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EXCHANGE RATES DYNAMICS OF THE REFORMING VIETNAM: AN EVIDENCE OF LONG-RUN EQUILIBRIUM

(Continued)

VUONG QUAN HOANG

4. EMPIRICAL FINDINGS

4.1. Non–stationary univariate time series for all RER, NER, and CPI in levels while all change series exhibit stationarity.

Univariate time series in financial markets are in many cases exhibiting non–stationary characters. In numerous papers, researchers have found evidence of non–stationarity for exchange rates (log) in levels, as well as consumer prices measured as index. Table [1] below presents results of our unit roots tests on univariate time series, in levels and first differences, before moving on to work with the models in focus. The series that are put to test span over the sample of 1986:01 to 2002:12; that is the entire reform period of Vietnam.

Table [1] – Unit root tests for RER, nominal and prices series Period 1986:01–2002:12

RER	ADF	PP	RER	ADF	PP
US	–2.281563	–2.186644	ΔUS	–8.504968*	–11.05621*
UK	–2.239055	–2.144665	ΔUK	–8.690370*	–10.88917*
JP	–2.364021	–2.207391	ΔJP	–8.631092*	–10.94557*
EU	–2.107712	–2.092097	ΔEU	–7.328626*	–10.04807*
Nominal ER	ADF	PP	Nominal ER	ADF	PP
US	–2.113743	–1.783928	ΔUS	–7.119171*	–11.38085*
UK	–2.218590	–1.892893	ΔUK	–7.040271*	–11.16233*
JP	–2.236723	–1.903101	ΔJP	–7.156835*	–11.24721*
EU	–2.624890	–2.301729	ΔEU	–5.904018*	–10.24907*
CPI	ADF	PP	CPI	ADF	PP
US	–0.660043	–0.031474	ΔUS	–9.081898*	–10.23048*
UK	–1.076937	–0.754681	ΔUK	–4.328766*	–11.86841*
JP	–1.630198	–2.162647	ΔJP	–10.86389*	–12.57107*
EU	–0.007153	–0.300107	ΔEU	–9.629268*	–13.53560*
VN	0.885414	1.730725	ΔVN	–2.833471*	–3.575707*

*(**) Null hypothesis of unit root rejected at 1(5)% level. ADF regression specification is based on minimizing AIC/SC. Rates for RER, nominal ER and CP index are provided in natural logarithms.
RER: ADF critical values for US. UK, JP at 1(5)% level: –4.0070(–3.4333); for EU: –4.0179 (–3.4385); PP critical values for US, UK, JP at 1(5)% level: –4.0061(–3.4329); for EU: –4.0158(–3.4376).
Nominal ER: ADF critical values for US, UK, JP at 1(5)% level: –4.0070(–3.4333); for EU: –3.4727 (2.8798); PP critical values for US, UK, JP at 1(5)% level: –4.0061(–3.4329); for EU: –3.4708(–2.8789).
CPI: ADF critical values for US, UK, JP, EU at 1(5)% level: –4.0070(–3.4333); PP critical values for US, UK, JP at 1(5)% level: –4.0061(–3.4329); for EU: –4.0061(–3.4329).

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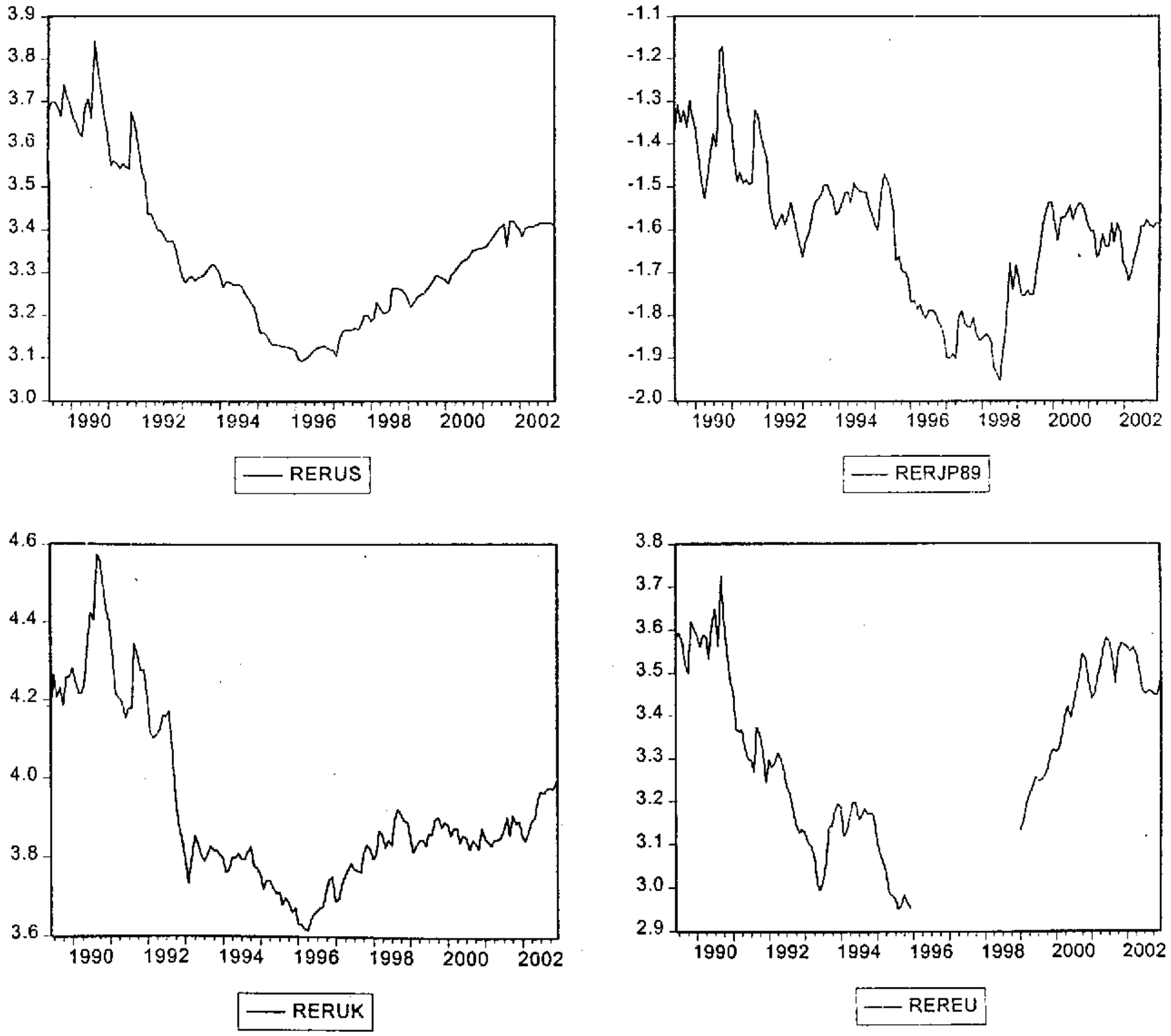
These tables indicate that exchange rates (both real and nominal log) and consumer prices in levels are non–stationary, given the optimal lag truncation, according minimizing AIC and SC. Both ADF and PP test statistics cannot reject the null of unit root existence in univariate series in levels.

