Do Prices Determine Exchange Rate?
The Japanese Evidence for Purchasing Power Parity

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Abstract

The longtime perplexing purchasing power parity (PPP) puzzle has been recently resolved empirically by using the pure price inflation rates extracted and estimated by a pioneering financial-asset pricing approach. Applying the same extracted inflation rates, we estimate a vector error-correction (VEC) model of prices and the Japanese yen per U.S. dollar exchange rate, and find strong evidence supportive of (i) the PPP restriction which yields *the equilibrium error in the form of a real exchange rate.* Further, documented under the PPP relationship so detected are (ii) the impulse responses *of exchange rate to prices* and *between prices* that would imply exchange rates channeling inflations into countries, and (iii) the impulse responses *of prices to exchange rate* (i.e., *exchange rate effects on prices*) that would usefully indicate the degree of exchange rate pass-through by Japanese exporters. Together, the findings lend, in the VEC context, desired PPP-theoretic content to the pure inflation rate estimates used.

*The present paper is a revised version of Kojima (2006a), a research conducted while I was Visiting Scholar at The John E. Anderson Graduate School of Management at UCLA during the 2005-2006 academic year. Two working papers were written based on Kojima (2006a) while I was at UCLA Anderson: Kojima (2006b, 2006c), which are, respectively, *Finance Working Paper Nos. 11-06* and *12-06*, downloadable at the UCLA Anderson Website *http://www.anderson.ucla.edu/x5962.xml* or at the Social Science Research Network (SSRN) Website *http://ssrn.com/author=649009.* Kojima (2006b) in particular is a revised, shortened version of Kojima (2006a) and hence of the present paper. I am grateful to my faculty sponsor Richard Roll for constructive comments and suggestions, and for helping download from the American Economic Association Website the data and program files written for Chowdhry, Roll and Xia (2005), without which the present research would not have been initiated. All remaining errors are solely my own.

1 Introduction

Lying at the heart of the purchasing power parity (PPP) puzzle are such inconsistencies between the PPP hypothesis and the empirical findings as the empirical gap in the speeds of adjustment between official price indices and exchange rates: official price indices such as the CPI and WPI (macroeconomic variables) tend to move slowly, whereas an exchange rate (a financial asset) moves much faster. The puzzle was studied in the open macroeconomic framework by Dornbusch (1976) who proposed the now popular (Dornbusch) overshooting model. More recently, the puzzle has been explored, now in the financial-asset context, by Chowdhry, Roll and Xia (2005) (C-R-X) in a novel way that bridges the gap in the speeds of adjustment between prices and exchange rates. Emphasizing that the gap in the speeds of adjustment is due to official price indices moving too slowly and that, to bridge the gap, price indices of *financial-asset* nature should be considered, they extract estimates for realized pure price inflation rate from stock returns. The estimates so extracted turn out to be sufficiently volatile that, using their extracted inflation rates, they document evidence in support of the short-run, relative PPP, thereby resolving the PPP puzzle.

Still another difficult question that remains is whether prices determine exchange rates. Ito (2005, pp.4-5) argues that "Even when the PPP is shown to hold, it is often difficult to determine whether domestic prices adjust to exchange rate changes or the exchange rate is determined by the gap between domestic and foreign prices. That is, proving that the PPP holds does not automatically prove the causality from prices to exchange rates. ... If the causality runs from the exchange rates to the domestic prices, as is feared in currency crisis countries, the estimated PPP relationship is not a theory of exchange rate determination. ... This process occurred in Indonesia, Russia and Argentina." Ito (p.7) adds that "The episode of the hyperinflation resulting in the exchange rate adjustment of a similar magnitude is the best example of the PPP relationship."

Yet, empirically, exchange rate moving fast as a financial variable can *lead* changes in goods prices: exchange rate effects on goods prices are indeed studied extensively in the empirical (and theoretical) literature on exchange rate pass-through. One of recent studies on pass-through is Landon and Smith (2006) who estimate the exchange rate effects on the industry-level investment good prices, for a panel of OECD countries, finding that an exchange rate depreciation [appreciation] leads to a significant rise [fall] in the prices of the investment goods used by most sectors.[1]

The present paper thus focuses on and explores the impulse responses of exchange rate and prices (i.e., impulse responses of exchange rate to prices, those between prices and those of prices to exchange rate) as well as the PPP relationship, and, in so doing, applies C-R-X's pure price inflation rate estimates and a vector error-correction (VEC) framework. Specifically, twofold research objectives are set:

(i) to explore the long-run structure of the yen per dollar exchange rate and C-R-X's extracted price indices, which is a VEC-based cointegration test of the PPP; and

(ii) to conduct the analyses of variance decomposition and impulse response functions in a cointegrated system of the yen per dollar rate and the extracted price indices, which is a study of the short-run structure of the time series, given the long-run structure estimated in (i).[2]

To my knowledge, the paper is the very first attempt to apply C-R-X's extracted inflation rates in investigating these issues for the yen per dollar rate in the VEC model. In the past PPP literature, long-run analyses of real exchange rate have been the main focus of study. For example, a test of the long-run PPP was conducted by Ito (1997) using the real effective yen exchange rate, and the unit-root test was employed to test the long-run constancy for the 1879 through 1995 period, whose results vary depending on the price indices used, the CPI or the WPI.[3] Another important, relevant past work here is Roll (1979) presenting an innovative, efficient markets view of the PPP that the real exchange rate should follow a random walk process. The present paper attempts to contribute to the literature by newly building and estimating a *multi-*

[1] Further, Taylor (2000) suggests, and presents evidence on, a hypothesis that a low inflationary environment results in a low degree of exchange rate pass-through to domestic prices. See also Marston (1990) and Kojima (1995) on two critical patterns of corporate pricing behavior, i.e., exchange rate pass-through and pricing-to-market.

[2] Note that we do not test for the Granger non-causality null hypothesis in the VEC framework; rather, impulse response functions of exchange rate and prices will be computed and studied within the short-run structure. Test for the Granger non-causality in our cointegrated system will be remarked subsequently in section 3.2.3.

[3] For a brief useful survey of existing studies on PPP, see Hausman, Panizza and Rigobon (2006, p.95). For example, the panel data-based unit-root tests for real exchange rates, as a test of the long-run PPP, are conducted by Wu (1996); Kojima (1993) attempts, by a univariate time series analysis, to detect structure changes in the yen per dollar real exchange rate movement that likely induce nonstationarity.

variate system of exchange rate and C-R-X's extracted price indices and interpreting its estimated short- as well as long-run structures in the contexts of PPP and impulse responses. Investigating whether the real exchange rate defined with C-R-X's extracted price indices still obeys a random walk process constitutes another important research topic and this is studied elsewhere, by Kojima (2006c).

To be specific, the paper considers a system of three economic variables (all logged): a nominal exchange rate s_t of a home currency (Japanese yen) against a foreign currency (U.S. dollar), a foreign price index p_t^* and a home price index p_t, where the price indices are constructed from C-R-X's extracted inflation rates. We will proceed by analyzing a series of the following problems:

First (as part of a preliminary analysis), we ask whether the two series $s_t + p_t^*$ and p_t have a strong positive association at all. If they do, does the series $e_t = s_t + p_t^* - p_t$ appear to be a stationary process? (If it does, then the two series $s_t + p_t^*$ and p_t will be cointegrated.) A preliminary analysis of these questions is important in the context of the PPP relationship, since, with the real exchange rate defined as a particular linear combination $r_t \equiv s_t + p_t^* - p_t$, absolute PPP asserts that $s_t + p_t^* = p_t$, that is, $r_t = 0$.[4] Based on the preliminary results obtained, we will next carry out a formal cointegration analysis of the entire three series (s_t, p_t^* and p_t).

Second, are the entire three series (s_t, p_t^* and p_t) cointegrated? In the cointegration analysis of the three series, a particular restriction motivated by our economic arguments will be the (strong) PPP restriction $(1 \quad 1 \quad -1)$, which is in fact a vector of coefficients on the right-hand side of the real exchange rate's definition above. If the PPP restriction is supported by the data, then C-R-X's estimated inflation rates will turn out to have desired PPP-theoretic content, here in the VEC context. This would add to C-R-X's evidence of the same theoretical content documented in their single equation-based analysis of the PPP relationship.

Third, if the three series are cointegrated to have a long-run structure characterized as a PPP relationship, then we next turn to their short-run

[4]Relative PPP requires, in terms of percentage, that $\Delta s_t + \Delta p_t^* = \Delta p_t$ where Δ is the first difference operator. See, for example, MacDonald and Marsh (1994, pp.24-25) and Hausman, Panizza and Rigobon (p.94). C-R-X investigate the PPP puzzle, focusing on the failure of relative PPP and successfully resolve the puzzle in the short run by using their extracted inflation rates as Δp_t^* and Δp_t (C-R-X, pp.260-261).

structure and ask whether the short-run Δs_t equation in the cointegrated system contains any statistically significant short-run effects of Δp^*_{t-l} and Δp_{t-l} with $l > 0$. If it does, the responses of exchange rate to prices will be likely observed. Also, if the short-run price equations contain effects of Δs_{t-l}, Δp^*_{t-l} and Δp_{t-l} with $l > 0$, then exchange rate effects (i.e., the responses of prices to exchange rate) and the responses between the prices are likely detected.

