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Nominal exchange rate flexibility and real exchange rate adjustment: New evidence from dual exchange rates in developing countries

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Abstract

This study investigates whether greater nominal exchange rate flexibility aids real exchange rate adjustment based on data from dual exchange rates in developing countries. Specifically, we analyze whether the more flexible parallel market rate produces faster real exchange rate adjustment than the less flexible official rate does. Half-life estimates of adjustment speeds are obtained from fractional time series analysis. We find no systematic evidence that greater exchange rate flexibility tends to produce either faster or slower real exchange rate adjustment, albeit there is substantial cross-country heterogeneity in speed estimates. With official rates pegged to the dollar, many developing countries use parallel exchange markets as a back-door channel to facilitate real exchange rate adjustment. The evidence suggests, however, that these parallel markets often fail to speed up real rate adjustment.

JEL classification : C22; F31

Keywords: Real exchange rate; Fractional time series; Half life; Adjustment speed

1. Introduction

The role of exchange rate flexibility in economic adjustment has long been a hotly contested issue. The issue brings back the old debate between Nurkse (1944) and Friedman (1953) on the

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stabilizing or destabilizing nature of speculation. According to Nurkse (1944), speculative exchange rate movements would tend to amplify and prolong disequilibria rather than accelerate economic adjustment. In contrast, Friedman (1953) suggested that it would be easier for the economy to adjust to shocks by securing needed changes in the real exchange rate through exchange rate than through price adjustment. Instead of fearing the instability flexible rates might bring, speculative forces could actually quicken exchange rate adjustment and hasten the equilibrating process. Since the move to the modern float, real exchange rate changes seem to have grown more persistent. As Rogoff (1996) observes, the speed of real exchange rate adjustment has been glacially slow among industrialized countries under the current float. The issue then arises as to whether greater nominal rate flexibility promotes real rate adjustment.

In this study, we provide new alternative evidence on the issue in exchange rate flexibility based on a special set of data on dual exchange rates from developing countries. We do not analyze exchange rate flexibility in usual terms of exchange rate regimes (floating as opposed to fixed), but in terms of dual exchange rates (market-determined as opposed to government-set). Not only during the Bretton Woods (BW) period but also afterwards, parallel markets for foreign exchange – especially for the U.S. dollar – were common among developing countries. Unlike the official rate, which is fixed and occasionally reset by the relevant monetary authority, the parallel rate is determined by market supply and demand in which speculative forces can play a significant role. With limited access to the official exchange market, the parallel market serves to meet unsatisfied demand for foreign currency. Many developing countries use the dual exchange rate system as a tool to stabilize the real economy and to insulate real economic activity from the volatility of financial markets (Pozo and Wheeler, 1999). As noted by Reinhart and Rogoff (2004), parallel exchange rates provide a form of ‘‘back-door’’ floating in a lot of countries where an official peg is adopted.

The spread between the parallel and the official exchange rate – referred to as the parallel market premium – often works as an indicator of exchange rate misalignments and has been used as a guide to realigning the official rate. According to Reinhart and Rogoff (2004), the parallel exchange rate is ‘‘a far better barometer of monetary policy than is the official exchange rate’’ and that the parallel market premium often correctly predicts realignments in the official rate and anticipates future official rate changes. Earlier studies of the parallel market premium (Dornbusch et al., 1983; Kamin, 1993; Montiel and Ostry, 1994; Pozo and Wheeler, 1999) also suggest an important role for the expectations of future official rate changes in driving the premium. Ghei and Kamin (1999) recognize that the parallel exchange rate is a good, though not entirely perfect, proxy for the free-market currency value.

To examine whether greater exchange rate flexibility leads to quicker real exchange rate adjustment, this study analyzes data on dual exchange rates for 24 developing countries. For each of these countries, a parallel market for foreign exchange exists alongside the official one during both BW and post-BW periods. With official and parallel rates being available for the same historical period, this special data set permits intra-period analysis. We can evaluate the relative adjustment speed of the real official and the real parallel rate for each country within a given time period. This minimizes the need to control for any inter-period differences in global and domestic economic conditions. Two questions of interest are: Do the official and the parallel market rate revert toward one another over time? Does the flexible parallel market rate bring about a faster speed of real exchange rate adjustment than the less flexible official rate?

In addition to analyzing the difference in adjustment speed between real official and parallel rates on a country-by-country basis, this study also shows that the speed of adjustment for either rate can vary considerably across countries, even within the same historical period. A cross-section analysis will be conducted to evaluate how much the cross-country variation in adjustment speeds is attributable to structural differences in the corresponding economies.

2. Empirical methodology

The adjustment dynamics of economic processes will be evaluated based on fractional time series models. Fractionally integrated processes can capture a wider variety of adjustment dynamics than usual time series processes (Diebold et al., 1991; Cheung and Lai, 1993). Indeed, fractionally integrated processes provide a better approximation to the Wold decomposition of time series dynamics than conventional processes (Granger and Joyeux, 1980; Hosking, 1981). The use of fractional time series models is particularly attractive in our application here. By permitting the order of integration to be a non-integer, fractional models cover a more general class of mean-reverting processes than usual autoregressive moving-average (ARMA) models, on which standard unit-root tests are all based. Such fractional integration models enable us to uncover a broad range of subtle mean-reverting dynamics, which can neither be described by ARMA models nor be detected by standard unit-root tests. Moreover, developing countries are known to have much greater cross-country heterogeneity in real exchange rate dynamics than developed countries. The additional modeling flexibility gained from using fractional integration models is thus useful for accommodating the substantial heterogeneity in real exchange rate dynamics. As reported later, our results do indeed confirm the presence of significant cross-country variation in real exchange rate behavior.

Fractionally integrated processes are in general represented by

$$B(L)(1-L)^d y_t = D(L)v_t \tag{1}$$

where y_t is the time series under consideration, L the standard lag operator, $B(L) = 1 - \beta_1 L - \ldots - \beta_p L^p$, $D(L) = 1 - \delta_1 L - \ldots - \delta_q L^q$, all roots of $B(L)$ and $D(L)$ are stable, v_t is the random error term, and the fractional differencing part is

$$(1-L)^d = \sum_{m=0}^{\infty} \frac{\Gamma(m-d)L^m}{\Gamma(m+1)\Gamma(-d)} \tag{2}$$

with $\Gamma(\cdot)$ being the gamma function. This model describes a broad class of time series processes known as autoregressive fractionally integrated moving-average (or ARFIMA (p, d, q)) processes. It extends the standard ARMA (p, q) and ARIMA (p, 1, q) models by permitting non-integer values of d. Such extended flexibility in modeling dynamics can be important for a proper evaluation of economic adjustment. The long-run reversion property of y_t is determined by the fractional integration parameter, d. Cheung and Lai (1993) show that mean reversion occurs so long as $d < 1$.