We will be taking on a later starting date for our time series of RER for the reason that Vietnam officially recognized other hard currencies than the old Soviet rubble since June–1989. For this, trade and commerce in new hard currencies become more active, and the exchange rates reflect better commercial nature of the exchange rates dynamics. It is also important that during this recent period of the financial market evolution, the difference between "black market" and official market exchange rates is quite negligible so that our data of official market is relevant to reflect the exchange rates in Vietnam's general economic settings. The figure [8] below shows the changes of RER of four key currencies of our study over the period 1989:06 to 2002:12. One could hardly imagine stationarity in RER variables.

Figure [8] – Real effective exchange rates since June–1989 in Vietnam



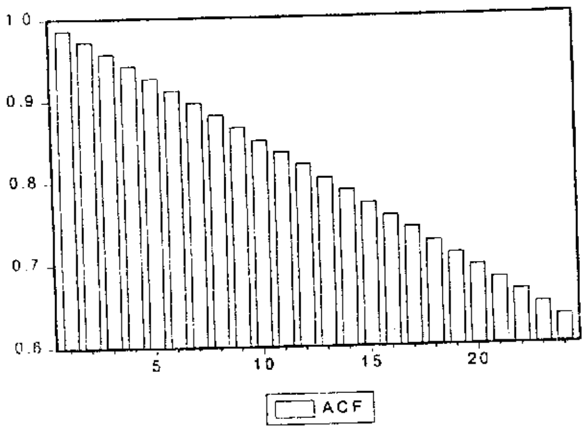
The RER series, although moving in different ways over the considered period, still indicate somewhat similar behavior, observing peaks and valleys. Netting effect of prices logarithms has partially wiped out divergent behaviors of nominal exchange rates of these currencies.

Table [2] - Unit root tests on RER period 1989:06–2002:12

RER	κ	ADF	AIC(SC)	PP	RER	κ	ADF	AIC(SC)	PP
US	3	–1.96419	–4.2784 (–4.1835)	–2.5078	ΔUS	2	–10.909*	–4.2665 (–4.1906)	–13.1370*
UK	5	–1.5626	–3.6146 (–3.4817)	–1.9221	ΔUK	7	–6.4224*	–3.6076 (–3.4368)	–11.2432*
JP	5	–2.0335	–3.4123 (–3.2794)	–2.5368	ΔJP	4	–8.1425*	–3.3984 (–3.2845)	–11.2815*
EU	1	–1.0516	–3.5418 (–3.4740)	–1.1022	ΔEU	1	–9.3634*	–3.5445 (–3.4763)	–10.9964*

*(**) Null hypothesis of unit root rejected at 1(5)% level. ADF regression specification is based on lag truncation (K) that minimizes AIC/SC. Rates for RER in natural logarithms.
RER levels: ADF critical values for US, UK, JP at 1(5)% level: –3.4715(–2.8792); for EU: –3.4835(–2.8845);
PP critical values for US, UK, JP at 1(5)% level: –3.4715(–2.8792); for EU: –3.4831(–2.8844);
RER first–order difference: ADF critical values for US, UK, JP at 1(5)% level: –3.4715(–2.8792); for EU: –3.4839(–2.8847); PP critical values for US, UK, JP at 1(5)% level: –3.4715(–2.8792); for EU: –3.4835(–2.8845);

Again, this sub–sample unit root test cannot reject our null hypothesis of non–stationarity in all exchange rates series in levels, even at 10%. It is because all ADF and PP test statistic values fall well onto the acceptance region of the empirical distribution. However, test statistics decisively reject unit root hypothesis in all time series of first–order differences, at 1% level. We can also observe the correlogram of RER and CPI in level, say, for US, below, which also exhibit quintessential pattern of non–stationarity, where ACF does not die out after quite a number of lags. For instance, at lag 24, the autocorrelation coefficient is still significant, with rather large (positive) magnitude.



In brief, we recognize that all series appear to have followed a first–order integrated, i.e. $I(1)$. At this point, we are ready for further examination of cointegration and other relationship in equilibrium.

4.2. Co-integrated relations among e_t p_t *and* p_t^*

This part of our test in effect examines weaker condition for PPP theorem to hold. The Engle–Granger (1987) work indicates that although non–stationary, these types of economic variable can be constructed into a linearly combined system that is stationary. Depending upon the specific test outcome, such a stationary linear combination can verify the extent to which long–run equilibrium will be reaching, given short–run data set. Outcomes derived from both Engle–Granger and Johansen's co–integration tests follow.

Co-integration under Engle-Granger two-stage analysis:

Engle–Granger (1987) method proposes a framework for bivariate relationship that presents a long–run equilibrium, in which case our consideration of exchange rates can be implemented. The trivariate system among

s_t p_t and p_t^* will be grouped into the following:

$$s_t = \alpha_1 + \alpha_2 \left(p_t - p_t^* \right) + u_t$$

$$s_t = \alpha_1 + \alpha_2 d_{pp^*} + u_t$$

The last equation indicates that nominal exchange rates are not fixed (for instance in government's price peg policy); that is, allowed to fluctuate. A visual check for possibly interrelated s_t and d_{pp^*} is provided in figure [13] of the appendix. This situation is true for VND during the sub–sample period 1989:06–2002:12. Results of regression based on the last equation are summarized in the table [3] as follows.

Table [3] – E–G first–stage bivariate OLS regression

N.E.R.	Estimation			
s_{USD}	$s_t =$	$+1.289198$	$+1.331015 d_{pp^*}$	$+EOT_{USD}$
		(0.139690)	(0.025163)	
		[9.235575]	[52.89629]	
s_{GBP}	$s_t =$	$+1.737678$	$+1.369444 d_{pp^*}$	$+EOT_{GBP}$
		(0.144531)	(0.026522)	
		[12.02289]	[51.26697]	
s_{JPY}	$s_t =$	-3.769239	$+1.341055 d_{pp^*}$	$+EOT_{JPY}$
		(0.132990)	(0.023326)	
		[-28.34235]	[57.49279]	
s_{EUR}	$s_t =$	$+1.376404$	$+1.313861 d_{pp^*}$	$+EOT_{EUR}$
		(0.169018)	(0.029640)	
		[8.649363]	[44.32721]	
Note:	Standard error (s.e.) is reported in parentheses; t–Stat. in square brackets.			

We note that the test statistic for error–correction term $\hat{u}_{s,t}$ does not follow standard distribution as in usual ADF and PP test. Thus, our statistical inference has to revert to Phillips–Ouliaris (1990) critical values for m variables (Appendix-Table [13]; in our bivariate model, $m = 2$).

Table [4] - Engle–Granger second–stage bivariate regression using ADF paradigm (ADF test statistic comparable to Phillips–Ouliaris simulated critical values)

	Equation				ADF	k	Min. AIC(SC)
EOT_{USD}	$\Delta\hat{a}_t =$	-0.005741	$-0.109596\hat{a}_{t-1}$	$-0.300557\Delta\hat{a}_{t-1}$	–4.346062*	1	–0.5598
		(0.01279)	(0.02522)	(0.06527)			(–0.5106)
		[–0.44895]	[–4.34606]	[4.60488]			
EOT_{GBP}	$\Delta\hat{a}_t =$	-0.005326	$-0.10648\hat{a}_t$	$+0.31642\Delta\hat{a}_{t-1}$	–4.275919*	1	–0.5181
		(0.01306)	(0.0249)	(0.0653)			(–0.4689)
		[-0.4079]	[-4.276]	[4.8462]			
EOT_{JPY}	$\Delta\hat{a}_t =$	-0.00566	$-0.11915\hat{a}_t$	$+0.3165\Delta\hat{a}_{t-1}$	–4.531211*	1	–0.5403
		(0.0129)	(0.0263)	(0.0653)			(–0.4912)
		[–0.4309]	[-4.5312]	[4.8602]			
EOT_{EUR}	$\Delta\hat{a}_t =$	-0.008302	$-0.112168\hat{a}_t$	$+0.28699\Delta\hat{a}_{t-1}$	–4.03133*	1	–0.3337
		(0.015876)	(0.027824)	(0.072184)			(–0.2770)
		[–0.52291]	[–4.03183]	[3.97581]			

*(**) Null hypothesis of non–stationarity rejected at 1(5)% level; comparable to Phillips–Ouliaris c.v., for m=2. Tests are conducted for period 1986:01–2002:12. Standard error (s.e.) is reported in parentheses; t–Stat. in square brackets. $\hat{a}_t$ denotes error–correction term from the first–stage E–G regression.