Finally, to further explore responses of exchange rate and prices, the variance decomposition and impulse response functions are computed and studied. Specifically, if the variance of the one-step forecast error for s_t is accounted for by innovations of price indices p^*_t and/or p_t, rather than by own innovations, the responses of exchange rate to prices are likely confirmed. In the impulse response functions analysis, we will interprete the confidence bands as indicating the degree of uncertainty about the shape of impulse responses estimated and, based on the shape of the impulse responses, derive implications both on exchange rate pass-through and exchange rates channeling inflations into countries.

The remaining of the paper is organized as follows: section 2 summarizes the data including C-R-X's inflation rates extracted from stock returns. The cointegration analysis of the yen per dollar rate and the extracted price indices is conducted in a VEC framework in section 3. Given the long-run structure (the PPP relationship) estimated in section 3, an analysis of short-run effects is carried out in section 4, and the impulse responses of exchange rate and prices are further explored in section 5, along with variance decomposition. The final section gives some concluding remarks and a summary of findings. Compiled and tabulated in the data appendix is a complete set of data used in the paper.

2 Data and Inflation Rates Extracted from Stock Returns

The system of monthly exchange rate determination to be investigated in the paper is a system of three economic variables, s_t, p^*_t and p_t (Japanese yen per U.S. dollar nominal rate, U.S. price index and Japanese price index, respectively). The underlying vector autoregressive (VAR) model for our system will be eq. (6) in section 3.1.1. The data period is May 1983 through December 1999. The data sources for the system variables,

s_t, p_t^* and p_t, are here described.

Tables 33 and 34 in the data appendix compile and tabulate all the data the present paper uses for the system, along with their means and standard deviations to do the data replication check with C-R-X (Tables 1 and 6, ps.262 and 266, in particular).

The source of the data for the yen per dollar exchange rate is the same as that in C-R-X (pp. 261-262): the Database Retrieval System (v2.11), available at http://pacific.commerce.ubc.ca/xr/. The monthly percentage changes are computed between the *ends* of two adjacent months as $(s_t - s_{t-1}) \times 100$ with s_t denoting logged end-of-month exchange rate; the monthly percentage changes and the raw (unlogged) month-end yen per dollar rates are provided, respectively, under columns "jpusfx"and "jpusfxr" in Table 33.

The U.S. and Japanese price indices (p_t^* and p_t) are logs of price indices constructed from (and hence *implied* by) the estimates of pure inflation rates that C-R-X extracted from the stock returns; these estimated pure inflation rates are also provided in Tables 33 and 34; how C-R-X (p.274) construct the price indices will be shown later. The official CPI inflation rates (in percent per month), tabulated in Tables 33 and 34, are also considered and their data source is also exactly the same as that used by C-R-X (p.261). While the official CPI inflation rates are being saved in one of the data files downloadable from the American Economic Association (AEA) Website, C-R-X's estimated pure inflation rates extracted from the stock returns are not and must be computed and saved by one of the program files downloaded from the AEA Website. The details of the files involved in the latter will be described in the next subsection.

Investigated in the following subsections are the time series features of C-R-X's extracted inflation rates and how they are related to the official inflation rates.

2.1 Time series features of inflation rates extracted from stock returns and official inflation rates

C-R-X extract a proxy for realized pure inflation rates from stock returns, which they call the "extracted risk-free rate," denoted by $\hat{R}_{ft}$. The time-series of the extracted risk-free rate is, however, not explicitly displayed in any tabular or graphical format in C-R-X.[5] Tables 33

[5]Its summary statisitics are provided for four countries studied; see their Tables 1 and 6.

and 34 in the data appendix are thus constructed to provide a full set of the extracted risk-free rate time series for Japan and the U.S. The extracted risk-free rate will play a critical role in the present analysis as well, so it is worthwhile studying the extracted series as a supplementary investigation.

C-R-X (Table 11, pp.273-274) provide evidence based on the Johansen (1988, 1991) multivariate cointegration test (of the null of no cointegration) that for each of the four countries studied there is a long-run relation between the price index implied by the extracted risk-free rate and the official CPI. The graphical check of this long-run relation, which is not presented by C-R-X, will be supplemented here.

2.1.1 Summary statistics: Data check for replication

The extracted risk-free rates $\hat{R}_{ft}$ are computed and saved by C-R-X's program "us3cff3ftabvi.m," which is written in MATLAB to construct their Tables 1 and 6.[6] The extracted risk-free rates thus saved and tabulated in Tables 33 and 34 in the data appendix are examined, for replication, by comparing their means and standard deviations computed by the program "us3cff3ftabvi.m," with the counterparts in C-R-X's Tables 1 and 6; Table 6 reports the summary statisitics of the differentials defined as "the difference between a foreign and the U.S. extracted risk-free rates" $\hat{R}^i_{ft} - \hat{R}^U_{ft}$ with i and U denoting, respectively, a foreign country and the U.S. (C-R-X, p.265).

Table 1 associates notation used in the graphs of the present paper with that in C-R-X.

C-R-X's Table 1 One minor or subtle difference has been found in the standard deviations between their Table 1 and the computational results from C-R-X's program "us3cff3ftabvi.m";[7] see below for details. On the other hand, no such differences are detected in the summary statistics of the official (CPI) inflation rates, three Fama-French factors, T-bills and exchange rate changes data, all of which are actually being stored in such input data files in text format as "cpippi.txt," to be simply read

[6] The last update date of the program "us3cff3ftabvi.m" submitted to the *AER* is July 9, 2004 (as remarked in the program). I am grateful to Richard Roll, one of the coauthors, who helped download all the files from the AEA Website and check them with me.

[7] The same holds with the extracted risk-free rates for U.K. and Germany.

Table 1 Notation

Notation in the Graphs		Notation Following C-R-X[a]
eRf1, ERF1		$\hat{R}^J_{ft}$
eRf2, ERF2		$\hat{R}^U_{ft}$
gp1_c, GP1_C		$\pi^J_{CPI,t}$
gp2_c, GP2_C		$\pi^U_{CPI,t}$
p1, P1	p_t	log of $P^J_{R,t}$
p2, P2	p^*_t	log of $P^U_{R,t}$
p1_c	cpi_t	log of $P^J_{CPI,t}$
p2_c	cpi^*_t	log of $P^U_{CPI,t}$
e12, E12	s_t	log of month-end yen per dollar exchange rate

[a]Superscripts, J and U, denote, respectively, Japan and U.S.

into the program; for their means and standard deviations, see Tables 33 and 34 in the data appendix.

Shown in Table 2 are the summary statistics actually computed by the program "us3cff3ftabvi.m" (except for Skewness, Kurtosis and Jarque-Bera which are computed additionally by my RATS program).[8] Throughout the paper (except the data appendix), multiplying $\hat{R}^J_{ft}$, $\pi^J_{CPI,t}$, $\hat{R}^U_{ft}$ and $\pi^U_{CPI,t}$ by 100 gives percent-per-month figures. The logged price indices p^*_t and p_t are logs of price indices $P^U_{R,t}$ and $P^J_{R,t}$ as constructed by C-R-X (p.274): with $P^i_{R,0} = 1$, $P^i_{R,t} = P^i_{R,t-1}(1 + \hat{R}^i_{ft})$. The CPIs are similarly constructed: with $P^i_{CPI,0} = 1$, $P^i_{CPI,t} = P^i_{CPI,t-1}(1 + \pi^i_{CPI,t})$. p_t, cpi_t, p^*_t and cpi^*_t are the logs of these constructed price indices, as defined in Table 1.

Asterisked in Table 2 are the statistics that coincide with those in C-R-X's (p.262) Table 1, whereas the standard deviations with superscript † differ slightly from theirs. C-R-X's standard deviations for $\hat{R}^J_{ft}$ and $\hat{R}^U_{ft}$ are, respectively, 0.0873 and 0.0715.[9] (See also the histograms drawn later for graphical checks of Skewness, Kurtosis(Exc) and Jarque-Bera

[8]See Doan (2004, *RATS*) for RATS program.

[9]C-R-X's standard deviations for the extracted risk-free rates for U.K. and Germany are, respectively, 0.0822 and 0.0737 in their Table 1, whereas those computed by the program "us3cff3ftabvi.m" are, respectively, 0.0802 and 0.0682. The difference thus prevails across all four countries. The reason behind the difference observed across all the four countries, though with perfect matches in the means, is unclear.

Table 2 Summary Statitics: Monthly Data From 1983:05 To 1999:12 (Observations=200)

Time Series	Sample Mean	Standard Deviation	Skewness[a]	Kurtosis(Exc)[b]	Jarque-Bera[c]
$\hat{R}^J_{ft}$	0.00297*	0.07790†	-0.24072 (0.167)[d]	1.81136 (0.000)[e]	29.27347 (0.000)[f]
$\pi^J_{CPI,t}$	0.00100*	0.00468†	0.72408 (0.000)	1.45584 (0.000)	35.13870 (0.000)
$\hat{R}^U_{ft}$	0.00482*	0.06802†	0.56775 (0.001)	3.26206 (0.000)	99.42001 (0.000)
$\pi^U_{CPI,t}$	0.00266*	0.00202†	0.26016 (0.136)	1.57887 (0.000)	23.02987 (0.000)

[a]See Doan (2004, *RM*, p.395) for the formulas.