In our statistical analysis, a frequency-domain maximum likelihood procedure is used to estimate the ARFIMA (p, d, q) model. Following Fox and Taqqu (1986), we utilize the property that maximizing the likelihood function is asymptotically equivalent to:

$$\text{Minimizing} \sum_{k=1}^{T-1} \frac{I_y(2\pi k/T)}{f_y(2\pi k/T;\xi)} \text{ with respect to } \xi = (d, \beta_1, \ldots, \beta_p, \delta_1, \ldots, \delta_q) \tag{3}$$

where

$$f_y(\lambda_k, \xi) = |1 - \exp(i\lambda_k)|^{-2d} |B^{-1}(\exp(-i\lambda_k)) D(\exp(-i\lambda_k))|^2 \tag{4}$$

$$I_y(\lambda_k) = \frac{1}{2\pi T} \left| \sum_{t=1}^{T} y_t \exp(i\lambda_k t) \right|^2 = \frac{1}{2\pi T} \left[\left(\sum_{t=1}^{T} y_t \cos(\lambda_k t) \right)^2 + \left(\sum_{t=1}^{T} y_t \sin(\lambda_k t) \right)^2 \right] \tag{5}$$

with $\lambda_k = 2\pi k/T$ and i being the imaginary part of the complex number. Specifically, $I_y(\lambda_k)$ gives the periodogram of y_t at the kth Fourier frequency, and $f_y(\lambda_k, \xi)$ is proportional to the spectral density of y_t at frequency λ_k. The resulting estimator for d is consistent and has an asymptotic normal distribution (Tanaka, 1999). Cheung and Diebold (1994) show that this maximum likelihood procedure for estimating fractional processes has good finite-sample properties. Cheung and Lai (2001) provide an empirical application of this fractional model estimation procedure.

3. Preliminary analysis on the parallel market premium

Before analyzing the relative adjustment speed of the real official and the real parallel exchange rate, we first address some known issues concerning the behavior of the parallel market premium. Let e_{Ot} and e_{Pt} be, respectively, the official and the parallel market rate in logarithms. The parallel market premium is given by $e_{\mathrm{Pt}} - e_{\mathrm{Ot}}$. Although infrequently done, a monetary authority may adjust the official rate to reduce currency misalignments. If the parallel market premium is a good gauge of the degree of misalignment and a useful guide for setting the official rate, the official and parallel rates should not drift too far apart. It follows that a long-run stationary relationship likely exists between these two rates. Theoretical models – including portfolio balance models (Dornbusch et al., 1983; Phylaktis, 1991) and monetary models (Phylaktis, 1996; Kouretas and Zarangas, 1998) – also suggest the parallel market premium be stationary (see also Appendix A).

This study examines time series data from 24 developing countries for which exchange rate and price data are available for both BW and post-BW periods. Data on both monthly consumer prices and official exchange rates were obtained from the IMF's International Financial Statistics database. Data on parallel rates were taken from Reinhart and Rogoff's (2004) dataset, which was compiled from various issues of *Pick's Currency Yearbook* and *Pick's World Currency Report* (later became the *World Currency Yearbook*). All exchange rates are expressed in units of foreign currency per U.S. dollar. Dictated by data availability, the sample data cover the period January 1957 through December 1998.[1] The countries under study include Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, Egypt, El Salvador, India, Israel, Korea, Malaysia, Mexico, Morocco, Pakistan, Paraguay, Philippines, South Africa, Sri Lanka, Thailand, Uruguay, and Venezuela. Even during the post-BW period, these countries chose to adopt official pegs or crawls (with fixed or limited flexibility in rates) against the U.S. dollar most of the time.[2]

[1] The *World Currency Yearbook* had stopped being published after 1998. In addition, due to limited data availability of parallel market rates, a slightly shorter sample period that began later than 1957 was used in the cases of Dominican Republic, El Salvador, Morocco, and Venezuela. For several other countries (Brazil, Costa Rica, Korea, and Thailand), moreover, consumer price data were not available for the sample period. In these cases, we used either wholesale or producer price data.

[2] The Reinhart–Rogoff dataset also contained exchange rate data for developing countries that had their official rates pegged against some foreign currencies other than the U.S. dollar. We chose not to consider them so that we would not need to control for any differences in the base currency in our cross-country analysis.

Table 1
The mean absolute size of nominal exchange rate movements

Country	Bretton Woods		Post-Bretton Woods	
	Official rate	Parallel rate	Official rate	Parallel rate
Argentina	2.38	4.02	8.24	9.40
Bolivia	0.59	1.59	4.20	5.47
Brazil	3.09	4.61	9.12	9.78
Chile	2.42	5.42	2.56	3.57
Colombia	1.25	2.44	1.61	2.27
Costa Rica	0.09	1.55	1.52	2.83
Dominican Republic	0.00	3.19	1.47	2.38
Ecuador	0.27	1.97	1.96	3.00
Egypt	0.11	3.83	0.77	2.74
El Salvador	0.00	1.25	0.49	3.94
India	0.45	2.85	1.31	2.84
Israel	0.44	1.83	3.50	4.98
Korea	1.13	3.08	1.06	2.59
Malaysia	0.16	0.62	1.26	1.73
Mexico	0.00	0.03	2.71	3.46
Morocco	0.16	1.98	1.82	2.58
Pakistan	0.56	3.07	0.58	2.35
Paraguay	0.39	1.38	1.19	2.67
Philippines	0.70	2.41	1.28	2.79
South Africa	0.19	1.51	2.10	5.31
Sri Lanka	0.22	3.62	1.54	2.33
Thailand	0.04	0.99	0.83	2.46
Uruguay	3.14	4.52	3.39	3.91
Venezuela	0.24	0.33	1.89	3.34

All the numbers reported in this table are expressed in percentage terms.

To allow for any possible changes in real exchange rate behavior across historical periods, we will examine the BW and the post-BW period separately. Our study builds on the basic premise that official rates are less flexible than parallel rates. The dual-rate analysis requires neither an assumption of official exchange rates being absolutely fixed over time nor an assumption of parallel exchange rates being entirely flexible.[3] Table 1 shows the average size (measured as the mean absolute change) of official and parallel rate movements over the two time periods. The data are consistent with the conventional wisdom that market-determined parallel rates move more freely than government-set official rates.

Table 2 reports some descriptive statistics showing the average size and the standard deviation of the parallel market premium during the BW and post-BW periods. The parallel market premium varied greatly both in size and in variability across countries, even within the same historical period. Moreover, there was no systematic pattern of change in the average premium size across the two time periods. The average parallel market premium rose in 9 countries but fell in 15 countries when moving from the BW to the post-BW period. The data also revealed no systematic pattern of cross-period change in variability, with increases for 10 countries and decreases for 14 countries from the BW to the post-BW period.

[3] Most nominal exchange rates are neither completely fixed nor completely flexible over time. This is true for both the BW period and the post-BW period. This is also true for both developing and industrial countries.

Table 2
The average size and variability of the parallel market premium

Country	Bretton Woods		Post-Bretton Woods	
	Mean	S.D.	Mean	S.D.
Argentina	11.60	23.60	26.34	35.25
Bolivia	12.83	15.91	23.50	47.61
Brazil	10.26	14.34	21.06	21.26
Chile	40.17	49.46	15.24	14.72
Colombia	21.22	18.37	7.68	6.80
Costa Rica	20.68	11.46	13.60	12.71
Dominican Republic	24.38	9.07	23.90	23.23
Ecuador	15.82	7.80	19.88	21.81
Egypt	74.99	23.84	42.39	39.10
El Salvador	16.60	6.64	41.09	33.35
India	38.05	15.88	12.49	7.20
Israel	16.92	9.37	9.10	12.53
Korea	34.94	32.38	4.40	6.52
Malaysia	1.95	1.91	0.63	2.04
Mexico	-0.08	0.04	8.37	13.85
Morocco	12.05	6.87	4.39	4.30
Pakistan	54.60	18.31	12.98	10.66
Paraguay	14.06	13.53	31.58	34.18
Philippines	22.36	25.88	6.36	6.49
South Africa	6.87	4.51	9.16	9.58
Sri Lanka	67.59	28.46	20.34	20.13
Thailand	1.26	2.10	0.13	3.39
Uruguay	35.56	48.68	8.77	10.52
Venezuela	10.41	15.55	24.69	38.72

The parallel market premium is measured in percentage terms.