We have obtained empirical results from implementing E–G framework that indicate co–integration between relative price difference and spot exchange rates for all data series in question. Next, estimation of trivariate systems is provided.

Table [5] - Engle–Granger first–stage trivariate OLS regression

N.E.R.	Estimation				
s_{USD}	$s_t =$	+2.290992	$-1.577618 p_{US}$	$+1.360767 p_{VN}$	$+ECT_{USD}$
		(2.061967)	(0.606630)	(0.047758)	
		[1.111071]	[–3.112521]	[28.28360]	
s_{GBP}	$s_t =$	+2.414354	$-1.530858 p_{UK}$	$+1.376679 p_{VN}$	$+ECT_{GBP}$
		(1.917186)	(0.484999)	(0.055475)	
		[1.259322]	[–3.156417]	[24.81628]	
s_{JPY}	$s_t =$	–2.149773	$-1.716036 p_{JP}$	$+1.356548 p_{VN}$	$+ECT_{JPY}$
		(3.093860)	(0.716037)	(0.037690)	
		[–0.694861]	[–2.396405]	[35.99266]	
s_{EUR}	$s_t =$	–2.847658	$-0.275682 p_{EU}$	$+1.228819 p_{VN}$	$+ECT_{EUR}$
		(2.234669)	(0.643776)	(0.053665)	
		[–1.274314]	[–0.502357]	[22.89778]	

The system separates domestic and foreign prices to reflect the relative PPP relation more clearly. (Standard errors in this situation are biased; reported for completeness.)

Table [6] - Engle–Granger second–stage trivariate regression using ADF paradigm (ADF test statistic comparable to Phillips–Ouliaris simulated critical values)

	Equation				ADF	K	Min. AIC(SC)
ECT_{USD}	$\Delta\hat{e}_t =$	–0.006797	$-0.111129\hat{e}_t$	$+0.302170\Delta\hat{e}_{t-1}$	–4.398784*	1	–0.559690
		(0.012789)	(0.025263)	(0.066126)			(–0.510558)
		[–0.453320]	[–4.398784]	[4.639797]			
ECT_{GBP}	$\Delta\hat{e}_t =$	–0.006375	$-0.107563\hat{e}_t$	$+0.317886\Delta\hat{e}_{t-1}$	–	1	–0.517462
		(0.013060)	(0.024927)	(0.066182)	4.315829**		(–0.468329)
		[–0.411571]	[–4.315829]	[4.876926]			
ECT_{JPY}	$\Delta\hat{e}_t =$	–0.006582	$-0.120845\hat{e}_t$	$+0.318421\Delta\hat{e}_{t-1}$	–4.585588*	1	–0.539348
		(0.012917)	(0.026353)	(0.066118)			(–0.490215)
		[–0.432126]	[–4.585588]	[4.889888]			
ECT_{EUR}	$\Delta\hat{e}_t =$	–0.007917	$-0.108769\hat{e}_t$	$+0.281820\Delta\hat{e}_{t-1}$	–	1	–0.336426
		(0.015853)	(0.027971)	(0.072708)	3.888662**		(–0.279722)
		[–0.499418]	[–3.888662]	[3.876071]			

*(**) Null hypothesis of non–stationarity rejected at 1(5)% level; comparable to Phillips–Ouliaris c.v., for m=3. Tests are conducted for period 1986:01–2002:12. Standard error (s.e.) is reported in parentheses; t–Stat. in square brackets. $\hat{e}_t$ denotes error–correction term from the first–stage E–G regression.

For both original Engle–Granger bivariate and extended trivariate regressions, we mostly cannot accept the null hypothesis of non–stationarity nature of the error–correction terms derived from empirical data at 1% level; except only log spot Euro and British pound at conventional 5%. The optimal lag truncation in the two models for ADF tests on ECT is 1, minimizing both AIC/SC. Therefore, our empirics on the co–

integration of the spot exchange rates and relative prices difference, and separate domestic–foreign prices, under E–G(1987), has verified the long–run equilibrium between the variables in the system. In reality, our short–run data although cannot confirm stationarity of RER levels, does unveil the behavior that endogenous variables of the system are interrelated and interact with one another in the long run to shape a stationarity over a long time span. This proves a relatively weak version of PPP theorem. It is also not uncommon that this two–stage co–integrating test may help alleviate the low power problem of ADF and PP unit root test frameworks.

Error-correction mechanism in modified ADF regression:

The problem of low power in testing for non-stationarity is well known. We have in the previous conduct of Engle-Granger (1987) two-stage regression found an evidence that NER, prices and relative price ratio will tend to error-correct towards some future equilibrium. Now we implement a modified ADF test to indicate some empirical values of such tendency.

The major motivation for this modified paradigm is described in the following. In our original theory, strong-form PPP implies (3.1). An augmented Dickey-Fuller regression (for first-order autoregression) is:

$$\Delta RER_t = \alpha + \delta t + \gamma RER_{(t-1)} + \sum_k \gamma_i \Delta RER_{(t-i)} + v_t$$

δt should not be present in the test for consistency of PPP theory. Now we construct a slightly different analysis to better reflect the error-correction mechanism. Clearly, the new construction reads:

$$\Delta s_t = \alpha + \gamma RER_{(t-1)} + \phi \Delta dpp*_{(t)} + \sum_k \left(\psi_{si} \Delta s_{(t-i)} + \psi_{dppi} \Delta dpp*_{(t-i)} \right) + \sum_k \delta_i \Delta dpp*_{(t+i)} + v_t$$

where s_t is spot exchange rate (in logarithms); RER real exchange rate; and $dpp*$ relative price ratio. Lead term in relative price ratio has been added as well as its lagged values up to order k. Our empirical results are in the table [7], examining the model for $i = 1$ in all lead and lag terms.

Table [7] – A modified ADF analysis

$$\Delta s_t = \alpha + \gamma RER_{(t-1)} + \phi \Delta dpp*_t + \sum_1 \left(\psi_{si} \Delta s_{(t-i)} + \psi_{dppi} \Delta dpp*_{(t-i)} \right) + \sum_1 \delta_i \Delta dpp*_{(t+i)} + \varepsilon_t$$

	α	γ	ϕ	ψ_{s1}	ψ_{dpp*1}	δ_1	Q_6	Q_{12}
USD	0.406902* (5.1242)	–0.117053* (–4.9949)	–0.366671 (–0.4947)	0.214487* (3.2832)	0.061878 (0.1011)	–0.503665 (–0.8327)	1.74 (0.94)	21.12 (0.05)
GBP	0.452088* (4.9967)	–0.111406* (–4.8762)	–0.222998 (–0.2978)	0.234402* (3.5820)	–0.050719 (–0.0826)	–0.532069 (–0.8791)	3.12 (0.79)	21.96 (0.04)
JPY	–0.184559* (–4.5086)	–0.125864* (–5.2671)	–0.302001 (–0.4223)	0.231126* (3.5562)	–0.305011 (–0.5122)	–0.297616 (–0.5109)	1.09 (0.98)	17.92 (0.12)
EUR	0.439262* (4.8921)	–0.126337* (–4.7529)	–0.468250 (–0.5644)	0.185142* (2.5538)	0.079089 (0.1157)	–0.573860 (–0.8530)	1.09 (0.98)	15.59 (0.21)

(*) denotes significance at 1% level. t–Stat. in parentheses. Q_6 and Q_{12} denote Ljung-Box statistic for serial correlation of up to 6 and 12 lags.