[b]Excess Kurtosis.

[c]See Jarque and Bera (1987).

[d]P-value (Sk=0). Zero in Sk=0, Ku=0 and JB=0 is the population value if the associated series is i.i.d. Normal; see Doan (*RM*, p.393).

[e]P-value (Ku=0).

[f]P-value (JB=0).

in the table.)

C-R-X's Table 6 Further, C-R-X's program "us3cff3ftabvi.m" computes the summary statisitics for the differentials being as defined earlier; here, though not reported, perfect matches are observed between them and those in their Table 6.

Based on this perfect match for C-R-X's Table 6, we could safely accept the estimated inflation rates $\hat{R}^i_{ft}$, $i = J, U$, computed by the program "us3cff3ftabvi.m" (and tabulated as percent per month in Tables 33 and 34), as the estimates of pure inflation rates that C-R-X actually extracted.

2.1.2 Time series plots and histograms

Drawn in Figures 1 through 8 is a set of plots for $\hat{R}^i_{ft}$ and $\pi^i_{CPI,t}$, $i = J, U$, and another set for p_t, cpi_t, p^*_t and cpi^*_t; see Table 1 for the symbols used in the graphs. And Figures 9 and 10 plot the histograms for $\pi^i_{CPI,t}$ and $\hat{R}^i_{ft}$, for Japan and the U.S. The relationships between the two inflation rates and between the price indices are statistically studied in the following section.

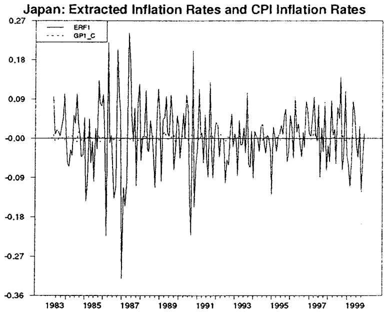

Figure 1 Japan: Extracted Inflation Rates $\hat{R}^J_{ft}$ (eRfl) and CPI Inflation Rates $\pi^J_{CPI,t}$ (gpl_c). See also Table 1 for the notation.

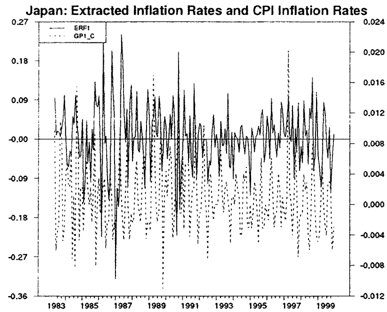

Figure 2 Japan: Extracted Inflation Rates and CPI Inflation Rates, with $\pi^J_{CPI,t}$ on the Right-side Scale. See Figure 1 for the notation.

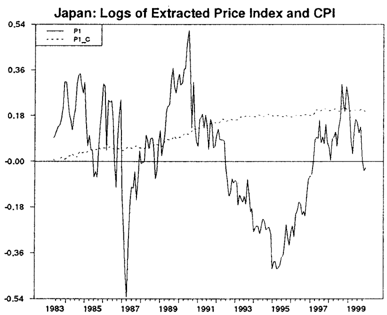

Figure 3 Japan: Logs of Extracted Price Index p_t (pl) and CPI cpi_t (pl_c). See also Table 1 for the notation.

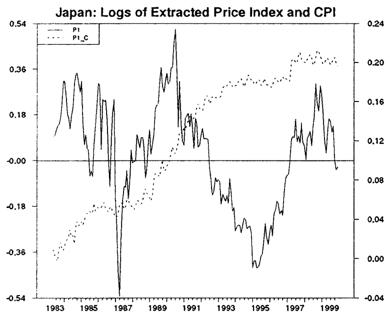

Figure 4 Japan: Logs of Extracted Price Index and CPI, with cpi_t on the Right-side Scale. See Figure 3 for the notation.

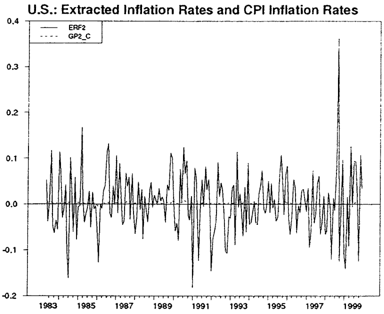

Figure 5 U.S.: Extracted Inflation Rates $\hat{R}^U_{ft}$ (eRf2) and CPI Inflation Rates $\pi^U_{CPI,t}$ (gp2_c). See also Table 1 for the notation.

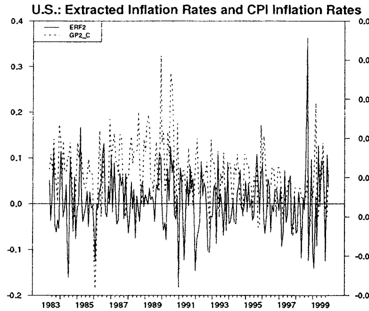

Figure 6 U.S.: Extracted Inflation Rates and CPI Inflation Rates, with $\pi^U_{CPI,t}$ on the Right-side Scale. See Figure 5 for the notation.

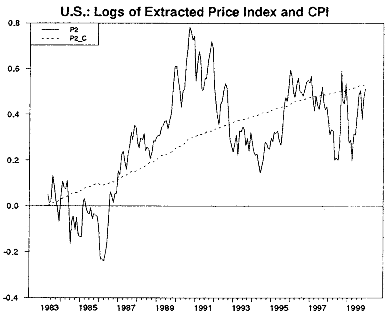

Figure 7 U.S.: Logs of Extracted Price Index p^*_t (p2) and CPI cpi^*_t (p2_c). See also Table 1 for the notation.

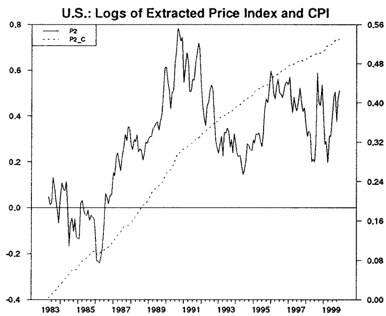

Figure 8 U.S.: Logs of Extracted Price Index and CPI, with cpi^*_t on the Right-side Scale. See Figure 7 for the notation.

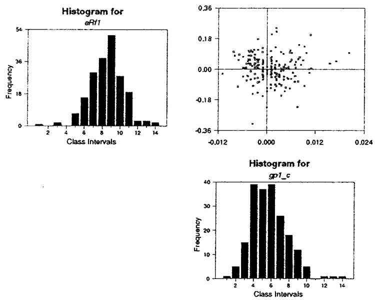

Figure 9 Japan: Histograms and Scatter Diagram for Extracted Inflation Rates $\hat{R}^J_{ft}$ (eRf1) and CPI Inflation Rates $\pi^J_{CPI,t}$ (gp1_c).

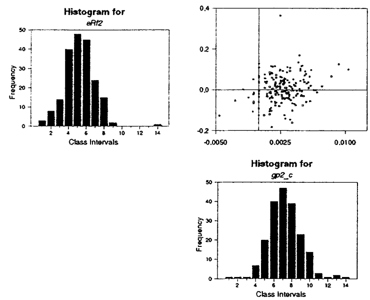

Figure 10 U.S.: Histograms and Scatter Diagram for Extracted Inflation Rates $\hat{R}^U_{ft}$ (eRf2) and CPI Inflation Rates $\pi^U_{CPI,t}$ (gp2_c).

2.2 Time series relationships between extracted and official inflation rates/price indices

2.2.1 Inflation rates

Scatter diagrams Figures 9 and 10 plot the scatter diagrams for CPI inflation rates $\pi^i_{CPI,t}$ and extracted inflation rates $\hat{R}^i_{ft}$, for Japan and the U.S.: as also documented in the correlation matrix in Table 3, *in short run*, the U.S. data exhibit a positive relation, while the Japanese a weaker positive relation, both of which supplement C-R-X's (Table 11, pp.273-274) Johansen test-based evidence, *in levels*, of the presence of *long-run* cointegration relation between the official and extracted price indices, for Japan and U.S. (and GM and UK as well).

Sample cross correlations Plotted in Figure 11 is the sample cross correlation function (SCCF) at lag l (or lead $|l|, l < 0$) that is an SCCF between the extracted $\hat{R}^i_{ft}$ and official inflation rates $\pi^i_{CPI,t-l}$; the contemporaneous cross correlations, SCCF at lag $l = 0$ are exactly equal to those computed in Table 3. We see that the Japanese $\hat{R}^J_{ft}$ and $\pi^J_{CPI,t-l}$ tend to be correlated at *more distant* lags $|l|$ than their U.S. counterparts, $\hat{R}^U_{ft}$ and $\pi^U_{CPI,t-l}$; the contemporaneous correlation is not significantly different from zero for the Japanese, while it is for the U.S. counterparts.