Table 3 summarizes the test results from fractional integration analysis by country. The unit-root null hypothesis of $d = 1$ is tested against the mean-reverting alternative of $d < 1$. For the BW period, the unit-root hypothesis can be rejected for all but one of the 24 countries at the 5 percent significance level. The post-BW data also widely reject the unit-root hypothesis. In all but two cases can the unit-root hypothesis be rejected at the 5 percent level. In general, the test results strongly support that the differential between the official and the parallel market rate is stationary. Indeed, our findings suggest that these two exchange rates are fractionally cointegrated—a more general notion of cointegration than what has been considered in previous studies (e.g., Booth and Mustafa, 1991; Kouretas and Zarangas, 1998).[4]

4. Adjustment speeds of real official and parallel exchange rates

The adjustment behavior of real official and real parallel exchange rates is examined next. We first check the stationarity of individual real exchange rate series. If PPP holds in the long run, the

[4] There is a vast literature on dual exchange rates. In addition to the long-run relationship between official and parallel market rates, previous studies have examined additional issues, including market efficiency, causality, volatility transmission, capital controls, and purchasing power parity. We will move away from these known issues and explore a different issue, namely, the relative speed of real rate adjustment produced by dual exchange rates.

Table 3
Testing for mean reversion in the parallel market premium

Country	Bretton Woods		Post-Bretton Woods	
	$d-1$	t-Stat	$d-1$	t-Stat
Argentina	0.00	(0.01)	−0.20	(−2.83)*
Bolivia	−0.06	(−3.54)*	−0.38	(−4.21)*
Brazil	−0.64	(−5.16)*	−0.32	(−2.94)*
Chile	−0.06	(−1.98)*	−0.41	(−3.59)*
Colombia	−0.35	(−10.58)*	−1.09	(−12.58)*
Costa Rica	−0.79	(−6.39)*	−0.23	(−3.00)*
Dominican Republic	−0.27	(−3.48)*	−0.99	(−6.92)*
Ecuador	−0.30	(−4.85)*	−0.20	(−2.38)*
Egypt	−1.16	(−5.34)*	0.09	(1.00)
El Salvador	−0.04	(−2.20)*	−0.12	(−1.99)*
India	−1.06	(−7.91)*	−1.08	(−8.80)*
Israel	−1.01	(−5.20)*	−0.09	(−1.15)
Korea	−0.84	(−42.63)*	−1.33	(−26.61)*
Malaysia	−0.38	(−6.63)*	−1.35	(−3.43)*
Mexico	−0.25	(−8.54)*	−0.64	(−8.62)*
Morocco	−0.32	(−4.12)*	−1.44	(−13.10)*
Pakistan	−1.41	(−5.97)*	−0.18	(−3.56)*
Paraguay	−0.89	(−4.44)*	−0.44	(−4.12)*
Philippines	−0.27	(−3.35)*	−1.10	(−12.42)*
South Africa	−0.94	(−6.62)*	−0.41	(−8.74)*
Sri Lanka	−0.08	(−2.32)*	−0.46	(−7.77)*
Thailand	−1.26	(−7.84)*	−1.36	(−21.38)*
Uruguay	−0.98	(−7.08)*	−0.33	(−3.82)*
Venezuela	−1.44	(−7.71)*	−0.77	(−3.78)*

The unit-root null hypothesis of $d-1=0$ is tested against the mean-reverting fractional alternative of $d-1<0$. The numbers in parentheses are the t-statistics for the corresponding estimates. Statistical significance is indicated by an asterisk (*) for the 5 percent level.

real official and the real parallel exchange rate should exhibit mean reversion. In analyzing the adjustment behavior of real exchange rates in industrial countries, previous studies generally report greater difficulty in finding mean reversion in flexible-rate than in fixed-rate data. In terms of historical periods, it also seems much harder to detect mean reversion during the post-BW period as opposed to other historical periods. Do similar results apply to the behavior of real exchange rates in developing countries? Compared to those for industrial countries, empirical findings for developing countries have been relatively limited and less extensive. In this regard, our results from dual exchange rates in developing countries may offer an alternative perspective on the dynamics of real exchange rate adjustment.

Table 4 contains the results from tests for stationarity in real exchange rates. For the BW period, the unit-root hypothesis can be rejected at the 5 percent significance level in 18 out of 24 cases for real official rates and in 19 out of 24 cases for real parallel rates. For the post-BW period, the unit-root hypothesis can be rejected in 18 out of 24 cases for real official rates and in 17 out of 24 cases for real parallel rates. Hence, the real official and the real parallel exchange rate show little difference in terms of the number of unit-root rejection cases. In the large majority of cases, both real official and parallel rates exhibit mean reversion. This applies to the BW and the post-BW period alike.

Table 4
Testing for mean reversion in real official and parallel exchange rates

Country	Bretton Woods				Post-Bretton Woods			
	Real official rate		Real parallel rate		Real official rate		Real parallel rate	
	$d-1$	t-Stat	$d-1$	t-Stat	$d-1$	t-Stat	$d-1$	t-Stat
Argentina	−0.18	(−6.10)*	−0.14	(−2.05)*	−0.25	(−3.08)*	−0.14	(−2.83)*
Bolivia	−0.75	(−6.88)*	−0.06	(−1.92)	−0.50	(−8.63)*	−0.07	(−0.68)
Brazil	−1.21	(−8.56)*	−0.10	(−2.07)*	−0.74	(−6.22)*	−0.16	(−2.90)*
Chile	−0.95	(−4.73)*	0.05	(0.86)	0.06	(1.18)	−0.13	(−1.48)
Colombia	−0.19	(−3.31)*	−0.83	(−6.64)*	0.12	(1.71)	0.05	(0.99)
Costa Rica	−0.04	(−0.49)	−0.85	(−8.59)*	−1.02	(−5.09)*	−0.14	(−0.80)
Dominican Republic	−1.09	(−6.65)*	−1.02	(−7.48)*	−0.14	(−3.95)*	−0.93	(−12.62)*
Ecuador	−0.10	(−2.07)*	−0.03	(−0.73)	−0.11	(−1.39)	−0.24	(−3.65)*
Egypt	−0.82	(−5.74)*	−1.04	(−8.76)*	0.06	(0.84)	−0.20	(−3.04)*
El Salvador	−0.07	(−0.88)	−0.82	(−5.64)*	−0.92	(−15.97)*	−0.87	(−4.64)*
India	−0.94	(−3.53)*	−1.14	(−5.63)*	−0.74	(−3.27)*	−0.73	(−3.94)*
Israel	−0.10	(−1.64)	−0.70	(−4.24)*	−0.07	(−2.85)*	−0.26	(−17.18)*
Korea	−1.05	(−4.41)*	−0.92	(−9.24)*	−0.95	(−7.95)*	−1.02	(−4.88)*
Malaysia	−0.78	(−4.51)*	−0.11	(−2.78)*	−1.42	(−4.42)*	−0.92	(−6.75)*
Mexico	−0.03	(−0.81)	−0.03	(−24.32)*	−0.85	(−8.05)*	−0.10	(−2.22)*
Morocco	0.04	(−0.58)	−0.34	(−4.28)*	0.07	(1.39)	−0.03	(−0.26)
Pakistan	−0.03	(−3.94)*	−0.33	(−3.93)*	0.09	(1.64)	−1.17	(−7.56)*
Paraguay	0.02	(0.57)	−0.03	(−2.28)*	−0.04	(−4.59)*	−0.02	(−17.43)*
Philippines	−0.68	(−5.10)*	0.03	(0.58)	−0.80	(−4.07)*	−0.87	(−4.16)*
South Africa	−0.19	(−4.12)*	−1.29	(−5.78)*	−0.26	(−3.64)*	−0.19	(−4.08)*
Sri Lanka	−0.03	(−5.43)*	−0.06	(−1.92)	−0.14	(−2.40)*	−0.03	(−0.22)
Thailand	−1.03	(−11.99)*	−1.10	(−8.08)*	−0.96	(−12.34)*	−1.05	(−29.30)*
Uruguay	−1.09	(−4.96)*	−0.84	(−4.57)*	−0.75	(−7.20)*	0.04	(0.52)
Venezuela	−0.06	(−2.00)*	−0.26	(−2.97)*	−0.15	(−2.89)*	−0.15	(−2.54)*