The results unveil a pattern of the exchange rate dynamics that all γ_i are significant at 1% level and negative. The regression is satisfactory with residuals exhibit white-noise property; no serial correlation afterwards. In terms of absolute magnitude, they appear to have been close to one another. Absolute values of ϕ in all cases are well below the unity. ϕ and φ_0 bear opposite signs, that is, negative in ϕ and positive with φ_0 (correctly signed). For all tested currencies, their magnitudes, however, do not seem to equal, except the case of GBP.

The new dynamics for all exchange rates indicate that coefficient of first-order lagged RER in the regression is statistically significant and smaller than zero. The result is important in verifying the above E-G analysis because the negative coefficient in the long run becomes endogenous error-correction engine. Large deviation from PPP in the preceding term will contribute to reduce the divergence in the next term, a sufficient condition for the exchange rate to tend to some equilibrium in the future.

Co-integration test under Johansen's VAR paradigm:

We now report some empirical from max eigenvalue approach of VAR(2) analysis for searching for co-integrated equations amongst variables versus stationarity character of the data series used for the system. We repeat the previous note that the entire reform period of Vietnam can be seen as consisting of two major sub-periods, which are before 1989:06 and after. The main reasons of this is due to the liberalizing of the trade account since beginning of 1989, where international trading started becoming a more liberal field of activities and private forces could join the market efficiently. Therefore, we can test for PPP separately between the two sub-samples for a general understanding. This is equivalent to testing for a dummy of structural break.

Table [8] - Results from Johansen test for co-integrating equations. Sample period: 1989:06-2002:12

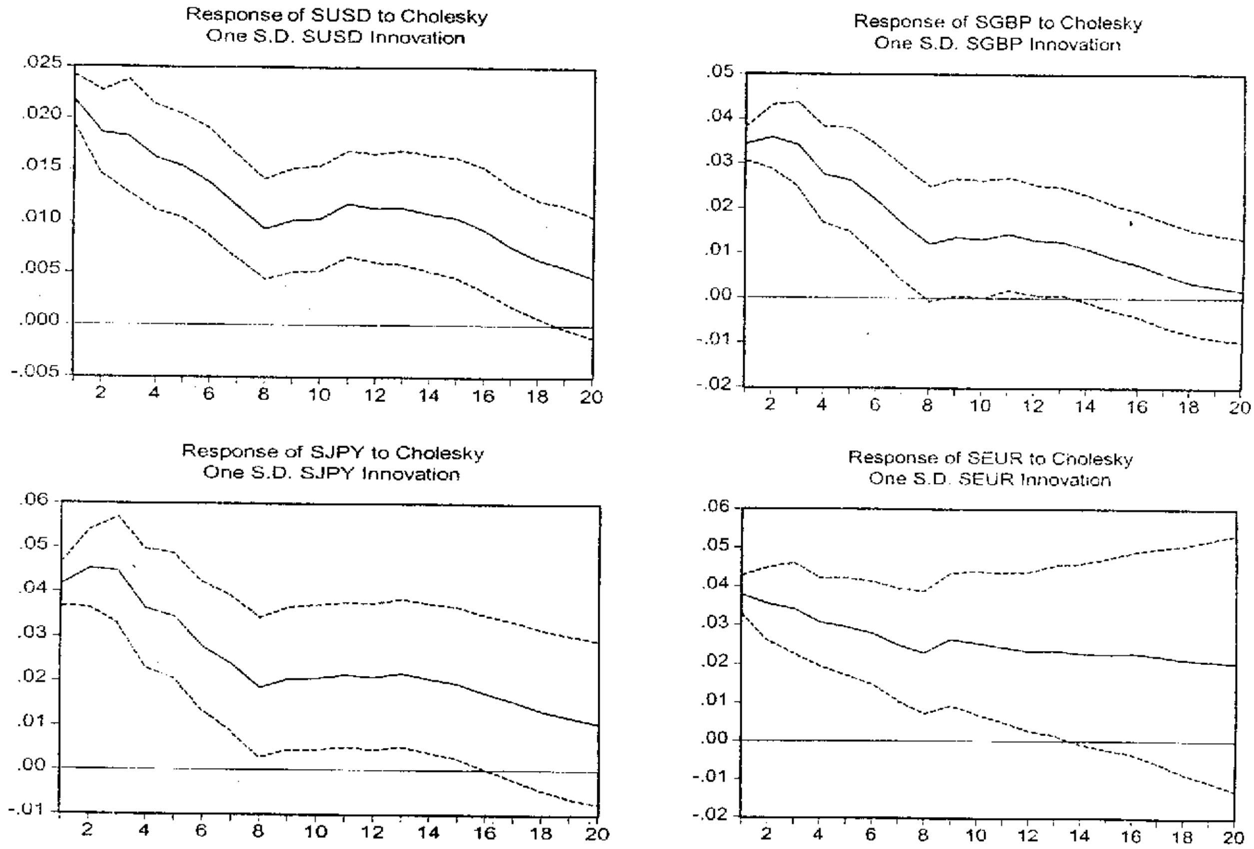
	Null	λ_{TR}	95% CV	No. of CE	Max-Eig.Val.	95% CV
USD	$r \leq 3$	11.83	12.25	2 (5%)	32.58	25.54
	$r \leq 1$	33.10**	25.32	1 (1%)	21.26*	18.96
	$r = 0$	65.68**	42.44		32.58**	25.54
	Co-integrating coefficient			Adjustment coefficient		
	USD	PUS	PVN	D(USD)	-0.089116 (0.01772)	
	1.000000	-3.055586	0.195094	D(PUS)	-0.000846 (0.00152)	
		(0.53449)	(0.13651)	D(PVN)	-0.007270 (0.00922)	
GBP	$r \leq 2$	2.22	3.76	2	2.22	3.76
	$r \leq 1$	24.69**	15.41	(5% &1%)	22.48**	14.07
	$r = 0$	52.03**	29.68		27.34**	20.97
	Co-integrating coefficient			Adjustment coefficient		
	GBP	PUK	PVN	D(GBP)	-0.004026 (0.00768)	
	1.000000	-0.938357	-2.030323	D(PUK)	0.003234 (0.00074)	
		(2.08206)	(0.59709)	D(PVN)	0.004407 (0.00204)	
JPY	$r \leq 2$	0.035031	3.76	2	0.035031	3.76
	$r \leq 1$	21.32**	15.41	(5% &1%)	21.28**	14.07
	$r = 0$	58.26**	29 68		36.94**	20.97
	Co-integrating coefficient			Adjustment coefficient		
	JPY	PJP	PVN	D(JPY)	-0.086534 (0.02621)	
	1.000000	-2.053206	-0.211476	D(PJP)	-0.008660 (0.00261)	
		(0.65684)	(0.10101)	D(PVN)	-0.019847 (0.00719)	
EUR	$r \leq 2$	0.84	3.76	1	0.84	3.76
	$r \leq 1$	13.07*	15.41	(5% &1%)	12.23*	14.07
	$r = 0$	61.77**	29.68		48.7**	20.97
	Co-integrating coefficient			Adjustment coefficient		
	EUR	PEU	PVN	D(EUR)	0.003298 (0.01428)	
	1.000000	-8.444299	-0.279437	D(PEU)	0.006368 (0.00099)	
		(1.36698)	(0 26734)	D(PVN)	0.007823 (0.00515)	

*(**) denotes null hypothesis rejected at 1(5)% level. Co-integrating coefficient (beta) and adjustment coefficient (alpha) vectors are presented for one co-integration equation. Standard errors reported in parentheses. D(□) denotes first-order difference of data in logarithms. VAR models selected for lag length of 10, although information criteria (LR, AIC and SC) indicate different lag choices. Johansen cointegration test is done with constant, but no trend in CE.