Table 3 Correlation Matrix: Monthly Data From 1983:05 To 1999:12

	$\hat{R}^J_{ft}$	$\pi^J_{CPI,t}$	$\hat{R}^U_{ft}$	$\pi^U_{CPI,t}$
$\hat{R}^J_{ft}$	1.000			
$\pi^J_{CPI,t}$	**0.019**	1.000		
$\hat{R}^U_{ft}$	-0.019	-0.026	1.000	
$\pi^U_{CPI,t}$	-0.039	0.055	**0.225**	1.000

Table 4 Ljung-Box Q-Statistics:[a] Monthly Data From 1983:05 To 1999:12

	$\hat{R}^J_{ft}$ and $\pi^J_{CPI,t}$	$\hat{R}^U_{ft}$ and $\pi^U_{CPI,t}$
Q(1 to 20)	30.900	29.697
	(0.056)[b]	(0.074)
Q(-20 to -1)	41.120	17.029
	(0.003)	(0.651)
Q(-20 to 20)	72.096	56.969
	(0.001)	(0.049)

[a]Ljung-Box Q-statistics for Figure 11.

[b]P-value.

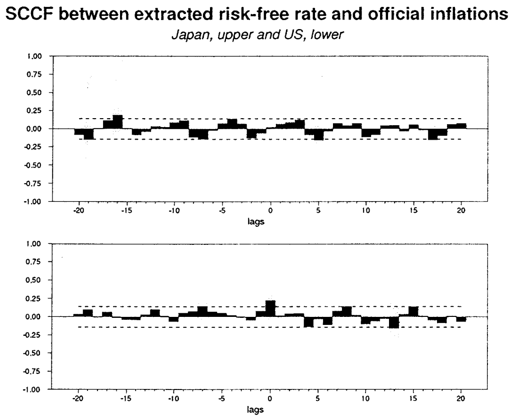

Figure 11 SCCF at Lag l (or Lead $|l|, l < 0$) is a Sample Cross Correlation Function between $\hat{R}^i_{ft}$ and $\pi^i_{CPI,t-l}$, $i = J, U$. Shaded in the figure are cross correlations greater than two standard errors ($= 2\times\sqrt{1/200} = 0.14142$, where 200 is the number of observations of $\hat{R}^i_{ft}$). See Table 4 for the Ljung-Box Q-Statistics of the SCCF here.

Simple regressions of extracted inflation rates on official inflation rates As reported in Table 5, a simple regression of extracted inflation rates $\hat{R}^i_{ft}$ on official inflation rates $\pi^i_{CPI,t}$ will also yield a positive relation between them, which is, again for $i = J$, statistically insignificant. The regression lines are drawn in Figures 12 and 14. The corresponding residuals are plotted in Figures 13 and 15: they appear stationary, though with one or two possible outliers being present in each residual series.

Table 5 Linear Regression - Estimation by Least Squares

Dependent Variable	$\hat{R}^J_{ft}$	$\hat{R}^U_{ft}$
Constant	0.003 (0.639)[a]	−0.015 (0.049)
$\pi^J_{CPI,t}$	0.323 (0.785)	
$\pi^U_{CPI,t}$		7.598 (0.001)
Monthly Data	1983:05 To 1999:12	1983:05 To 1999:12
Observations	200	200
Degrees of Freedom[b]	198	198
Adjusted R^2	−0.005	0.046
Residual Standard Deviation[c]	0.078	0.066
Regression F(1,198)	0.074 (0.785)	10.577 (0.001)
Durbin-Watson Statistic	1.964	1.972

[a]P-value.

[b]Degrees of freedom of residuals (= number observations − number of a constant and independent variables).

[c]Computed as [Sum of Squared Residuals/Residuals Degrees of Freedom]$^{\frac{1}{2}}$.

2.2.2 Price indices: The Engle-Granger cointegration test of extracted price index and official CPI

Turning now to the level data, p_t, cpi_t, p^*_t and cpi^*_t (which are, respectively, logs of $P^J_{R,t}$, $P^J_{CPI,t}$, $P^U_{R,t}$ and $P^U_{CPI,t}$, all constructed from $\hat{R}^i_{ft}$ and $\pi^i_{CPI,t}$, $i = J, U$), here we will look at an equilibrium/long-run regression line. Steps of the Engle-Granger (1987) single equation methodology of cointegration test are employed, though only first two steps will suffice here.[10]

[10]The four-step methodology is summarized by Enders (2004, pp.337-339); note that throughout the paper Enders (2004) will be referred to simply as Enders, to distinguish it from Enders (1995) which appears only in Table 6. Steps 3 and 4

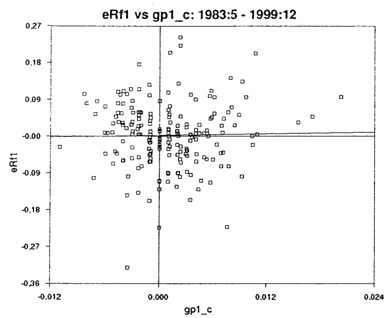

Figure 12 Japan: Extracted Inflation Rates $\hat{R}^J_{ft}$ (eRf1) and CPI Inflation Rates $\pi^J_{CPI,t}$ (gp1_c). The regression line as estimated in Table 5 is also drawn.

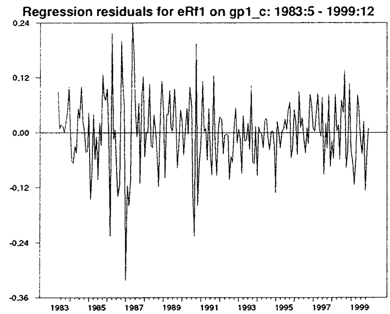

Figure 13 Japan: Residuals from the Regression in Table 5. The regression line is drawn in Figure 12.

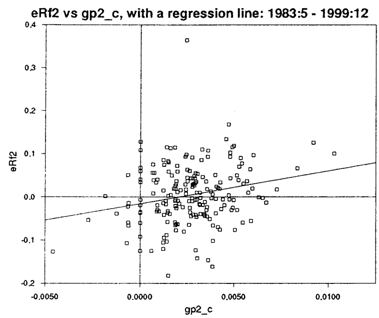

Figure 14 U.S.: Extracted Inflation Rates $\hat{R}^U_{ft}$ (eRf2) and CPI Inflation Rates $\pi^U_{CPI,t}$ (gp2_c). The regression line as estimated in Table 5 is also drawn.

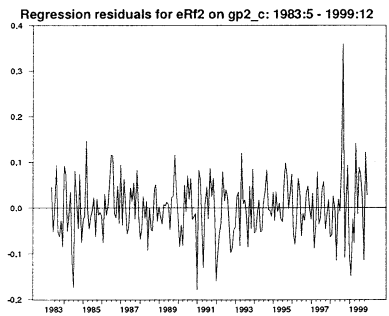

Figure 15 U.S.: Residuals from the Regression in Table 5. The regression line is drawn in Figure 14.

of the Engle-Granger methodology are omitted, for these steps involve building and estimating the VEC model which, for pairs of p_t and cpi_t and of p^*_t and cpi^*_t, does not constitute our main focus.

Table 6 Unit-Root Tests[a]

	Dickey-Fuller[b]	Phillips-Perron[c]	
Regression Run	1984:06 to 1999:12	1983:10 to 1999:12	1984:06 to 1999:12
Observations	188	200	200
With intercept and trend with	12 lags	4 lags in the error process	12 lags in the error process
T-test statistic for:			
p_t^*	−1.928	−2.645	−2.525
p_t	−2.071	−2.807	−2.821
cpi_t^*	−1.211	0.209	−0.071
cpi_t	−1.317	−0.689	−0.889
s_t	−2.437	−1.771	−1.833

[a]See Dickey and Fuller (1979), Phillips (1987) and Phillips, and Perron (1988). For the difference between the tests see Doan (*UG*, p.242); for inclusion of the trend term, see Doan (*UG*, p.246).

[b]The unit-root test for a single variable may be conducted, for example, following Enders (1995, pp.222-223, 256-258); the general form of univariate ADF regression is $\Delta y_t = \sum_{l=1}^{p} \phi_l^{\Delta} \Delta y_{t-l} + \alpha y_{t-1} + \mu + \psi t + u_t$ where Δ is the first difference operator. The null of a unit root is $\alpha = 0$. (The equation will be compared later with eq. (7) in section 3.) The ADF statistics adjust for autocorrelation using an autoregressive representation. Critical values that apply here are: 1%= −4.010, 5%= −3.435, 10%= −3.141.

[c]The PP statistics correct for autocorrelation using a non-parametric correction. Critical values that apply here are: 1%= −4.007. 5%= −3.433, 10%= −3.140

A single equation approach: Engle-Granger static modeling versus dynamic modeling The motivation behind using the Engle-Granger single equation methodology is primarily to complement C-R-X's (p.274) Johansen multivariate cointegration test.

Several remarks on the Engle-Granger single equation methodology are in order:

(i) Only when the number n of variables in the VEC model is 2 is it possible to show there is an unique cointegration vector, in which case estimating a single equation would be appropriate; (ii) if single equation methods are to be used, *dynamic* modeling approach is most likely to produce unbiased estimates of the long-run relationship and the test with that approach is more powerful against the usual Engle-Granger *static* model (Harris 1995, p.72); and (iii) as capturing the relationship between x_t and y_t requires more complicated dynamic models, such as eq. (3) later, then estimating instead the static model (1) below, to obtain an estimate of the long-run parameter β_1, will push more complicated dynamic terms into the residual, v_t, with the result that the static model can exhibit severe autocorrelation (Harris, p.60).