The unit-root null hypothesis of $d-1=0$ is tested against the mean-reverting fractional alternative of $d-1<0$. The numbers in parentheses are the t-statistics for the corresponding estimates. Statistical significance is indicated by an asterisk (*) for the 5 percent level.

4.1. Intra-period comparison of adjustment speeds by country

We next evaluate how fast the real official and the real parallel exchange rate adjust to shocks. To analyze whether the flexible parallel market rate produces quicker real exchange rate adjustment than the pegged official rate, their adjustment speeds – measured in terms of half-lives of shocks to real exchange rates – are computed using impulse response analysis. Based on the ARFIMA model, as described in (1), the half-life can be estimated from a moving average representation of the time series process:

$$(1-L)y_t = A(L)v_t \tag{6}$$

where $A(L) = 1 + \alpha_1 L + \alpha_2 L^2 + \alpha_3 L^3 + \ldots$ derived from

$$A(L) = (1-L)^{1-d}\Phi(L) \tag{7}$$

with $\Phi(L) = B^{-1}(L)D(L)$. The moving average coefficients $\{\alpha_1, \alpha_2, \alpha_3, \ldots\}$ are often referred to as impulse responses, and they track and measure how much the real exchange rate adjusts in subsequent periods after a unit shock. The calculated half-life indicates how long it takes for the impact of a unit shock to the real exchange rate to dissipate by half.

The question at issue is whether greater exchange rate flexibility facilitates faster real exchange rate adjustment. The conventional argument is that, in a world of sticky prices, the speed at which the real exchange rate adjusts should depend a lot on exchange rate flexibility. Under pegged exchange rates, real exchange rates are expected to adjust at a slow pace limited by price stickiness. Although exchange rate realignments may sometimes hasten adjustment, they occur infrequently. With flexible exchange rates, on the other hand, real exchange rates can adjust quickly through immediate changes in nominal exchange rates. Despite the intuitive appeal of the argument, its empirical validity remains open to question.

In our analysis of dual exchange rates, the adjustment speed of the real official exchange rate is compared to that of the real parallel market exchange rate. If greater nominal exchange rate flexibility really promotes real exchange rate adjustment, the more flexible parallel market rate should yield a significantly faster speed of real exchange rate adjustment than does the less flexible official rate. Empirically, different historical periods could have rather different global and domestic conditions that would alter real exchange rate behavior. To the extent that the adjustment speed of the real official rate is compared against that of the real parallel rate in the same country over the same historical period, our intra-period analysis here averts the need to control for any differences in economic conditions.

Table 5 presents the half-life estimates for both real official and parallel exchange rates of individual countries. Evidently, the empirical results show no systematic pattern for the country group as a whole. While the estimated half-life of the real parallel rate can differ significantly from that of the real official rate in a given country, there is no identifiable general pattern in the half-life difference between the two real rates. Interestingly, real parallel rates appear as likely to yield a shorter half-life as real official rates. In about half of the cases, real parallel rates may actually adjust more slowly rather than more quickly in comparison to real official rates. Qualitatively similar findings are obtained, regardless of whether the BW or the post-BW period is considered. Overall, there is no consistent evidence to support that greater exchange rate flexibility tend to generate faster or slower real exchange rate adjustment.

The foregoing analysis has been based on point estimates of half-lives. As discussed later, some recent studies (Cheung and Lai, 2000; Murray and Papell, 2002) investigate the potential uncertainty in estimating half-lives of real exchange rates. The allowance for uncertain half-life measurements will further reinforce our conclusion that there is little evidence of any systematic relationship between nominal exchange rate flexibility and the speed of real exchange rate adjustment.

4.2. *Departures from consensus estimates*

Previous studies of real exchange rates typically report slow adjustment speeds with half-lives estimated to be 3–5 years. When describing the ‘‘consensus’’ on half-life estimates as remarkable and puzzling, Rogoff (1996) points out the difficulty in reconciling the slow adjustment speed of real exchange rates with their immense short-term volatility. Although standard exchange rate models with sticky prices may account for the huge short-term volatility under monetary shocks, the consensus half-life estimates are, as Rogoff (1996) notes, still too slow to be explained by nominal rigidities. On the other hand, the slow adjustment speed may reflect the influence of real shocks. If real shocks are predominant, they can exert significant long-lasting effects on the real

Table 5
Estimates of adjustment speeds for real official and parallel rates

Country	Bretton Woods		Post-Bretton Woods	
	Real official rate (half-life)	Real parallel rate (half-life)	Real official rate (half-life)	Real parallel rate (half-life)
Argentina	13.97	9.86	1.29	5.31
Bolivia	5.16	100+	0.15	0.11
Brazil	0.61	35.30	2.08	4.88
Chile	0.73	∞	∞	100+
Colombia	6.25	1.12	∞	∞
Costa Rica	100+	1.95	0.57	80.51
Dominican Republic	0.69	0.44	7.02	2.84
Ecuador	58.83	100+	34.52	13.61
Egypt	2.56	0.48	∞	1.04
El Salvador	34.33	1.91	0.64	1.37
India	1.16	0.57	2.27	1.48
Israel	34.61	1.73	46.33	0.50
Korea	1.06	1.76	1.52	0.91
Malaysia	2.07	23.56	1.22	1.45
Mexico	100+	100+	2.65	59.28
Morocco	∞	0.73	∞	100+
Pakistan	100+	0.71	∞	0.32
Paraguay	100+	100+	100+	100+
Philippines	1.58	∞	1.89	1.11
South Africa	1.60	0.39	5.70	1.80
Sri Lanka	100+	100+	4.27	100+
Thailand	0.73	0.64	1.49	0.70
Uruguay	0.46	1.07	4.60	∞
Venezuela	100+	2.09	4.61	100+

All half-life estimates are expressed in years. For half-life estimates that are longer than 100 years, they are indicated by "100+" in the table. In cases in which the real exchange rate process yields an estimated integration order greater than one (i.e., $d > 1$), it is indicated by "∞" and considered as having an infinite half-life.

exchange rate. Nevertheless, existing models based on real shocks cannot account for short-term exchange rate volatility.