The test outcome supports the hypothesis that trivariate systems are co-integrated with reduced rank. In 3/4 exchange rates in the test, at 5% level, we accept the hypothesis that there exist two co-integrating relations among the three variables. In the case of EUR, the co-integration equation is one (at both 5 and 1% level). Normalized co-integrating vector and adjustment coefficient matrix are provided in table [8].

The co-integration equations reveal that these considered variables tend to combine well with leads in value to become a stationary linear combination, therefore supporting the mean-reverting nature of real exchange rates. The figure [9] below show partial picture of impulse response of the spot exchange rates (in level) to one standard deviation shock. Given a future period of 20 months, the IRF (Monte Carlo simulation using EViews random number generator; 10000 replications) shows that given other shocks held constant, effects of the spot exchange rates innovation alone on the future real exchange rates reduce over time, and fairly quickly. Similar observation can be seen with individual shocks of domestic and foreign inflation levels.

Figure [9] - IRF of spot exchange rates to unit innovation

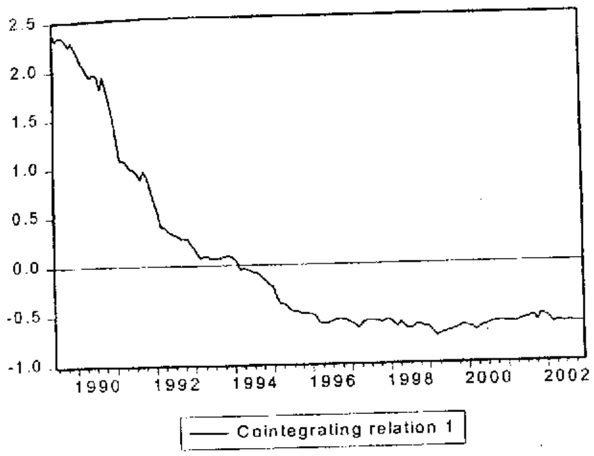


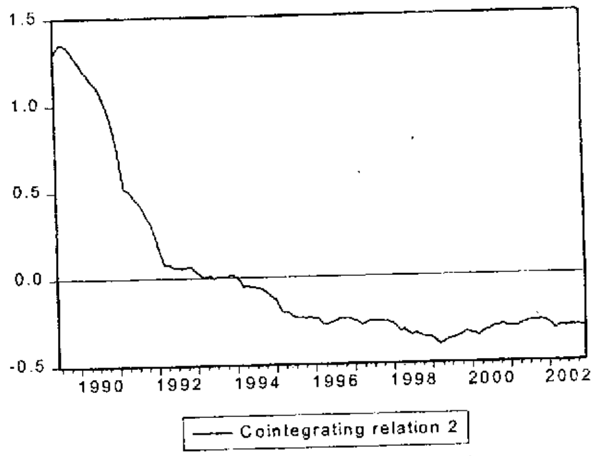
VECM representation:

Co-integrating between endogenous variables of RER implies VECM, as vector autoregressive multiple equations can be presented in VEC specifications; or a system of autoregressive distributed lags, with first difference of lagged variables introduced into the model. Below are representations of vector error correction regression; using Johansen's multivariate analysis. For brevity, we show only the first estimation, applicable in the case of flexible (movable) exchange rates, in which changes in spot exchange rates on the LHS, while ECT followed by lagged changes in values of endogenous variables. Lag structure is identified based on AIC-minimizing information, and in our situation, 3 for all data series.

USD:

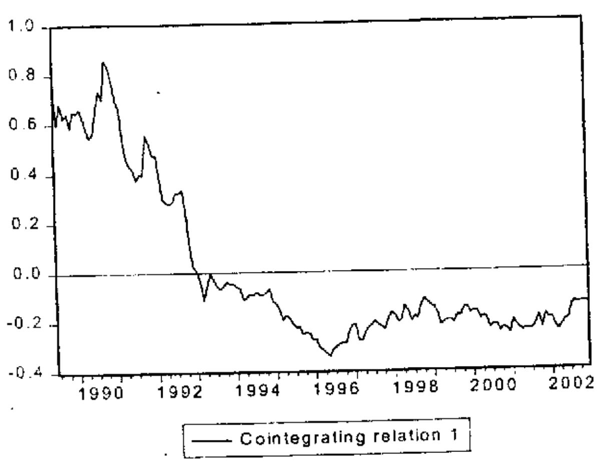
$$\Delta s_{USD} = -0.0419(s(-1) - 2.297 p_{VN}(-1) + 15.75) + 0.087(p_{US}(-1) - 1.13 p_{VN}(-1) + 7.41)$$
$$-0.076\Delta s_{USD}(-1) - 0.0014\Delta s(-2) - 0.104\Delta s(-3)$$
$$+1.634\Delta p_{US}(-1) - 0.675\Delta p_{US}(-2) + 0.776\Delta p_{US}(-3) + 0.445\Delta p_{VN}(-1)$$
$$+ 0.0283\Delta p_{VN}(-2) - 0.08462\Delta p_{VN}(-3) + 7.4254 \times 10^{-5}$$

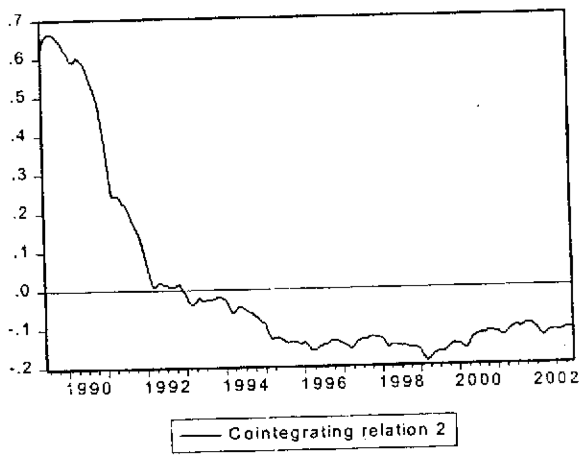




GBP:

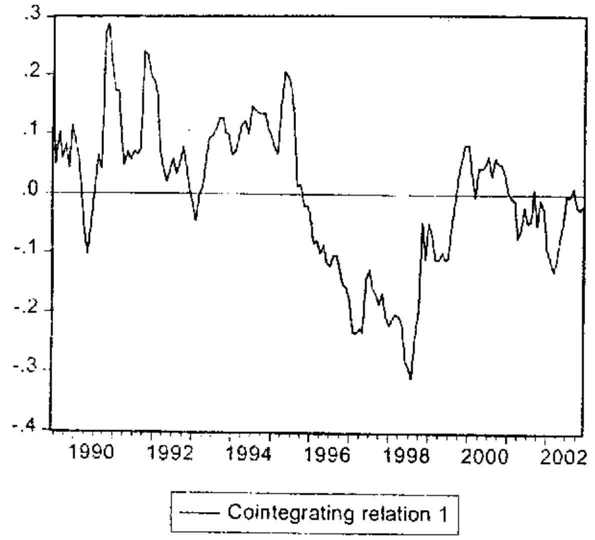
$$\Delta s_{GBP} = -0.0663(s(-1) - 1.06 p_{VN}(-1) + 1.78) + 0.098(p_{UK}(-1) - 0.671 p_{VN}(-1) + 2.273)$$
$$+0.067\Delta s(-1) + 0.06\Delta s(-2) - 0.11\Delta s(-3) + 1.085\Delta p_{UK}(-1)$$
$$+0.074\Delta p_{UK}(-2) + 1.66\Delta p_{UK}(-3) + 0.365\Delta p_{VN}(-1) + 0.096\Delta p_{VN}(-2)$$
$$-0.115\Delta p_{VN}(-3) - 0.0048$$

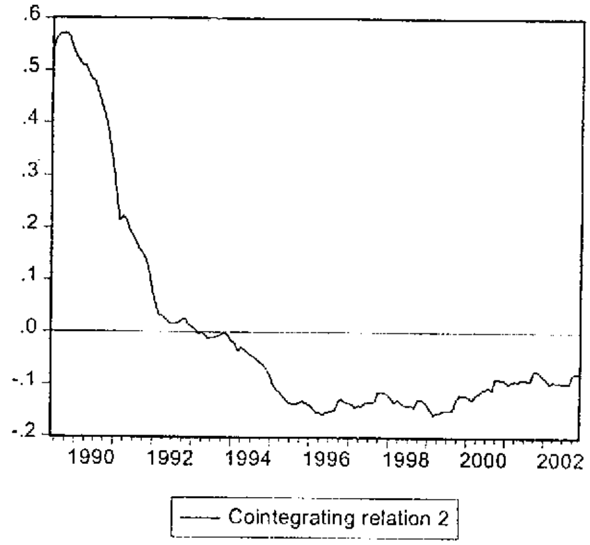




JPY:

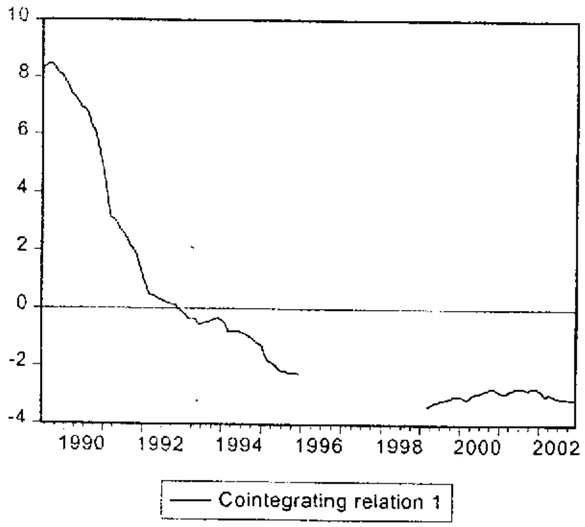
$$\Delta s_{JPY} = -0.077(s_{JPY}(-1) - 0.777 p_{VN}(-1) + 3.937) + 0.037(p_{JP}(-1) - 0.496 p_{VN}(-1) + 0.648)$$
$$+0.1467\Delta s_{JPY}(-1) + 0.66788\Delta p_{JP}(-1) + 0.30\Delta p_{VN}(-1) + 0.0019$$

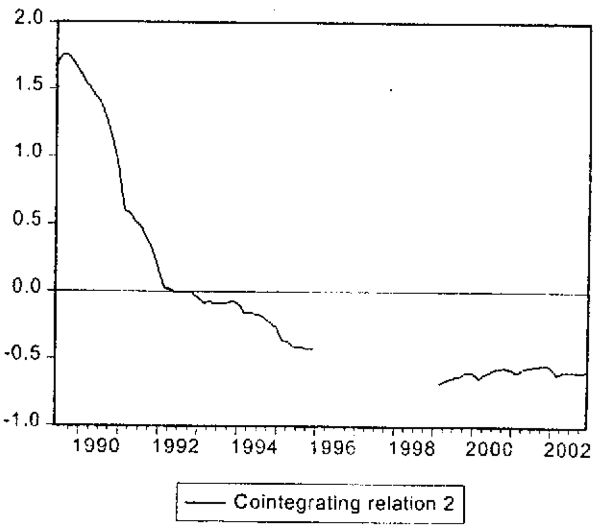




EUR:

$$\Delta s_{EUR} = -0.063(s_{EUR}(-1) - 7.203 p_{VN}(-1) + 68.766) + 0.308(p_{EU}(-1) - 1.634 p_{VN}(-1) + 11.66)$$
$$+0.0057\Delta s_{EUR}(-1) + 0.729\Delta p_{EU}(-1) + 0.629\Delta p_{VN}(-1) - 0.0023$$





In the estimating process, lag length of through 3 is chosen for USD and GBP, while 1 for JPY and EUR based on minimal AIC. However, our foremost concern is with the decomposed matrices α and β of $\Pi = \alpha\beta'$ (as reported in table [8], in the preceding section).

Restricted coefficients long-run equilibrium:

We now report further statistics that impose standard restriction for ECT of the long-run relationship. It is known that the coefficient vector of ECT in the PPP theorem must satisfy:

$$\beta' y_t = (s_t + p_t - p_t^*) \sim I(0)$$

Therefore, the restriction for absolute PPP (strong-form) becomes: $\beta' = (1,1,-1)$.

Below are test statistics for these restricted coefficients under Johansen VAR procedure:

	rank	lag	LogL	LR	d.f.
USD	1	10	1739.04	8.46 (0.015)	1
	2	10	1752.39	1.74 (0.187)	2
GBP	1	10	1596.14	1.36 (0.505)	2
	2	10	1607.70	0.71 (0.399)	1
JPY	1	10	1507.70	9.012 (0.011)	1
	2	10	1521.17	3.364 (0.067)	2
EUR	1	10	1120.46	18.93 (0.000)	1

The above results indicate that although we support the long-run equilibrium in all these series, the strict form of PPP where vector $\beta' = (1,1,-1)$ is defined theoretically by PPP theorem is rejected for JPY and EUR, while accepted for USD and GBP exchange rates.

4.3. Half-lives and convergence to long-run equilibria

We have found evidence that the exchange rates will be convergent in the future to some equilibrium. This expected equilibrium value will possibly be reached when the error-correction component tends to reduce the preceding deviation from PPP to a negligible magnitude. More exactly, the real exchange rate will come to approximate equilibrium point after some time. Such a process of reverting to an unbiased mean value (mean-reversion property of PPP) should indicate the speed at which deviation will be adjusted by the endogenous error-correction process.

In the table below, we summarize some values of half-lives (H-L) for different real exchange rates under a widely used D-F regression. The D-F framework does not suggest a time trend for the regression, however, due to a possible time trend control (Balassa-Samuelson effect), the inclusion of time is also conducted for comparison.

