hypothetical construct. In most countries we would expect to observe only the administered interest rate that prevails in the formal sector and the *ex post* uncovered interest parity rate. Thus an alternative method has to be used for obtaining an estimate of φ without the unobservable rate i and the unobservable rate i'.

As indicated earlier, the interest rate i' is that which would prevail in the domestic economy in the absence of private capital flows. To derive an expression for i', express the standard money supply identity:

$$M \equiv R + D, \tag{B4}$$

where M is the domestic money supply, R is the domestic-currency value of foreign exchange reserves, and D is the stock of domestic credit outstanding. This can be rewritten using the fact that current

reserves are equal to reserves at the end of the last period plus the changes in reserves in the current period as follows:

$$M \equiv R_{-1} + D + \Delta R, \tag{B5}$$

where Δ is the first difference operator. Using the balance of payments identity to substitute for the change in reserves ΔR, this can be written:

$$M \equiv R_{-1} + D + CA + KAG + KAP, \tag{B6}$$

where CA, KAG, and KAP are the domestic-currency values of the current account, public capital account, and private capital account respectively. If the capital account is closed, then private capital accounts (KAP) are by definition zero. Consequently, the money supply that would correspond to a situation with a closed private capital account, which we shall denote M', is given by:

$$M' \equiv R_{-1} + D + CA + KAG. \tag{B7}$$

Comparing (*B*6) and (*B*7), we can see that:

$$M' \equiv M - KAP, \tag{B8}$$

i.e., M' is equivalent to the actual money supply less the portion of reserve flows accounted for by private capital movements.

We specify a demand for money function quite conventionally as follows:

$$\log\left(\frac{M^d}{P}\right) = \alpha_0 + \alpha_1 i + \alpha_2 \log y + \alpha_3 \log\left[\frac{M}{P}(-1)\right], \tag{B9}$$

with real output denoted by y, the domestic price level given by P, and M^d denoting the demand for money. This money demand function is semi-logarithmic in interest rates and assumes standard stock adjustment behavior. Given the usual assumptions about the demand for money, the coefficients can be expected to be of the following signs:

$$\alpha_1 < 0, \quad \alpha_2, \quad \alpha_3 > 0$$

Given this money demand specification, recall that the interest rate i' is that value of the domestic interest rate that equilibrates the money market when the economy is closed. Consequently, with money demand given by equation ($B9$) and the money supply of a closed economy given by equation ($B7$), i' can be derived from the following equilibrium condition:

$$\log\left(\frac{M'}{P}\right) = \log\left(\frac{M^{d}}{P}\right). \qquad (B10)$$

By using ($B9$) in ($B10$) and solving for i' we can derive an expression for i':[7]

$$i' = -\frac{\alpha_1}{\alpha_0} + -\frac{1}{\alpha_1}\log\left(\frac{M'}{P}\right) - \frac{\alpha_1}{\alpha_2}\log y - \frac{\alpha_3}{\alpha_1}\log\left(\frac{M}{P}(-1)\right) \qquad (B11)$$

Edwards and Khan (1985) derive an expression similar to ($B11$) for the interest rate that would correspond to a closed private capital account. They then substitute for i' in ($B2$) and estimate the resulting expression directly, with the key parameter given by the coefficient of the uncovered parity interest rate i^*. This procedure, however, is not widely applicable in developing countries because the dependent variable in the resulting equation to be estimated is the actual domestic market-clearing interest rate i. As already indicated, in repressed financial systems, it is the "curb market", or free interest rate, that is the valid empirical counterpart of i. In countries where adequate series for the curb market rate are available, the Edwards and Khan procedure (i.e., ($B11$) substituted in ($B2$) and the direct estimation of the degree of capital mobility) can be used.