The remarkable consensus highlighted by Rogoff (1996) comes mainly from real exchange rate studies of industrial countries. Departing from the consensus range, the half-life estimates obtained from the study here are much more dispersed. As shown in Fig. 1, the half-life estimates for both real official and parallel exchange rates show substantial variation among the developing countries under study. In a large majority of cases, these half-life estimates fall far outside Rogoff's (1996) consensus range of 3–5 years. Indeed, the half-life estimates are as likely to be higher as to be lower than the previous consensus estimates. This applies to the BW and the post-BW period alike. For the BW period, the estimated half-lives of the real official and parallel rates are shorter than 2 years in 9–13 cases and longer than 6 years in 10–13 cases. For the post-BW period, the estimated half-lives of the real official and parallel rates are less than 2 years in 7–11 cases and more than 6 years in 9–10 cases.

All in all, the half-life consensus reported among previous studies on real exchange rates in industrial countries fails to prevail among developing countries, which display much greater heterogeneity. For a number of developing countries, the results may seem consistent with relatively fast adjustment speeds as suggested by sticky-price models. For other developing

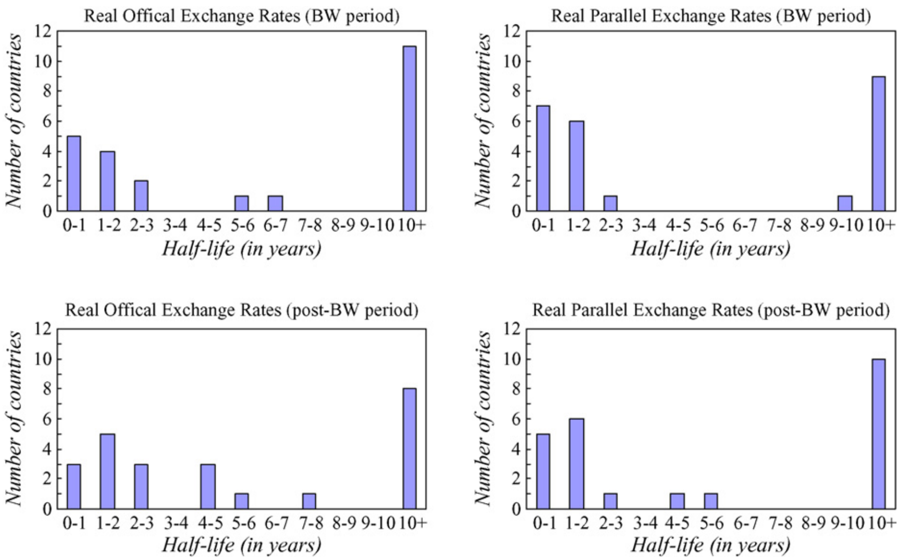

Fig. 1. Cross-country differences in half-life estimates.

countries, the data may yield sluggish rates of real exchange rate adjustment that are so slow that there is little long-run reversion. These findings are robust with respect to whether the real official or the real parallel exchange rate is considered.

4.3. Increased uncertainty in adjustment speed measurements

In analyzing the speed of real exchange rate adjustment during the post-BW period, Cheung and Lai (2000) evaluate the sample half-life measure and its estimation accuracy. To quantify the inevitable imprecision with which the adjustment speed is estimated, confidence intervals for half-life estimates are computed for several major industrial countries. These confidence intervals are found to be wide, indicating a high level of uncertainty in measuring the half-life of real exchange rate adjustment. The lower-bound estimates generally come in with half-lives less than 2 years, but the upper-bound estimates contain half-lives that last 7–9 years. Murray and Papell (2002) examine the post-BW data from 20 industrial countries and find even greater uncertainty in half-life measurements. While the lower-end estimates can reach as low as just one year, the higher-end estimates are often infinitely large.

In this study, we provide alternative evidence from the experience of developing countries with dual exchange rates over both the BW and post-BW periods. To explore whether the previous results on uncertain half-lives may similarly be observed in the data for developing countries, the 95 percent confidence interval for each individual half-life measurement is computed by Monte Carlo simulation. Using the estimated ARFIMA process for the respective data series as the data generating process, the half-life confidence interval is constructed based on 10,000 simulation replications in each case.

Table 6 gives the lower and upper bounds (denoted respectively by $LB_{95\ \text{percent}}$ and $UB_{95\ \text{percent}}$) of the half-life confidence interval. According to the lower-bound estimates, most half-lives can be as short as within a year for both real official and parallel rates. The upper-bound estimates of half-lives, in contrast, show no general pattern. Unlike those estimates previously reported for industrial

Table 6
Sampling uncertainty in half-life estimates

Country	Bretton Woods		Post-Bretton Woods	
	Real official rate [$LB_{95\ percent}$, $UB_{95\ percent}$]	Real parallel rate [$LB_{95\ percent}$, $UB_{95\ percent}$]	Real official rate [$LB_{95\ percent}$, $UB_{95\ percent}$]	Real parallel rate [$LB_{95\ percent}$, $UB_{95\ percent}$]
Argentina	[0.39, ∞]	[0.71, 100+]	[0.31, 17.94]	[0.31, 100+]
Bolivia	[0.63, 5.60]	[0.99, ∞]	[0.10, 0.25]	[0.08, ∞]
Brazil	[0.30, 0.71]	[0.51, 100+]	[0.79,17.07]	[0.45, 100+]
Chile	[0.32, 0.81]	[33.32, ∞]	[100+, ∞]	[0.89, ∞]
Colombia	[0.37, 100+]	[0.65, 2.10]	[3.78, ∞]	[100+, ∞]
Costa Rica	[1.75, 100+]	[1.19, 2.38]	[0.23, 4.03]	[1.07, 100+]
Dominican Republic	[0.24, 0.85]	[0.24, 0.53]	[0.55, 100+]	[0.94, 12.30]
Ecuador	[0.55, 100+]	[1.71, ∞]	[0.36, ∞]	[0.99, 100+]
Egypt	[0.80, 3.18]	[0.23, 0.64]	[0.63, ∞]	[0.23, 12.98]
El Salvador	[0.17, ∞]	[0.62, 2.15]	[0.39, 3.46]	[0.40, 100+]
India	[0.42, 1.24]	[0.33, 0.58]	[0.99, 9.93]	[0.61, 4.07]
Israel	[0.50, ∞]	[0.46, 2.40]	[1.41, 100+]	[0.18, 1.78]
Korea	[0.44, 1.09]	[0.63, 2.14]	[0.59, 6.37]	[0.49, 3.62]
Malaysia	[0.66, 2.54]	[0.47, 100+]	[1.07, 8.45]	[0.67, 12.18]
Mexico	[1.47, ∞]	[1.47, ∞]	[1.02, 53.96]	[0.95, 100+]
Morocco	[100+, ∞]	[0.22, 3.67]	[100+, ∞]	[0.63, ∞]
Pakistan	[1.46, ∞]	[0.24, 6.88]	[100+, ∞]	[0.18, 12.09]
Paraguay	[7.21, ∞]	[1.64, ∞]	[2.87, ∞]	[6.93, ∞]
Philippines	[0.77, 3.04]	[11.13, ∞]	[0.57, 12.58]	[0.46, 3.95]
South Africa	[0.23, 86.31]	[0.27, 0.42]	[0.78, 100+]	[0.31, 23.29]
Sri Lanka	[1.72, ∞]	[0.87, ∞]	[0.23, 100+]	[1.20, ∞]
Thailand	[0.41, 0.98]	[0.48, 1.06]	[0.69, 6.70]	[0.35, 32.61]
Uruguay	[0.24, 0.50]	[0.41, 1.26]	[1.23, 21.48]	[2.81, ∞]
Venezuela	[0.88, ∞]	[0.35, 41.25]	[0.48, 100+]	[1.91, ∞]