Two approaches to the cointegration test are thus employed later in Step 2: Step 2A is the usual Engle-Granger static, single-equation modeling, and Step 2B a dynamic, single-equation approach; the latter relies on the dynamic version of the Engle-Granger static modeling.

Step 1: Unit-root tests for the price series As reported in Table 6, the augmented Dickey-Fuller (ADF) test (with trend and lags=12) and the Phillips-Perron (PP) test (with trend and lags=4/lags=12) cannot reject, at any conventional levels of significance, the null of a unit root for all (logged) p_t, cpi_t, p_t^* and cpi_t^*. The unit-root test results here are consistent with C-R-X (p.274) for the logged series: all series appear to be integrated of order one, $I(1)$. (The Johansen approach-testing for unit roots will be carried out in section 3, yielding the same results as those here: see Table 15 there.) The same non-rejection of the null is also seen to result from the unit-root tests for (logged) exchange rate.

Step 2A: Unit-root tests for the residuals, to check on the cointegration: a static approach The Engle-Granger *static*, long-run model is estimated by an OLS method, with β_0 and β_1 being the cointegrating or long-run parameters:

$$y_t = \beta_0 + \beta_1 x_t + v_t, \tag{1}$$

with $y_t = p_t$ or p_t^*, and $x_t = cpi_t$ or cpi_t^*. Setting y_t and x_t as such follows C-R-X (pp.260-261) arguing that exctracted risk-free rates, based on which both p_t and p_t^* are constructed, "contain substantial noise."[11]

The *statistically significant, negative* slope coefficient in the p_t model reported in Table 7 appears *in*consistent with C-R-X's (Panel B of Table 11, p.274) *insignificant, positive* estimate of the cointegrating coefficient of the "cointegration regression" with an intercept. (The sign of the constant estimate here is also inconsistent with C-R-X's.)

On the other hand, the result of the p_t^* model appears consistent with C-R-X (Panel B of Table 11, p.274), though the sign and siginificance of the constant estimate is not consistent with C-R-X's.

The same inconsistency results (not reported here) as above apply to the (raw) unlogged price indices and C-R-X (Panel A of Table 11, p.274).

[11] Setting, instead, $y_t = cpi_t$ or cpi_t^*, and $x_t = p_t$ or p_t^*, will yield entirely different, and insensible, regression results with, for example, much worse Durbin-Watson statistics; the results are not reported here.

Table 7 Linear Regression - Estimation by Least Squares

Dependent Variable	p_t	p_t^*
Constant	0.163 (0.000)	0.038 (0.133)
cpi_t	-1.104 (0.000)	
cpi_t^*		0.902 (0.000)
Monthly Data	1983:05 To 1999:12	1983:05 To 1999:12
Usable Observations	200	200
Degrees of Freedom	198	198
Adjusted R^2	0.121	0.417
Residual Standard Deviation	0.195	0.172
Regression F(1,198)	28.417 (0.000)	143.433 (0.000)
Durbin-Watson Statistic	0.166	0.151

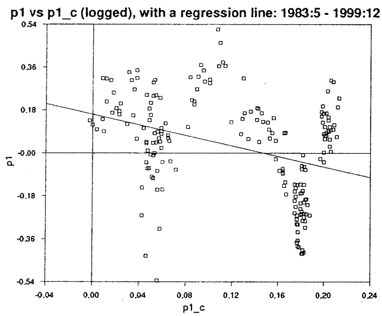

Figure 16 Japan: Logged Price Indices p_t (p1) versus cpi_t (p1_c). The Engle-Granger static, long-run model (1) is estimated: the regression line as estimated in Table 7 is also drawn.

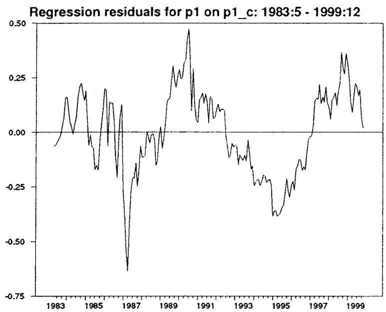

Figure 17 Japan: Residuals from the p_t Model (1). The regression line is drawn in Figure 16.

What accounts for all these inconsistencies? Given the satisfactory data check made for replication in section 2, one most likely explanation is that our eq. (1) could differ from C-R-X's (Table 11, p.274) "cointegration regression" equation. If so, then we would need not to be so concerned over the apparent inconsistencies here. What is more critical for the cointegration test of the extracted and official price indices is the unit-root test results for the residuals from the regressions, which *must* coincide with C-R-X. So we now turn to the test.

The regression residuals are plotted in Figures 17 and 19; the unit-

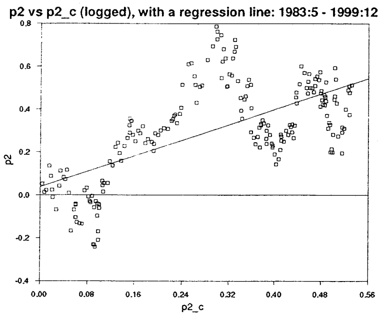

Figure 18 U.S.: Logged Price Indices p_t^* (p2) versus cpi_t^* (p2_c). The Engle-Granger static, long-run model (1) is estimated: the regression line as estimated in Table 7 is also drawn.

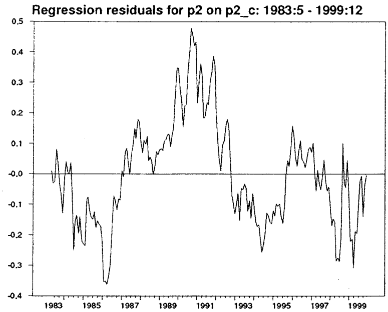

Figure 19 U.S.: Residuals from the p_t^* Model (1). The regression line is drawn in Figure 18.

root test for the residuals is carried out in Table 8, to check on the cointegration. Both the ADF and PP tests reject the null of a unit root at the 5-percent level (except for the p_t^* regression residuals under the PP test rejecting the null at the 10-% level), for both residuals from the p_t and p_t^* regressions. The residual sequences thus both appear stationary, which, together with p_t, cpi_t, p_t^* and cpi_t^* being all $I(1)$ as has been shown in Table 6, implies that (Japanese) p_t and cpi_t are cointegrated and so are (U.S.) p_t^* and cpi_t^*. This *is* consistent with C-R-X's (p.274) Johansen (multivariate-system) cointegration test rejecting the null of no cointegration for the Japanese and U.S. price series. (Similar results are obtained for the raw, unlogged series as well, though they are not reported here.)

Step 2B: Unit-root tests for the residuals, to check on the cointegration: a dynamic approach Step 2B here is designed following the remarks (i) through (iii) made at the begnning of the present subsection.

The dynamic version of the Engle-Granger static, long-run model *with first differences being likely to capture short-run dynamics* is now estimated: we start out with a simple dynamic model of short-run adjust-

Table 8 Unit-Root Tests for Residuals from Static Models (1)

	Dickey-Fuller[a]	Phillips-Perron[b]	
Regression Run	1984:06 to 1999:12	1983:10 to 1999:12	1984:06 to 1999:12
Observations	188	200	200
Without intercept and trend[c] with	12 lags		
With intercept and trend with		4 lags in the error process	12 lags in the error process
T-test statistic for:			
p_t regression residuals	−2.198	−2.882	−2.899
p_t^* regression residuals	−1.993	−2.777	−2.659

[a]Critical values that apply here are: 1%= −2.577, 5%= −1.941, 10%= −1.617.

[b]Critical values that apply here are: 1%= −3.464, 5%=−2.876, 10%= −2.574.

[c]Both intercept and trend are omitted. For a residual sequence, there is no need to include an iterecept term in the ADF test (Enders, pp.336). There is no such an option for the PP test in Doan (*RATS*).

ment and its rewritten version (Harris, pp.52-53):

$$y_t = \delta + \gamma_0 x_t + \gamma_1 x_{t-1} + \alpha y_{t-1} + u_t \tag{2}$$
$$y_t = \beta_0 + \beta_1 x_t + \lambda_1 \Delta x_t + \lambda_2 \Delta y_t + v_t, \tag{3}$$

where $\beta_0 = \delta/(1-\alpha), \beta_1 = (\gamma_0 + \gamma_1)/(1-\alpha), \lambda_1 = -\gamma_1/(1-\alpha), \lambda_2 = -\alpha/(1-\alpha), v_t = u_t/(1-\alpha)$, and with $y_t = p_t$ or p_t^*, and $x_t = cpi_t$ or cpi_t^*. First differences are introduced here (in eqs. (4), and (5) below) on a purely statistical ground (without assuming an original, economic model like (2)), that is, based on those statistically significant SCCFs between the extracted risk-free rate R_{ft}^i and official inflation rate $\pi_{CPI,t-l}^i$ at lag or lead l, as drawn in Figure 11.