Since published interest rate data for the informal loan market are rarely available, the procedure of Edwards and Khan is not readily applicable in a developing-country setting.[8] To circumvent this problem, one can proceed in two steps: First, substitute ($B6$) into ($B1$) to derive an expression for the unobserved informal market-clearing interest rate, i:

$$i = \varphi i^* - (1-\varphi)\left[\frac{\alpha_0}{\alpha_1} + \frac{1}{\alpha_1}\log\left(\frac{M'}{P}\right) - \frac{\alpha_2}{\alpha_1}\log y - \frac{\alpha_3}{\alpha_1}\log\left(\frac{M}{P}(-1)\right)\right] \qquad (B12)$$

This yields an expression for the unobserved market clearing interest rate, i, which can be substituted into the money-demand expression ($B9$) to obtain:

$$\log\left(\frac{M}{P}\right) = \lambda_0 + \lambda_1 i^* + \lambda_2 \log\left(\frac{M'}{P}\right) + \lambda_3 \log y + \lambda_4 \log\left(\frac{M_{-1}}{P_{-1}}\right), \qquad (B13)$$

where

$$\varphi_0 = -\alpha_0(1-\varphi),$$

$$\lambda_1 = \alpha_1\varphi < 0,$$

$$\lambda_2 = 1-\varphi; \quad 0 \leqslant \lambda_2 \leqslant 1,$$

$$\lambda_3 = \alpha_2\varphi > 0,$$

$$\lambda_4 = \alpha_3\varphi > 0.$$

All the variables in equation ($B13$) are directly observable in developing countries with repressed financial markets. Estimation of the equation can be conducted to obtain the structural parameters of the model. If the interest is in obtaining all the parameters of the model – those of the money demand function, the α's, as well as the degree of capital mobility, φ – estimation could be conducted using nonlinear instrumental variables. However, the degree of capital mobility alone can be estimated using only ordinary least squares.

4 Some Evidence from Developing Countries

The approach described above was used by Haque and Montiel (1990, 1991) to estimate the effective degree of interest arbitrage between the informal loan market and external financial assets in a diverse set of developing countries. Two separate approaches were used. First, country-by-country estimation was conducted using annual data from 15 developing countries for the years 1969 to 1987. The size of the sample provides some basis for judging the degree of capital mobility in developing countries more generally. The countries included in this sample were Indonesia, Malaysia, Philippines, Sri Lanka, India, Kenya, Tunisia, Morocco, Zambia, Uruguay, Guatemala, Brazil, Malta, Turkey, and Jordan. The choice of countries was dictated by the availability of internally consistent time series of reasonable length and the desire to maintain a geographical balance and to obtain a sample that was illustrative of various categories of developing countries. The sample consists of six Asian countries, four African countries, three Latin American countries and two European countries. In addition, the sample included four low-income countries and three heavily indebted countries.[9]

The dependent variable in the estimating equations was the log of the real money supply, measured as $M1$ divided by the consumer price index (CPI). The independent variables included the logs of lagged real money, real GNP, (i.e., GNP divided by the CPI), the real value of M' (M1 minus the domestic-currency value of private capital inflows and divided by the CPI), and an interest rate variable. The interest rate variable was defined by the uncovered interest parity condition. This variable was measured by the US Treasury Bill rate plus the expected depreciation in the exchange rate, which was proxied by the actual exchange rate change that took place between periods.

Since the specification included a lagged dependent variable ($\log(M_{-1}/P_{-1})$) as well as endogenous variables ($\log(M'/P)$ and $\log y$), and incorporated a rationally-expected variable, a generalized nonlinear instrumental variables procedure was used in the F5 estimation.[10] The instruments used were the lagged values of the external interest rate, real GDP, investment, the money supply, the consumer price index, imports, foreign exchange reserves, and industrial country real GNP.

The results of the nonlinear instrumental-variable estimation of equation ($B13$) are presented in table $B.1$ taken from Haque and Montiel (1991). The underlying model appears to fit the data fairly well. In almost all cases, the coefficient of the interest rate variable, α_1, and the income variable, α_2, were of the right sign and of conventional magnitudes, in keeping with those that are commonly available for most developing countries. The estimated semi-elasticity of the interest rate in the money demand function was less than one, in keeping with commonly available estimates.[11] While the estimates of the money-demand parameters were statistically significant at conventional levels only in a few cases, this is not surprising in view of the relatively small sample size. The estimated Q-statistics were consistent with the absence

Table *B*.1 Estimates of the parameters of the model[d]