All half-life estimates are expressed in years. For half-life estimates that are longer than 100 years, they are indicated by "100+" in the table. [$LB_{95\ percent}$, $UB_{95\ percent}$] indicates the lower and upper bounds of the 95 percent confidence interval for the corresponding half-life estimate.

countries (Cheung and Lai, 2000; Murray and Papell, 2002), these upper-bound half-life estimates vary widely among developing countries, ranging from less than a year to infinitely many years.

A closer examination of our results suggests a possible difference in pattern between the BW and the post-BW period. Fig. 2 summarizes the results on half-life uncertainty by looking at the relative width of half-life confidence intervals for each historical period. The interval width, which is the difference between $UB_{95\ percent}$ and $LB_{95\ percent}$, shows the level of uncertainty in estimating the half-life. A wider (narrower) confidence interval indicates greater (lesser) uncertainty. Compared to those for the BW period, the half-life confidence intervals for the post-BW period are more likely to be wider than to be narrower. This observed pattern holds for both real official and parallel exchange rates. For the post-BW data, in only one out of the 24 cases can the width of the half-life confidence interval be shorter than 3 years. For the BW data, in contrast, the width of the confidence interval can be less than 3 years in at least 10 cases.

The foregoing results raise interesting questions about the source of the increased uncertainty in measuring real exchange rate adjustment speeds. In our study, the post-BW sample period yields a larger number of observations than the BW sample period does. Although the width of the confidence interval depends partly on the sample size, the difference in sample size cannot

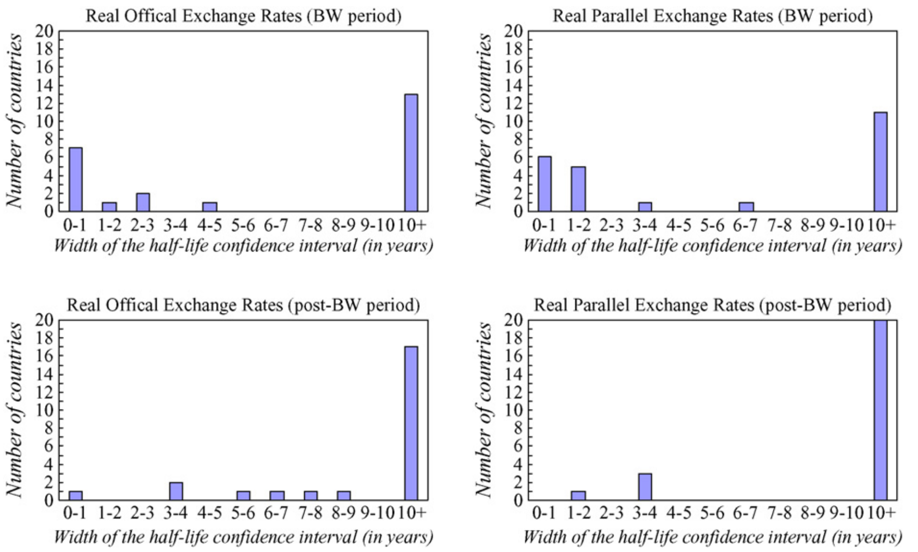

Fig. 2. Cross-country variation in the uncertainty in half-life estimation.

account for the observed difference in the interval width between the two historical periods. For a given confidence level, the width of the confidence interval decreases, not increases, as the sample size grows. All else being equal, we should actually expect to see lower measurement uncertainty for the post-BW data than for the BW data. The empirical evidence shows just the opposite. Accordingly, the increased uncertainty may reflect the difference in variability of shocks rather than the difference in sample size. This suggests that some not yet identified factors can be responsible for the increased uncertainty in measuring real exchange rate adjustment speeds during the post-BW period. Bayoumi and Eichengreen (1994), for example, observe that real shocks have become more dominant during the post-BW period. The source of the increased sampling variability poses an interesting issue for future research.

5. Further analysis of the cross-country variation in adjustment speeds

With the estimated half-life of the real exchange rate varying greatly from country to country, it is instructive to investigate how much the observed cross-country variation is ascribable to inter-country differences in structural economic characteristics. We will examine several key country characteristics that have been identified in the literature as possible factors influencing real exchange rate dynamics. The first structural characteristic is the country's openness to trade. A basic element of the PPP adjustment process is that goods market arbitrage affects trade flows and induces real exchange rate adjustment. To ascertain whether openness to trade facilitates real exchange rate adjustment, the degree of trade openness is measured by the ratio of total trade (imports plus exports) to the country's GDP.

The second structural characteristic under study is the country's productivity growth. This supply-side factor is the focus of the Balassa–Samuelson hypothesis, which posits that productivity growth can induce sustained changes in real exchange rates due to its differential impact on the prices of traded and non-traded goods (Balassa, 1964; Samuelson, 1964). Empirical evidence for the hypothesis has been presented by, e.g., Canzoneri et al. (1999) and Chinn and Johnston (1999) based on OECD data.

Money growth is another country characteristic to be explored. A country's money growth rate is an indicator for the general stance of monetary policy: a lower (higher) rate of money growth means a tighter (easier) monetary policy. Unlike productivity changes, which are a source of real shocks, money supply changes represent monetary shocks. If nominal exchange rate and price changes are dominated by monetary rather than real shocks, real exchange rates are expected to adjust relatively fast. It is thus interesting to see if there is any significant negative relationship between half-lives of real exchange rates and money growth rates across countries.

The fourth country characteristic under consideration is government spending, a demand-side factor included in some structural models of real exchange rates (e.g., Frenkel and Razin, 1996; Froot and Rogoff, 1991; Obstfeld and Rogoff, 1996). Unlike private spending, government spending tends to fall more heavily on non-traded goods. Consequently, government spending can affect the relative demand for – and thus the relative price of – traded and non-traded goods. Balvers and Bergstrand (2002) also highlight the complementarity of private and government consumption as an important channel through which government spending can influence the equilibrium real exchange rate. In addition to short-run effects, Alesina and Perotti (1995) show that government spending, if financed by distortionary taxes, can have long-run real effects. De Gregorio et al. (1994) examine panel data from OECD countries and report empirical evidence in support of a positive relationship between government spending and the relative price of traded and non-traded goods. Chinn and Johnston (1999) also find significant evidence from OECD data that government spending affects real exchange rates.