For $y_t = p_t$ and $x_t = cpi_t$,

$$y_t = \beta_0 + \beta_1 x_t + \lambda_1 \Delta x_{t+19} + \lambda_2 \Delta x_{t-5} + \lambda_3 \Delta y_t + v_t \tag{4}$$

where Δx_{t+19} and Δx_{t-5} are included based on the statistically significant SCCFs at $l = -19$ and 5 (nineteen leads and five lags) as shown in Figure 11.[12]

[12]With $\Delta x_{t-l} = x_{t-l} - x_{t-l-1}$ and $x_t = cpi_t, x_{t-l} = x_{t-l-1}(1 + \pi_{CPI,t-l})$. Incidentally, in Figure 11 SCCFs at lead $l = -16$ and lag $l = 17$ are also statistically significant, and yet their corresponding slope coefficients turn out insignificant (the results are not reported here); they are thus omitted from the model (4).

Again, the statistically significant, negative slope coefficient of cpi_t in the p_t regression results in Table 9 appears inconsistent with C-R-X's (Panel B of Table 11, p.274) insignificant, positive estimate of the cointegrating coefficient of the "cointegration regression" (with an intercept).

Table 9 Linear Regression - Estimation by Least Squares

Dependent Variable	Eq. (4): p_t	Eq. (5): p_t^*
Constant	0.205 (0.000)	-0.111 (0.007)
cpi_t	-1.653 (0.000)	
Δcpi_{t+19}	-7.852 (0.040)	
Δcpi_{t-5}	9.561 (0.007)	
Δp_t	0.517 (0.003)	
cpi_t^*		1.014 (0.000)
Δcpi_t^*		23.769 (0.000)
Δcpi_{t-13}^*		19.302 (0.002)
Δp_t^*		0.447 (0.018)
Monthly Data	1983:11 To 1998:05	1984:07 To 1999:12
Usable Observations	175	186
Degrees of Freedom	170	181
Adjusted R^2	0.268	0.448
Residual Standard Deviation	0.185	0.165
Regression F(4,170)	16.922 (0.000)	38.587 (0.000)
Durbin-Watson Statistic	0.192	0.242

For $y_t = p_t^*$ and $x_t = cpi_t^*$,

$$y_t = \beta_0 + \beta_1 x_t + \lambda_1 \Delta x_t + \lambda_2 \Delta x_{t-13} + \lambda_3 \Delta y_t + v_t \tag{5}$$

where Δx_t and Δx_{t-13} are included based on the statistically significant SCCFs at $l = 0$ and 13 (lags), as drawn in Figure 11.[13]

In Table 9, the p_t^* regression result appears consistent with the U.S. result in C-R-X (Panel B, Table 11, p.274), with respect to both sign and statistical significance of the intercept and the coefficient of cpi_t^*, in which sense it improves upon the previous, corresponding Step 2A result, although the magnitude of each parameter is somewhat different between the two results.

The residuals from the p_t and p_t^* regressions are plotted, respectively, in Figures 20 and 21; the unit-root test for each of the residuals is carried out now in Table 10, to check on the cointegration. Both the ADF and PP tests reject the null of a unit root at either the 5-percent or 10-percent level, for either of residuals from the p_t and p_t^* regressions. Both

[13]The statistically significant SCCF is also observed at $l = -7$ (leads), but including the corresponding Δx_{t+7} does not lead to a significant slope coefficient; the results are not reported here.

residual sequences thus appear stationary, which, together with p_t, cpi_t, p_t^* and cpi_t^* being all $I(1)$ as has been shown in Table 6, implies that (Japanese) p_t and cpi_t are cointegrated and so are (U.S.) p_t^* and cpi_t^*.

As desired, the results are (again in Step 2B here) consistent with C-R-X's (Panel B in Table 11, p.274) Johansen (multivariate-system) cointegration tests rejecting the null of no cointegration for the Japanese as well as U.S. price series.

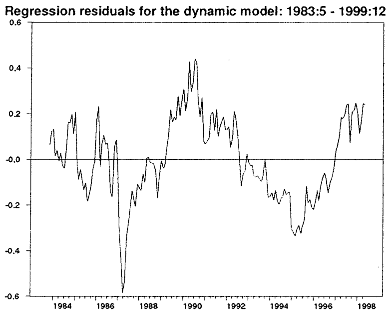

Figure 20 Japan: The p_t Regression Residuals. The *dynamic* version (4) of the Engle-Granger static, long-run model is estimated: see Table 9 for the estimated results.

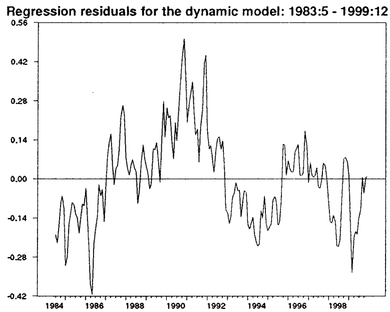

Figure 21 U.S.: The p_t^* Regression Residuals. The *dynamic* version (5) of the Engle-Granger static, long-run model is estimated: see Table 9 for the estimated results.

Summary of the data check The single equation-based unit-root test results for the residuals in Steps 2A and 2B (together with those in Step 1) do coincide with those of C-R-X's (Table 11) Johansen multivariate VAR-based cointegration test: this would satisfactorily confirm the validity of the price indices p_t and p_t^* that are constructed from the estimated pure inflation rates $\hat{R}^i_{ft}, i = J, U$ which are computed by C-R-X's program "us3cff3ftabvi.m" and as tabulated in Tables 33 and 34. In sum, we have replicated C-R-X's Tables 1, 6 and 11, though with some minor differences found.

Table 10 Unit-Root Tests for Residuals from Dynamic Models (4) and (5)

	Dickey-Fuller	Phillips-Perron	
For p_t regression (4) residuals[a]			
Regression Run	1984:12 to 1998:05	1983:10 to 1999:12	1984:12 to 1998:05
Observations	163	200	175
Without intercept and trend with	12 lags		
With intercept and trend with		4 lags in the error process	12 lags in the error process
T-test statistic	-2.097	-2.882	-2.769
For p_t^* regression (5) residuals[b]			
Regression Run	1985:08 to 1999:12	1984:12 to 1999:12	1985:08 to 1999:12
Observations	174	186	186
Without intercept and trend with	12 lags		
With intercept and trend with		4 lags in the error process	12 lags in the error process
T-test statistic	-1.815	-3.447	-3.358

[a] For the ADF, critical values that apply here are: 1%= −2.578, 5%= −1.942, 10%= −1.617. For the PP, critical values that apply here are: 1%= −3.464, 5%=−2.876, 10%= −2.574.

[b] For the ADF, critical values that apply here are: 1%= −2.577, 5%= −1.942, 10%= −1.617. For the PP, critical values that apply here are: 1%= −3.467, 5%=−2.877, 10%= −2.575.

3 Long-run Structure of Exchange Rate and Extracted Price Indices

To explore the long-run structure of the exchange rate and extracted price indices requires a test for unit roots, which is in the present paper equivalent to a multivariate cointegration testing of PPP. A better basis for examining the number of unit roots in a vector of variables is given by the multivariate cointegration methodology of Johansen (MacDonald and Marsh, p.33); specifically, the multivariate form of the ADF test will be used, with a null of stationarity (rather than a null of nonstationarity).

3.1 Models and preliminary study

3.1.1 The underlying VAR model and its VEC model

Multivariate cointegration tests of PPP are conducted using the Johansen method, to examine the long-run structure (the number of unit roots)[14] in the vector $\boldsymbol{y}_t = (s_t, p_t^*, p_t)'$, each element of which is a potentially endogenous variable[15] and assumed to be integrated of order 1, $I(1)$.[16] . To conduct the tests, we consider the VAR model including a constant and augmented with centered seasonal dummies:[17]

$$\boldsymbol{y}_t = \sum_{l=1}^{L} \boldsymbol{\Phi}_l \boldsymbol{y}_{t-l} + \boldsymbol{\mu} + \boldsymbol{\Psi} \boldsymbol{D}_t + \boldsymbol{u}_t. \tag{6}$$

The underlying VAR model is reformulated in the error-correction form as the VEC model:

$$\Delta \boldsymbol{y}_t = \sum_{l=1}^{L-1} \boldsymbol{\Phi}_l^{\Delta} \Delta \boldsymbol{y}_{t-l} + \boldsymbol{\Pi} \boldsymbol{y}_{t-1} + \boldsymbol{\mu} + \boldsymbol{\Psi} \boldsymbol{D}_t + \boldsymbol{u}_t \tag{7}$$

where: Δ is the first-difference operator; the *short-run* matrices $\boldsymbol{\Phi}_l^{\Delta}$ represent the short-run dynamics/adjustment to past change in $\boldsymbol{y}_t$, $\Delta \boldsymbol{y}_{t-l}$;[18] and the *long-run* marix $\boldsymbol{\Pi}$ represents long-run adjustment. The initial assumptions include, in particular, the white noise $\boldsymbol{u}_t \sim IN(\boldsymbol{0}, \boldsymbol{\Sigma})$ or

[14]For the long-run structure and/or the number of unit roots, see ⊂ 3 ⊃ in section 3.1.2, the paragraph "Restricting, jointly, $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$" in section 3.2.2, and the paragraph "Long-run structure and real exchange rate" in section 3.2.3.

[15]It could turn out wealky exogenous, as will be evidenced in section 3.2.2. The rationale behind choosing the ordering (s_t, p_t^*, p_t) instead of, for example, its reverse (p_t, p_t^*, s_t) is given in section 5.2.