Table [9] - Half-lives under D-F regressions

	Trend (no first difference) $\Delta RER_t = \alpha + \delta t + \gamma_i RER_{t-1} + \varepsilon_t$			No-trend (no difference) $\Delta RER_t = \alpha + \gamma_i RER_{t-1} + \varepsilon_t$			No-trend (with first difference) $\Delta RER_t = \alpha + \gamma_i RER_{t-1} + \psi_i \Delta RER_{t-1} + \varepsilon_t$		
	γ_i	$(1+\gamma_i)$	H-L	γ_i	$(1+\gamma_i)$	H-L	γ_i	$(1+\gamma_i)$	H-L
USD	−0.0496	0.950368	13.62	−0.0355	0.96455	19.20	−0.0450	0.95498	15.05
GBP	−0.0447	0.955275	15.15	−0.0339	0.96606	20.07	−0.0439	0.956131	15.45
JPY	−0.0478	0.952243	14.16	−0.0355	0.96450	19.18	−0.0456	0.954362	14.84
EUR	−0.0658	0.934158	10.18	−0.0392	0.96084	17.35	−0.0477	0.952281	14.18

In our computation, half-life is defined as the necessary lag length for the current deviation from PPP to decay by 50% of its original magnitude. It is quite straightforward to prove that the value of H-L depends on the point estimates of γ_i, in which:

$$\text{H-L} = \ln(1/2)/\ln(1+\gamma_i)$$

Where, shocks of the regression decay monotonically (as has been verified in previous examination of autocorrelation function).

The test results show that for all three cases summarized in the table, half-lives of exchange rates in consideration, range from 0.85 to 1.67 years. Specifically, the case of

DF regression with no first-order difference, the shortest time is 1.45 years (EUR) and longest 1.67 years (GBP). This speed of mean reversion is quite fast compared to those obtained by previous studies, where H-L is typically ranging from 3 to 5 years. Our estimated half-lives clearly indicate that their mean values are not only finite, but also fairly small in magnitude, compared to many previously reported results. For example, Murray and Papell (2002) reported H-L of exchange rates denominated in US Dollar in a range like 0.89 (FFR), to 11.2 years (JPY), after bootstrapping an annual data sample of period 1900-1996. Most of values vary around 2.0-4.0 years.

4.4. Beyond the analytics

Bootstrapping H-L:

Our analysis above is based on a sample period of 1986:01-2002:12. For many previous studies on long-run equilibrium of real exchange rates, data samples usually span over decades, or even centuries, in which researchers find evidence supporting PPP validity. In this study, the data sample is limited due largely to a short period of reform (since 1986). It will then be worthwhile that we conduct a replication of many more samples based on the actual one to observe any significant changes in our previous note of finite H-L for different currency rates.

The replication is done with a bootstrapping algorithm executed on the *Mathematica* environment (V4.0), where we construct populations from different replications, specifically 750, 1500, 5000 and 15000. Because bootstrap replication does not require known distribution of the stochastic variable, different percentile values of simulated mean γ_i are also reported. Thus, such as discussed in Murray and Papell (cf. 2001); a bootstrap for bounds of point estimates of lagged variable coefficient can help obtain asymptotically valid result.

Table [10] - Results from bootstrapping γ_i and corresponding H-L

		750	1500	5000	15000	Observed
JPY	γ_i	-0.0332448	-0.0341238	-0.0346953	-0.033385	**-0.0354967**
	0.5%	-0.185298	-0.220035	-0.2486	-0.363691	
	2.5%	-0.182793	-0.190434	-0.218846	-0.225442	
	5%	-0.159074	-0.172666	-0.188707	-0.185383	
	95%	0.0315635	0.027538	0.0275885	0.0277923	
	H-L	**20.50**	**19.96**	**19.63**	**20.41**	**19.18**
USD	γ_i	-0.0341519	-0.0329634	-0.0343204	-0.0335253	**-0.0354534**
	0.5%	-0.221611	-0.253531	-0.260319	-0.31165	
	2.5%	-0.171767	-0.226552	-0.238483	-0.227406	
	5%	-0.142112	-0.182305	-0.194928	-0.18639	
	95%	0.0440411	0.0260389	0.0286264	0.0297169	
	H-L	**19.95**	**20.68**	**19.85**	**20.33**	**19.20**
GBP	γ_i	-0.035126	-0.0331505	-0.0328467	-0.0328393	**-0.0339426**
	0.5%	-0.194006	-0.233934	-0.290493	-0.302865	
	2.5%	-0.190955	-0.202818	-0.213739	-0.216312	
	5%	-0.160449	-0.183399	-0.182572	-0.180472	
	95%	0.0320275	0.0271793	0.0285752	0.0282177	
	H-L	**19.38**	**20.56**	**20.75**	**20.76**	**20.07**
EUR	γ_i	-0.0350281	-0.035505	-0.0373075	-0.037321	**-0.0391623**
	0.5%	-0.237784	-0.241543	-0.295558	-0.340962	
	2.5%	-0.237784	-0.238357	-0.240514	-0.225166	
	5%	-0.198972	-0.189881	-0.19127	-0.192136	
	95%	0.0281187	0.0284251	0.0311187	0.0302162	
	H-L	**19.44**	**19.17**	**18.23**	**18.22**	**17.35**

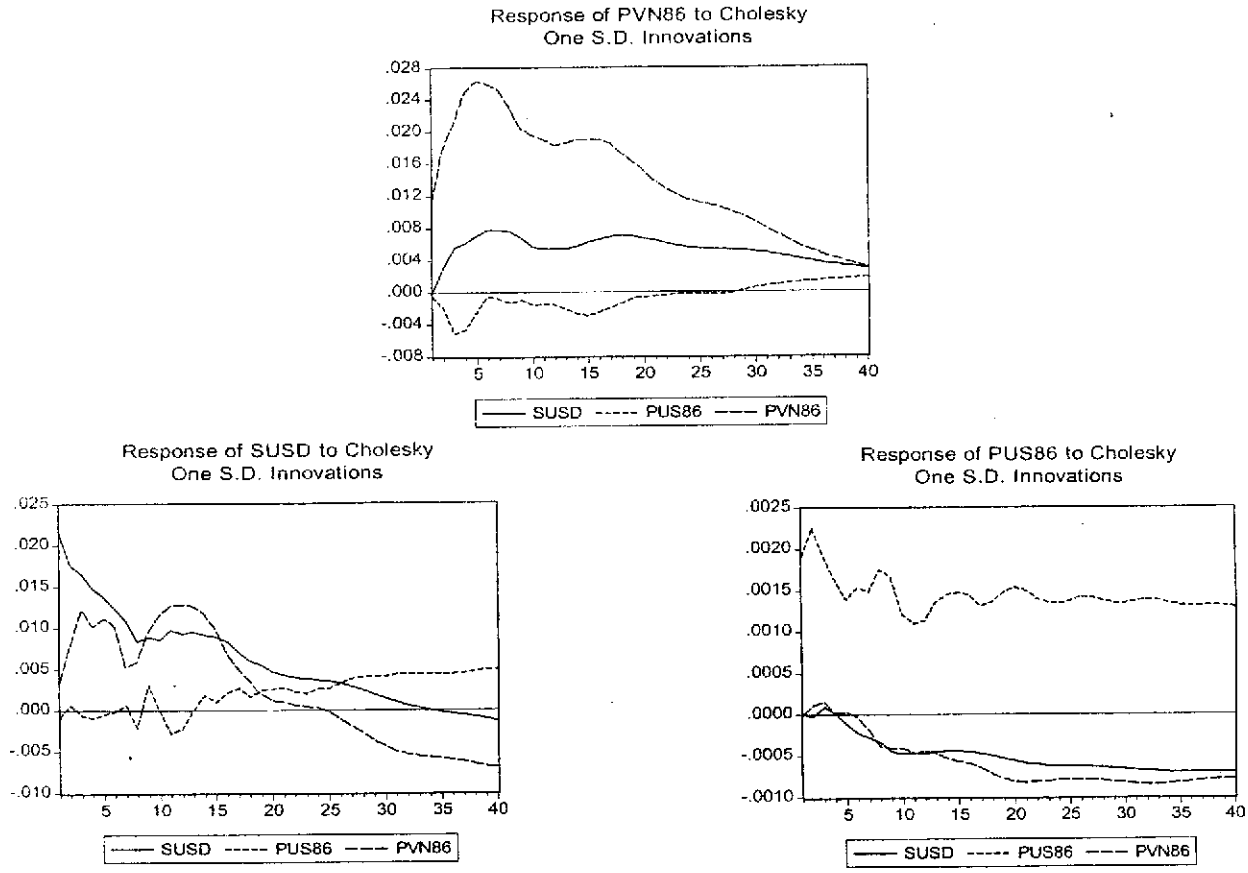
Percentile of bootstrap mean provided in the second column, above H-L.