Table 7 shows the differences among the 24 developing countries in terms of structural economic characteristics. The trade openness variable represents the average ratio of total trade to GDP over the relevant sample period. For the productivity growth variable, the average annual rate of growth in real per capita GDP over the sample period is used as a broad proxy. The money growth variable measures the average annual rate of M2 money growth (adjusted for real GDP growth) over the sample period. The G/GDP growth variable gives the average annual rate of growth in government spending as a share of GDP over the sample period. In general, the cross-country data display huge variability in structural economic characteristics among the developing countries under study.

An inspection of the cross-country variation in either adjustment speeds or structural characteristics suggests that the usual distributional assumption on normality is not tenable for these data. This calls for the use of non-parametric methods in our statistical analysis. We first employ Spearman's rank correlation analysis to gauge the strength (and direction) of the relationship between the half-life of the real exchange rate and each structural economic variable. The rank correlation method makes no assumptions about the data distribution, and it does not require the underlying relationship between the variables to be linear.[5] The rank-based method also works well when the data are not given in precise sample values.

To implement the Spearman analysis, all the observations are ranked from the smallest to the largest for each data series. In case when a tie in rank occurs, the observations involved are assigned the average value of the ranks they would receive as if they were in successive order. Considering a pair of variables, say (x_1, x_2), the rank correlation statistic (denoted by ρ) is given

[5] The Pearson product–moment correlation coefficient, which is the standard correlation statistic, measures how well a linear equation describes the relation between two variables. The Spearman rank correlation coefficient, in contrast, measures how well an arbitrary monotonic function can describe the relationship between two variables. The monotonic function can be nonlinear.

Table 7
Cross-country differences in structural economic characteristics

Country	Bretton Woods				Post-Bretton Woods			
	Trade openness	Productivity growth	Money growth	G/GDP growth	Trade openness	Productivity growth	Money growth	G/GDP growth
Argentina	22.05	14.93	18.22	−10.59	17.32	0.16	183.55	−1.21
Bolivia	43.58	3.25	22.45	1.70	47.70	0.13	94.72	1.35
Brazil	13.57	6.03	43.28	−1.33	17.65	1.52	266.31	3.01
Chile	74.31	13.86	42.38	3.22	52.73	2.95	46.45	−1.30
Colombia	27.37	3.94	19.56	3.47	31.22	1.69	34.29	3.57
Costa Rica	54.74	3.16	7.93	2.90	73.87	1.24	25.53	−0.41
Dominican Republic	39.62	2.72	14.66	−2.81	64.02	1.97	17.29	−0.88
Ecuador	35.54	2.95	10.50	−0.97	52.53	0.92	10.23	−0.09
Egypt	37.84	N.A.	N.A.	3.27	53.13	2.75	14.75	−3.45
El Salvador	51.78	2.32	7.66	0.06	57.62	0.41	10.95	−0.45
India	10.60	1.46	9.80	1.99	16.60	2.72	14.84	1.54
Israel	65.06	4.61	0.71	4.56	91.06	2.10	61.91	−1.36
Korea	26.68	4.51	42.07	−0.69	68.59	5.76	15.04	0.59
Malaysia	87.13	7.01	11.50	1.96	130.66	3.97	9.95	−1.87
Mexico	20.59	3.32	9.46	4.43	34.16	1.21	38.72	0.54
Morocco	42.52	1.22	9.72	0.13	50.04	1.52	14.67	1.70
Pakistan	20.62	1.12	14.37	2.45	33.81	2.37	12.71	0.57
Paraguay	30.80	1.89	24.79	0.17	54.74	1.53	22.47	1.70
Philippines	30.23	1.78	10.91	1.51	59.56	0.62	19.25	1.66
South Africa	52.29	2.19	6.55	2.15	50.56	−0.40	11.81	1.74
Sri Lanka	54.02	3.96	7.14	−0.45	70.60	3.76	17.70	−0.70
Thailand	37.48	4.24	15.02	1.43	63.38	4.85	12.44	0.70
Uruguay	26.35	−0.24	35.45	1.51	39.70	2.18	63.65	−0.74
Venezuela	48.52	2.05	7.75	0.25	53.25	−0.37	26.22	−1.70

All the above numbers are given in percent per year, and they represent average values over the corresponding sample period.

by

$$\rho = 1 - \frac{6\sum_{j=1}^{N}[R(x_{1j}) - R(x_{2j})]}{N(N^2 - 1)} \tag{8}$$

where N is the number of observations for each variable, $R(x_{1j})$ the rank order of the jth observation of variable x_1, and $R(x_{2j})$ the rank order of the jth observation of variable x_2. The null hypothesis of $\rho = 0$ (i.e., of no correlation between the two variables) can be tested against the alternative hypothesis of $\rho \neq 0$. Based on the rank correlation estimates (not reported but available upon request), little correlation could be found in all but one case. In the exception case, a significantly negative correlation was found between the money growth rate and the half-life of the real official exchange rate. In all the other cases, the correlation estimates were either statistically insignificant or having ambiguous signs or both.

A more formal analysis of the cross-country data is performed using the multiple rank regression method. Similar to rank correlation analysis, rank regression analysis is based on the rank-ordered data instead of the original data. The regression equation includes the various economic factors as explanatory variables:

$$\mathrm{HL}_j = \zeta_0 + \zeta_1\mathrm{OPEN}_j + \zeta_2\mathrm{PROD}_j + \zeta_3\mathrm{GOVT}_j + \zeta_4\mathrm{MONEY}_j, \quad j = 1, 2, \ldots, N \tag{9}$$

where HL is the half-life of the real exchange rate of the corresponding country, OPEN the trade openness of the respective economy, PROD the country's average productivity growth rate, GOVT the average increase in government spending as a share of the country's GDP, and MONEY the average money growth rate (adjusted for real GDP growth) in the respective country.

Table 8 contains the results from rank regression analysis, and they are largely consistent with those from rank correlation analysis. The coefficients on the variables of trade openness, productivity growth, and government spending are all found to be statistically insignificant. Some of these coefficients may even have an ambiguous sign. For the money growth variable, the coefficient is statistically significant in the case involving the real official exchange rate. The negative coefficient in this case suggests that countries with higher money growth rates tend to have faster speeds of real exchange rate adjustment.[6] However, this finding does not apply to the real parallel exchange rate, which yields a money growth coefficient of an incorrect sign, though it is statistically insignificant. In addition to the separate regressions conducted for the BW and the post-BW period, regressions were also run using pooled data from both BW and post-BW periods together, with a time-period dummy being included as well. Qualitatively similar results were obtained from the pooled data that only the money growth coefficient was found to be statistically significant.

In summary, the estimates of real exchange rate adjustment speeds show substantial variation across developing countries. These different speed estimates are generally found to differ from Rogoff's (1996) range of consensus estimates identified for industrial countries. Although the cross-country variation can be partially attributed to inter-country differences in money growth, much of the heterogeneity remains unaccounted for. This is certainly an area that warrants additional investigation in future research.[7]

[6] The negative relationship found between the adjustment speed and the rate of money growth is consistent with previously reported findings that PPP holds particularly well for economies with high inflation rates (e.g., McNown and Wallace, 1989).

[7] A new study by Benigno (2004) illustrates how a country's choice of monetary policy rules may influence real exchange rate persistence.