[16]Hansen and Juselius (1995, p.1) remark that "... we assume that $\boldsymbol{y}_t$ is at most $I(1)$ ... However not all the individual variables included in $\boldsymbol{y}_t$ need be $I(1)$, as is often incorrectly assumed. To find cointegration between nonstationary variables, only two of the variables have to be $I(1)$."

[17]For *centered* seasonal dummies, see Hansen and Juselius (p.8) and Doan (*RM*, p.84, pp.367-368); Harris (p.81) remarks that "Seasonal dummies are centered to ensure that they sum to zero over time, and thus they do not affect the underlying asymptotic distributions upon which tests (including tests for cointegration rank) depend."

[18]The terms "short-run matrices" and "short-run dynamics" are those used by Hansen and Juselius (ps.29, 71).

$u_1, ..., u_T$ are $niid(\mathbf{0}, \Sigma)$; the dependence is allowed among the white-noise disturbance terms $u_{1t_1}, u_{2t_2}, u_{3t_3}$ for any t_i, $i = 1, 2, 3$. For monthly data, L may be set at 12; it will be far smaller for our set of the data, however, as shown later.

Short-run effects/dynamics/matrices $\mathbf{\Phi}_l^{\triangle}$, the short-run dynamics/adjustment to past changes in $\boldsymbol{y}_t$, and their estimates are crucial in our analysis of short-run PPP, for C-R-X has shown, using the pure inflation rate they extracted from stock returns, that the short-run PPP is strongly supported.

Note that the analysis of the short-run structure (consisting of short-run effects $\mathbf{\Phi}_l^{\triangle}, l = 1, ..., L-1$) here will be made *after* the modeling of the long-run structure is completed: the estimated cointegration vectors in the long-run structure will be considered as given or known, in the subsequent short-run analysis (in section 4).

Long-run adjustment If the long-run marix $\mathbf{\Pi}$ is either zero or non-zero and full-rank, it is of no use to write the VAR in form (7) rather than (6), to begin with; if it is non-zero but less than full-rank, then it is usefully written as[19]

$$\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}' \tag{8}$$

where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $3\times r$ matrices, with r being the rank of $\mathbf{\Pi}$.[20] Following Engle and Granger's (1987) definition of equilibrium error,[21] $\boldsymbol{\beta}'\boldsymbol{y}_{t-1} \neq 0$ in eq. (7) is interpreted as an equilibrium error, with $\boldsymbol{\beta}$ being a matrix representing long-run coefficients such that the term $\boldsymbol{\beta}'\boldsymbol{y}_{t-1}$, the deviation from long-run equilibrium embedded in eq. (7), represents up to $(n-1)$ cointegration relationships in the multivariate model which ensure that the $\boldsymbol{y}_t$ converge to their long-run steady-state solutions. The rank r indicates the number of cointegration relations $\boldsymbol{\beta}'\boldsymbol{y}_{t-1}$. Assuming

[19] See Doan (*UG*, p.360). In this case, eq. (7) *without* the term $\mathbf{\Pi}\boldsymbol{y}_{t-1}$ would be a misspecified model.

[20] $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are matrices of full rank (see Hansen and Juselius, p.2). The decomposition in eq. (8) is not unique; where r is one, it is unique up to a scale factor in the two parts (see Doan, *UG*, p.360).

[21] The very beginning of Engle-Granger's formal analysis is to consider a set of n economic variables in long-run equilibrium when $\sum_{i=1}^{n} \beta_i y_{it} = 0$; the equilibrium error is a disequilibrium defined as a deviation from long-run equilibrium and given by $e_t \equiv \sum_{i=1}^{n} \beta_i y_{it}$; in the long run, $e_t = 0$.

$\boldsymbol{y}_t$ is a vector of nonstationary $I(1)$ variables, then all the terms in (7) which involve $\Delta \boldsymbol{y}_t$ are $I(0)$, while $\boldsymbol{\Pi y}_{t-1}$ must also be stationary for $\boldsymbol{u}_t \sim I(0)$ to be white noise (Harris, p.79).[22]

$\boldsymbol{\alpha}$ is a matrix representing a measure of the average speed of convergence towards the long-run equilibrium (i.e., the speed of adjustment to disequilibrium).[23] The elements of $\boldsymbol{\alpha}$ will be shown in section 3.2.3 to indicate how rapidly a current deviation from PPP is offset in the future.

$\boldsymbol{\Pi}$, the long-run adjustment, has been the major topic of interest in the cointegration and error-correction model analysis of PPP. In a way, this is due to the lack of short-run support for PPP in the past PPP literature. Now that C-R-X have found strong support for PPP in the short-run and made available more appropriate inflation rate data for the first time, it is an interesting and meaningful work, using their extracted price data, to statistically examine the short-run dynamics based on the VAR model and its error-correction representation.

Conditional/partial version If one (or more), say s_t, of the three variables turns out to be weakly exogenous to the system under study, then the conditional (or partial) version of eq. (7), conditioned on the weakly exogenous variable s_t, is written as

$$\Delta \boldsymbol{z}_t = \sum_{l=0}^{L-1} \boldsymbol{\Theta}_l^{\triangle} \Delta s_t + \sum_{l=1}^{L-1} \tilde{\boldsymbol{\Phi}}_l^{\triangle} \Delta \tilde{\boldsymbol{y}}_{t-l} + \tilde{\boldsymbol{\Pi}} \tilde{\boldsymbol{y}}_{t-1} + \boldsymbol{\mu} + \boldsymbol{\Psi D}_t + \boldsymbol{u}_t, \qquad (9)$$

where $\boldsymbol{z}_t = (p_t^*, p_t)'$, $\tilde{\boldsymbol{y}}_t = (p_t^*, p_t, s_t)'$ with s_t being the *last* element now and $\tilde{\boldsymbol{\Pi}} = \tilde{\boldsymbol{\alpha}}\tilde{\boldsymbol{\beta}}'$ where $\tilde{\boldsymbol{\alpha}} = \boldsymbol{\alpha}$ with the very *last* row vector being $\mathbf{0}$ and the *last* element of $\tilde{\boldsymbol{\beta}}$ corresponds to the exchange rate s_t. The conditional version of the model (9) will be investigated later in section 4.2.

3.1.2 Preliminary analysis

We conduct the long-run and short-run analyses ignoring the possible presence of outliers in endogenous variables.[24] Carried out at the outset

[22]That the deviation from long-run equilibrium is stationary means that the deviation is temporary in nature. The stationarity requirement imposed on $\boldsymbol{\Pi y}_{t-1}$ is investigated by Kojima (2006c).

[23]See Hansen and Juselius (pp.2-3) and Harris (pp.77-78).

[24]The same results were obtained, too, for the analyses that do take into account possible outlier/non-normality in price time series. Their details are thus not reported here.

is the step-by-step preliminary analysis (which will be later followed by the cointegration test of PPP):

⊂ 1 ⊃ **Graphical analysis** Figure 22 plots the three time series s_t, p^*_t and p_t: it is difficult to check whether they are cointegerated or not, based on the figure. Let us now draw the two series $s_t + p^*_t$ and p_t; recall here that a real exchange rate is defined as a particular linear combination $r_t \equiv s_t + p^*_t - p_t$.

From Figure 23, the two series $q_t = s_t + p^*_t$ and p_t are seen to have a strong positive association, and we could draw an equilibrium regression line indicating the long-run equilibrium relationship between the two. The deviations from the line are called equilibrium error (defined as $e_t \equiv s_t + p^*_t - p_t$) and plotted in Figure 24: the equilibrium error does appear to be an $I(0)$ (i.e., sationary) process and hence, at least, the two series q_t and p_t seem cointegrated.[25] The equilibirium error drawn in Figure 24 will then be the real exchange rate r_t, and the stationarity of the equilibrium error is consistent with absolute PPP that asserts $s_t + p^*_t = p_t$, that is, $r_t = 0$ (MacDonald and Marsh, p.24). (The cointegration of the two series q_t and p_t can be formally tested, with a special focus on the real exchange rate behavior. Such a test does constitute a research topic in Kojima 2006c .)

This preliminary graphical result motivates us to further explore a formal cointegration analysis of the entire three series (s_t, p^*_t and p_t). (Recall that $\boldsymbol{\Pi y}_{t-1}$ which contains an equilibrium error $\boldsymbol{\beta}'\boldsymbol{y}_{t-1}$ must be stationary for $\boldsymbol{u}_t \sim I(0)$ to be white noise.)

⊂ 2 ⊃ **Cointegration rank tests** For (6) and (7), we set $L = 4$, which will be shown later to be long enough based on the residual diagnostic checks (especially, the residual autocorrelaiton tests). L is determined here by Table 12, in such a way that the number of eigenvalues of the companion matrix that are close to unity is sensible: with L=12, for instance, no such sensible number would obtain.[26] In Hansen and Juselius'

[25] See Enders (pp.324-325) for similar simulated illustrations.

[26] As many as eight out of 36 ($= 3 \times L$) roots are close to unity, thought it is not clear why: see column $L = 12$ in Table 12. For $L = 6$, about four roots are closer to unity: see column $L = 6$ in Table 12. For $L = 5$, as in the case of $L = 4$, only two roots are close to unity: see column $L = 5$ in Table 12.