Country	φ	α_1	α_2	α_3	α_0	*Q(3)*[e]
Brazil	0.72[c]	−0.01	0.25	0.48	2.23	0.46
	(8.7)	(−0.4)	(1.2)	(1.9)	(1.7)	
Guatemala	0.71[a]	−0.02	0.09	0.91	0.14	7.8
	(4.0)	(−1.6)	(0.1)	(3.0)	(0.1)	
India	0.16[b]	−0.02	3.2	−0.53	−0.17	3.6
	(1.4)	(−0.9)	(1.7)	(−0.58)	(−1.7)	
Indonesia	0.87[a]	−0.04	1.15	0.35	−7.1	2.6
	(8.5)	(−4.0)	(4.1)	(2.5)	(−3.5)	
Jordan	0.5[c]	−0.01	0.30	0.80	−0.13	3.3
	(2.6)	(−4.06)	(1.5)	(5.9)	(−0.8)	
Kenya	0.6[a]	−0.01	0.21	0.61	0.76	0.4
	(2.6)	(−0.5)	(0.4)	(2.2)	(0.37)	
Malaysia	0.6[a]	−0.01	0.56	0.65	−1.32	8.8
	(2.9)	(−0.9)	(2.1)	(4.3)	(−1.7)	
Malta	0.4[c]	−0.3	0.92	−0.09	0.77	4.5
	(3.1)	(−3.3)	(3.1)	(0.3)	(3.3)	
Morocco	0.88[a]	−0.01	0.25	0.85	−0.64	3.7
	(7.1)	(−2.7)	(1.0)	(6.3)	(−0.78)	
Philippines	0.58[a]	−0.01	0.58	0.53	−1.4	3.2
	(2.1)	(−0.2)	(1.6)	(2.0)	(−0.1)	
Sri Lanka	0.64[a]	−0.01	0.17	0.89	−0.38	0.4
	(3.2)	(−0.8)	(0.07)	(2.6)	(−0.0)	
Tunisia	0.83[a]	−0.0	0.61	0.5	−0.76	3.1
	(5.3)	(−1.3)	(3.6)	(3.7)	(−3.0)	
Turkey	0.53[a]	−0.03	1.54	−0.72	−0.21	3.5
	(2.5)	(0.0)	(0.8)	(0.8)	(1.7)	
Uruguay	0.89[a]	−0.0	0.23	0.87	−0.72	2.8
	(14.1)	(−0.3)	(0.4)	(9.6)	(−0.2)	
Zambia	1.02[a]	−0.1	0.35	0.21	0.58	2.0
	(16.1)	(−0.14)	(2.1)	(1.2)	(0.79)	

[a] φ significantly different from 0 and insignificantly different from 1. Perfect capital mobility cannot be ruled out.
[b] φ insignificantly different from 0 and significantly different from 1. Perfect capital immobility cannot be ruled out.
[c] φ significantly different from both 0 and from 1. Both polar cases of perfect capital mobility and immobility cannot be ruled out.
[d] Ratios of coefficients to standard errors in parantheses.
[e] The Q-statistic is calculated for three lags of residuals. It is $\chi^2(3)$ and at 95 percent confidence level, the critical value is 7.81.
Source: Haque and Montiel (1991)

of serial correlation in the residuals for every case but that of Malaysia.

The key result of the estimation, however, was the measurement of the degree of capital mobility, given by the magnitude and statistical significance of φ. In all but one case the point estimate of this parameter was found, as implied by the specification of equation ($B13$), in the interval between zero and one. In Zambia alone the point estimate was marginally different from one - a difference which could easily result from sampling error.

In ten out of the fifteen countries, φ was significantly different from zero and insignificantly different from one, indicating that one polar case - that of a completely financially-closed economy - could be ruled out, but not the other - a completely financially-open economy. In these cases the tendency was towards openness and a high degree of capital mobility. The government in these countries, therefore, has little control over domestic interest rates and the money supply. In four countries, φ was significantly different from both zero and one, ruling out both the polar opposites described above.[12] In these countries the informal interest rate would be subject to partial influence by the government, at least in the short run. In only one country, India, did the estimate of φ suggest that capital was immobile. The parameter estimate was small in magnitude, insignificantly different from zero, and significantly different from one.[13]

These results suggest that, over this sample period, domestic market-determined interest rates for this rather diverse group of developing countries tended to move quite closely with their uncovered-parity foreign counterparts. Domestic interest rates were relatively less influenced by domestic financial developments - except to the extent that the latter were expected to lead to exchange-rate adjustments. We say "on average" because the effective degree of capital mobility was found to differ across countries and could have varied over time for some of the countries in the sample. However, the fairly uniform value of the estimate of φ across a large group of developing countries, as well as the absence of serial correlation in all but one equation, suggest that variations in the degree of capital mobility for individual countries, while undoubtedly present, are unlikely to have been large.