Table 8
Multiple rank regression analysis of the cross-country differences in adjustment speeds

	Bretton Woods		Post-Bretton Woods	
	Real official rate	Real parallel rate	Real official rate	Real parallel rate
Trade openness	−0.077 (0.218)	0.148 (0.261)	−0.347 (0.257)	−0.089 (0.263)
Productivity growth	−0.007 (0.203)	0.255 (0.243)	0.237 (0.229)	−0.038 (0.235)
Government spending	0.011 (0.188)	−0.032 (0.225)	−0.075 (0.241)	0.038 (0.247)
Money growth	−0.666 (0.216)*	0.187 (0.259)	−0.125 (0.231)	0.217 (0.237)
R^2	0.411	0.144	0.118	0.079

The dependent variable is the adjustment speed measured in terms of the half-life of the real exchange rate. Standard errors of the coefficient estimates are given in parentheses. Statistical significance is indicated by an asterisk (*) for the 5 percent level.

6. Concluding remarks

A longstanding issue in exchange rate economics concerns whether greater nominal exchange rate flexibility promotes real exchange rate adjustment—an important channel through which an open economy adjusts to disturbances. Some economists consider nominal exchange rate flexibility to be important for macroeconomic adjustment by accelerating the realignment of the real exchange rate, while others hold the opposing view that free exchange rate movement may actually disrupt and prolong the real exchange rate adjustment process because speculative forces can send the nominal exchange off its equilibrating path. These contrasting views reflect the old Nurkse versus Friedman debate about exchange rate flexibility and the effect of speculation.

In this study, we have analyzed whether the more flexible parallel market rate generates a faster speed of real exchange rate adjustment than the less flexible official rate does. We do observe substantially greater variability in parallel market rate changes than in official rate changes. Nonetheless, there is no significant evidence that greater nominal rate flexibility tends to yield faster real rate adjustment. Nor is there any significant evidence that greater nominal rate flexibility tends to produce slower real rate adjustment. In other words, based on the information from dual exchange rate systems, no systematic relationship can be found between nominal rate flexibility and the speed of real rate adjustment. The result holds for both the BW and the post-BW data.

Many developing countries have used parallel exchange markets as a tool to adjust to economic shocks and external imbalances. When official exchange rate adjustment is limited, changes in real exchange rates need to come mainly through price changes. In these countries where an official peg to the dollar is adopted, the presence of an active parallel exchange market offers a back-door channel that may help facilitate real exchange rate adjustment. Although parallel exchange rates can move much more freely than official rates, parallel rate movements are, in most cases, not found to generate faster real exchange rate adjustment.

The foregoing finding on relative adjustment speeds raises questions about the working of the parallel exchange market. A possible view may be that the parallel exchange rate tends to maintain a certain gap from the official exchange rate without bringing about a faster adjustment speed. The rate difference reflects, at equilibrium, the market price for restrictions associated with the official rate (the authors owe this point to an anonymous referee). Another possible view may be that strong speculative forces are at work in many parallel exchange markets, especially for short-term market movements. Currency traders can overreact to market shocks, thereby amplifying the short-term impact of these shocks on exchange rates. As a result, PPP deviations magnify first before diminishing. The rising importance of chartists in currency trading may further reinforce such overreacting behavior. When the market overreacts, it not only adds to the short-term volatility of the exchange rate but also prolongs the time it takes for the real rate to revert toward its long-run level. Accordingly, the parallel exchange market fails to produce significantly faster adjustment than the official exchange market.

A final remark about the scope of our analysis is in order. The central question we examined is: If exchange rates were operating under similar economic conditions, would a more flexible rate generally produce faster or slower real rate adjustment than a less flexible rate? We did not attempt to make any general inferences about the difference in the speed of real exchange adjustment between economies with fixed exchange rates and those with floating rates. Countries with different exchange rate systems usually differ considerably in economic conditions, which can cause different real exchange rate dynamics. Even among economies having the same exchange rate regime, there can be substantial variation in real exchange rate adjustment behavior. The developing countries examined in this study are characterized as having largely a fixed exchange rate under the official IMF classification, and yet these countries still yield a wide range of speed estimates for real exchange rate adjustment. Moreover, the *de facto* exchange rate regime can be very different from the officially stated regime. Reinhart and Rogoff (2004) observe that, when a parallel exchange market exists, a regime of an official peg might easily turn out to be a *de facto* float or a crawling band. All these empirical issues make a general determination of the regime effect rather difficult. This brings us back to the reason why we do not analyze exchange rate flexibility in broader terms of exchange rate regimes (floating-rate regimes as opposed to fixed-rate regimes), but in more specific terms of dual exchange rates (market-determined parallel rates as opposed to pegged official rates). The latter approach provides a more controlled study of the adjustment dynamics of exchange rates with different flexibility. Our empirical evidence suggests that the market-determined parallel rate does not often adjust faster than the government-set official rate.

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Appendix A. Parallel market premium and purchasing power parity

The parallel market premium may also reflect part of the deviation from purchasing power parity (PPP). To illustrate, we consider a simple variant of the parallel exchange rate models discussed by Kouretas and Zarangas (1998) and Diamandis (2003), under which both financial and goods arbitrages take place in the parallel foreign exchange market. All the model variables are in logarithms. Financial arbitrageurs exploit any divergence between the official and the parallel market rate, and their net supply of foreign exchange in the parallel market is described by

$$S_{\mathrm{Pt}} = \theta_0 + \theta_1(e_{\mathrm{Pt}} - e_{\mathrm{Ot}}), \quad \theta_1 > 0 \tag{A.1}$$

where e_{Pt} is the parallel market rate and e_{Ot} is the official rate (all expressed as domestic currency per U.S. dollar). This indicates that as the parallel market premium increases, so does the profit incentive to meet demand for foreign exchange in the parallel market. Goods arbitrages, on the other hand, are carried out based on the differential between the foreign price (p_{Ft}) and the domestic price (p_{Dt}). Since demand for foreign exchange rises as foreign goods become cheaper relative to domestic goods, the net demand for foreign exchange from goods arbitrageurs in the parallel market is specified as

$$D_{\mathrm{Pt}} = \zeta_0 + \zeta_1(p_{\mathrm{Dt}} - e_{\mathrm{Pt}} - p_{\mathrm{Ft}}), \quad \zeta_1 > 0 \tag{A.2}$$

where $p_{\mathrm{Dt}} - e_{\mathrm{Pt}}$ gives the domestic price in foreign currency units (i.e., in U.S. dollars). At market equilibrium (i.e., when $S_{\mathrm{Pt}} = D_{\mathrm{Pt}}$), Eqs. (A.1) and (A.2) combine to yield the following condition:

$$e_{\mathrm{Pt}} - e_{\mathrm{Ot}} = \varphi_0 + \varphi_1(\bar{e}_t - e_{\mathrm{Ot}}) \tag{A.3}$$

where $\varphi_0 = (\zeta_0 - \theta_0)/(\theta_1 + \zeta_1)$, $\varphi_1 = \zeta_1/(\theta_1 + \zeta_1)$, and $\bar{e}_t = p_{\mathrm{Dt}} - p_{\mathrm{Ft}}$ is the PPP-implied equilibrium rate. To the extent that PPP prevails in the long run, the parallel market premium, $e_{\mathrm{Pt}} - e_{\mathrm{Ot}}$, is stationary, implying the existence of a long-run equilibrium relation between e_{Pt} and e_{Ot}.

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