We note that for most cases, bootstrapped mean of half-lives is dragged on. The exception is for GBP 750-replication sample where H-L is little smaller than the observed value in D-F regression. The change, although small, in computed H-L is because absolute values of most autoregressive parameters get a bit smaller than the originally observed. It appears that the finiteness of H-L is confirmed while increase in H-L after the treatment does not change the picture substantially. Frequency distributions of the autoregressive parameters after being bootstrapped are provided in the Appendix, where we see clearly that γ_i are not normally distributed but skewed to the right, containing more mass. The speed of correction towards PPP is obviously quite slow, however, it could still show fast pace of adjustment if compared to other world results.

Impulse responses of VAR system:

We now pay more attention to the persistence of innovations to changes in economic variables over some period ahead. This should be done on discrete data with a Monte Carlo simulation, beyond the analytics. Simulated statistics will help observe the pattern of relative changes between variables in the system, provided that others remain unchanged; or in short, we analyze relationship between separate pair of variables with observed shocks.

Figure [10]



The difficulty of tracing effects of a random shock (impulse) of one equation on other endogenous variables of the VAR system is well known, because the error terms of equations can and in most cases of financial time series system are correlated. These correlated innovations may represent a dynamic system through which a one-time shock affects current and future values of one or more endogenous variables.

Beyond the analytics, which is hard to obtain given the trivariate VAR system, for the impulse-response relation, and a widely used approach is to simulate the impact of innovation whose value is equal to one standard deviation of the corresponding error term. Tracing the effect requires different decomposition methods, and the above graph shows Cholesky decomposition (in which the Cholesky ordering matters, and we follow the rule: domestic prices à foreign price à spot exchange rate).[3]

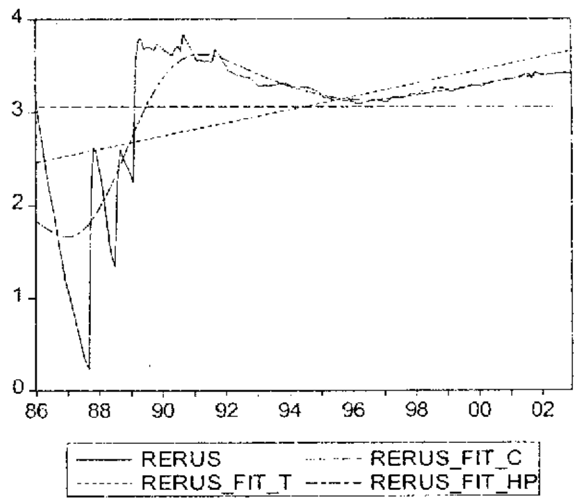
Recent RER misalignment?

In line with our analysis of the long-run equilibrium, where we establish that it is very likely that real exchange rate dynamics will bind stochastic variables to a stationary relationship with finite speed of convergence, it is possible to explore behavioral process of real exchange rates over the considered sample. It is critical for this exercise as through 17 years of the economy, several important events may represent shocks that could drive real exchange rates from deterministic theoretic equilibrium levels. Economists and policy makers should also be wary of significant departure (in magnitude) from considered equilibrium.

In this aspect, we examine three widely used benchmarks for detecting a departure from hypothesized equilibrium real exchange rates (ERER). The first is to regress RER data against a constant (i.e., a fixed ERER), the second on a constant but allowing for a time trend component (a linear time trend ERER), and the third by Hodrik-Prescott filtered data (fluctuating ERER).

Our full empirical analysis of RER misalignment is provided in table [14] of the Appendix, and below are several key observations. In fact, given the same original data of RER, the three fitting will provide for different understanding. It is not unexpected that while this shows an undervaluation of USD in parity with VND, the other version may unveil an overvaluation. We now detect the consensus in two ways. First, if any two of the three show an overvaluation, we take the point with some caution. If all three agree on over- versus undervaluation, we accept the conclusion. RER_USD shows that USD was undervalued against VND significantly in early years of reform. The average undervaluation of USD is about 60% in 1987; 20% 1988. However, in 1989, the situation changed and USD became more or less overvalued by 15%; over 20% in 1990; 15% 1991; <15% 1992; and less than 10% for the period 1993-95. Period 1996-2002 shows negligible disparity towards undervaluation of USD, with undervaluation is estimated below 5% (average level of the year).

Figure [11] - Misalignment from USD ERER: 1988-2002



The average undervaluation of GBP is about 50% in 1987; 15-20% 1988. Similar to USD, in 1989, GBP became more or less overvalued by 10-15%; over 20-30% in 1990; 15-20% 1991; above 15% 1992; and about 1-5% for the period 1993-99. Period 2000-01 shows small undervaluation of GBP, with

3 This is because we can reasonably assume that domestic price change will have a first and foremost impact on the exchange rate dynamics. (See Hamilton (2000) for further analysis)

undervaluation is estimated 1-5%. Behavior of EUR is also quite relevant to GBP, with significant undervaluation in 1987 (50-60%) and 1988 (30%). One difference is undervaluation of EUR in 1995, when USD and GBP were overvalued in our computation. However, the magnitude is quite small, less than 3%. In the year of new EUR launching, 1999, EUR again became undervalued by 3%, and the situation reversed in 2000, where EUR gained over VND by about 5%. As the confidence and reputation of EUR got stronger, EUR was overvalued in 2001 by substantial percentage of over 10%. In 2002, however, the undervaluation returned by negligible 1-5%.

The Japanese yen behavior looked somewhat different from other strong currencies in this study. Period 1987-88 saw a strong yen, with magnitude of overvaluation in between of 40-60%. Yen in 1989-95 depreciated quickly and the parity undervaluation was in the range of 15-30%. It is noted that this period is considered a competitive-yen policy while VND reacts more to USD and EUR than JPY. In period 1996-98, yen regained the value, and became a bit overvalued by about 5-10% against VND, together with move of FDI and trade increase between the two nations. 1999-2000 saw the trend reversed where a negligible depreciation in RER of JPY. In 2001-02, JPY bounced back and the overvaluation was about less than 10%.

(Continued)