Further evidence on the degree of capital mobility that prevails

in developing countries in general is developed in Haque and Montiel (1990), who estimated the model described above for pooled cross-section, time-series data for a large group of countries. In this case, the sample consisted of 27 countries with data for the years 1969 to 1987. They used a fixed effects model for the estimation of equation ($B13$) with the pooled data. The estimated coefficients and the associated t-ratios were as follows:

Coefficient	*Estimate*	*t-statistic*
φ	0.933	11.5
α_1	−0.005	−2.7
α_2	0.207	4.2
α_3	0.837	26.3
$R^2 = 0.97$	$F = 4660.6$	

Once again the model seems to explain the data rather well, with all the key parameters being of the right sign and magnitude. The average degree of capital mobility in this group of developing countries is again given by the magnitude and statistical significance of φ. The estimate turns out to be about 0.93 and was quite precisely estimated with a relatively small standard error. However, this coeficient was close enough to unity for it not to be distinguishable from it. Yet it was significantly different from zero, or perfect capital immobility. Consequently, the assumption of a financially closed economy (perfect capital immobility) can be ruled out in developing countries while that of financial openness (perfect capital mobility) cannot be ruled out.

5 Implications for Modeling of the Informal Sector

At the very least, the evidence examined here suggests that market-determined interest rates in informal markets can be expected to be influenced by rates of return on assets denominated in foreign exchange, as assumed in the models constructed in chapters 3 and 4. In fact, the extreme assumption of perfect arbitrage between the two types of asset is more consistent with the data than that of perfect insulation of interest rates on domestic assets. Since domestic market-determined (informal) interest rates appear to be so heavily influenced by the rates of return available on external assets it is clearly improper to exclude foreign-exchange denominated assets from the menu of portfolio choices available to domestic savers in models of the informal loan market. This implies that the premium on foreign exchange and the informal loan market interest rate must be determined simultaneously.

Notes

1 See the *Annual Report on Exchange and Trade Restrictions* published by the International Monetary Fund for a description of capital controls in developing countries.
2 See Khan and Haque (1985, 1987).
3 See the review by Kreinin and Officer (1978).
4 See also Edwards and Khan (1985).
5 In the case of industrial countries, of course, direct tests of arbitrage relationships such as (*B*1) or even (*B*2) are feasible. As Edwards and Khan (1985) show, it is possible to undertake direct tests in the case of those developing countries where market-determined interest rates pertain to the formal sector.
6 See Kouri and Porter (1974).
7 Note that the lagged dependent variable in equation (*B*6) is $\log(M_{-1}/P_{-1})$. The reason is that the current demand for money in equation (*B*9) depends on the actual money stock in the previous period rather than on the money stock that would hypothetically have emerged with zero cumulative private capital mobility up to the previous period.
8 Edwards and Khan limited their study to Colombia and Singapore, for which market-determined interest rates were available.
9 Low income countries are defined in the *World Economic Outlook* of the International Monetary Fund as those with per capita income less than $425. In our sample these include India, Kenya, Sri Lanka, and Zambia. The three heavily-indebted countries in our sample (Brazil, Morocco, and the Philippines) are drawn from the category of the fifteen heavily indebted countries as described in the *World Economic Outlook*.
10 The assumption of rational expectations also complicates estimation. As Wickens (1982) has shown, instrumental variables can be used in an errors in variables approach to estimating the rationally expected variable.

11 Note that the semi-elasticity is obtained by multiplying α_1 by 100, since the interest rate was entered in terms of percentage points.
12 In these cases the point estimate is also generally smaller in magnitude than the point estimate in countries where perfect capital mobility cannot be ruled out.
13 On the basis of generally-accepted priors, India would have been one of the countries whose capital immobility would have been expected, as is in fact suggested by the test. In general, the test seems to have discriminated fairly well, relative to generally-accepted priors for the group of countries in the sample.

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