L may be either 4 or 5; we choose the former. It is important that, ex post, $L = 12$ turned out an insensible choice purely for a numerical reason, although, ex ante, it

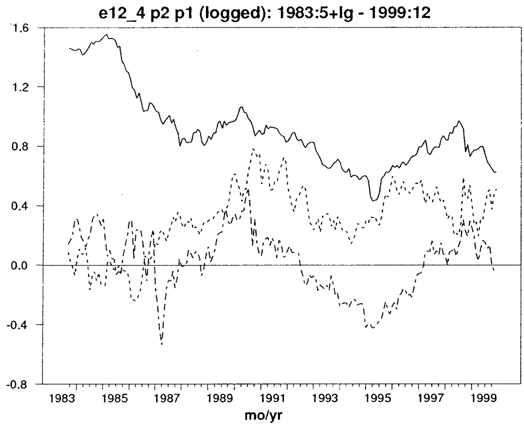

Figure 22 From Top Line: s_t, p_t^* and p_t (e12, p2 and p1) for 1983:9 [=1983:5 plus L (=4 months)] - 1999:12. s_t is shifted vertically downward by 4. See Table 1 for the notation.

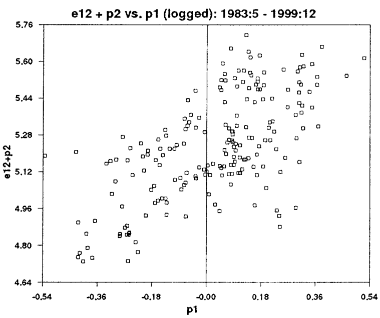

Figure 23 $s_t + p_t^*$ (e12+p2) versus p_t (p1): $1983:5 - 1999:12$.

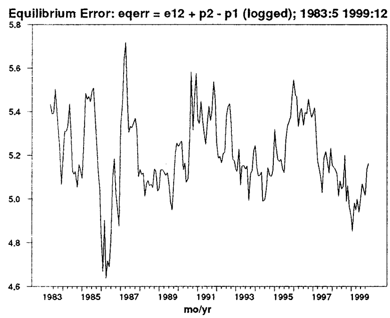

Figure 24 Equilibrium Error (Real Exchange Rate) $s_t + p_t^* - p_t : 1983:5 - 1999:12$.

software, one can choose from between two assumptions about trend: one that each endogenous variable contains linear trends, but such trends are not present in the cointegration relations: and the other that the trends

should be a first choice for a statistical seasonality reason.

exist also in the cointegration space. The former assumption is made in the remaining analysis.

Tables 11 through 13 document the preliminary results of the cointegration analysis and are entirely for the *unrestricted* model where no restrictions are yet imposed on either $\boldsymbol{\alpha}$ or $\boldsymbol{\beta}$ (both $3\times r$ matrices). An unrestricted model is written as eq. (6), the underlying, Sims-type VAR in level and it is equivalent to (7) with $r=3$ (i.e., $\boldsymbol{\Pi}$ of full rank) and the variables in $\boldsymbol{y}_t$ being $I(0)$ to begin with. (We will turn to the *restricted* model after Table 14.) The cointegration rank tests in Table 11 suggests strongly that the null of r (the rank of $\boldsymbol{\Pi}$) $=1$ is not rejected.

Table 11 Cointegration Analysis [1]: Estimated Results of Unrestricted[a] Model (7) with $r=3$ (i.e., $\boldsymbol{\Pi}$ of Full Rank)

Endogeneous series	$s_t\ p_t^*\ p_t$						
Deterministic series	Unrestricted constant; 11 centered seasonal dummies						
Effective sample	1983:09 TO 1999:12						
Lag(s) in VAR-model	4						
No. of observations	196						
Obs.- no.of variables	172						
$I(1)$ Analysis							
Cointegration Rank Tests	Eigenv.	L-max[b]	Trace[c]	H_0: r^d	$p-r^e$	L-max90	Trace90
	0.104	21.45	30.33	0	3	13.39	26.70
	0.028	5.47	8.88	1	2	10.60	13.31
	0.017	3.41	3.41	2	1	2.71	2.71
$\boldsymbol{\beta}'$							
	s_t	p_t^*	p_t				
	6.328	5.921	-6.590				
	1.356	-3.512	-2.268				
	-2.541	-1.423	-2.650				
$\boldsymbol{\alpha}$							
s_t	-0.001	-0.004	0.003				
p_t^*	-0.014	0.006	0.003				
p_t	0.016	0.004	0.007				
$\boldsymbol{\Pi}$							
	s_t	p_t^*	p_t				
	-0.018	0.006	0.008				
	-0.091	-0.111	0.073				
	0.091	0.073	-0.133				

[a]No restrictions are imposed on either $\boldsymbol{\alpha}$ or $\boldsymbol{\beta}$.

[b]For L-max, the null versus alternative hypotheses are $r=0$ vs $r=1$, $r=1$ vs $r=2$, and $r=2$ vs $r=3$.

[c]For Trace, the null versus alternative hypotheses are $r=0$ vs $r\geq 1$, $r\leq 1$ vs $r\geq 2$, and $r\leq 2$ vs $r\geq 3$.

[d]r is the rank of $\boldsymbol{\Pi}$.

[e]p is the number of endogenous variables, which is 3 in the present study.

Further, it is worth noting here at this point that, in Table 11, the very first row vector, $(6.328, 5.921, -6.590)$, of unrestricted β' has the desired signs of its elements, since our hypothesis of PPP restirction on the cointegrating vector will be later shown to be $(1, 1, -1)$. In fact, it is critical at this point to have the signs $(+, +, -)$ for the first row vector of unrestricted β': with differently ordered signs, the hypothesis of PPP restriction would be most likely rejected.

Table 12 Cointegration Analysis [2]: The Eigenvalues of the Companion Matrix; Unrestricted Model (7) with $r = 3$

$L = 4$				$L = 5$			
real	complex	modulus	argument	real	complex	modulus	argument
0.975	0.014	0.975	0.014	0.969	-0.025	0.970	-0.026
0.975	-0.014	0.975	-0.014	0.969	0.025	0.970	0.026
0.717	-0.149	0.732	-0.205	0.735	-0.145	0.749	-0.195
0.717	0.149	0.732	0.205	0.735	0.145	0.749	0.195
0.387	-0.358	0.527	-0.747	0.499	0.491	0.700	0.777
0.3870	0.358	0.527	0.7470	0.499	-0.491	0.700	-0.777
0.055	0.482	0.485	1.456	Remaining 9 Rows Not Reported			
0.055	-0.482	0.485	-1.456				
-0.255	-0.406	0.479	-2.132				
-0.255	0.406	0.479	2.132				
-0.435	-0.181	0.471	-2.748				
-0.435	0.181	0.471	2.748				
$L = 6$				$L = 12$			
real	complex	modulus	argument	real	complex	modulus	argument
0.968	-0.015	0.969	-0.016	0.978	0.000	0.978	0.000
0.968	0.015	0.969	0.016	0.953	0.062	0.955	0.065
0.813	-0.142	0.825	-0.173	0.953	-0.062	0.955	-0.065
0.813	0.142	0.825	0.173	-0.866	0.263	0.905	2.847
0.585	0.531	0.790	0.738	-0.866	-0.263	0.905	-2.847
0.585	-0.531	0.790	-0.738	0.867	0.195	0.889	0.221
Remaining 12 Rows Not Reported				0.867	-0.195	0.889	-0.221
				0.655	-0.599	0.888	-0.741
				Remaining 28 Rows Not Reported			

⊂ 3 ⊃ **Infering** r In Table 12 (listing $3\times L$ roots), column $L = 4$ shows that the number of roots close to unity is two, which is supposed to be equal to $3 - r$, yielding $r = 1$:[27] there is then only one cointegration

[27]Note that: if $r = 0$ (i.e., $\mathbf{\Pi}$ has zero rank), the desired model is the VAR in first differences but *without* long-run (cointegration) elements; the lack of cointegration here suggests that the variables could wander arbitrarily far from each other. If $r = n$ (i.e., $\mathbf{\Pi}$ has full rank), the variables in $\boldsymbol{y}_t$ are $I(0)$ and the underlying, Sims-type VAR in level and the model in error-correction form are equivalent, either of which can be a desired model. If $0 < r < n$ (i.e., $\mathbf{\Pi}$ has reduced rank), which is the only interesting case, the desired model is the VAR in first differences and *with* long-run

relation/vector. Notice that "$r = 1$" here does coincide with the value of r inferred above from the rank tests in Table 11.

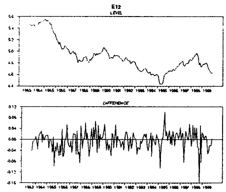

Figure 25 s_t (E12). Upper and lower plots are, respectively, level and first difference; the level covers the period of 1983:9 [=1983:5 plus L (=4 months)] - 1999:12 (the same applies to the two figures that follow).

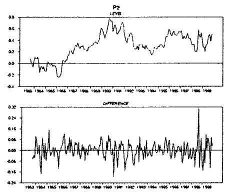

Figure 26 p_t^* (P2).

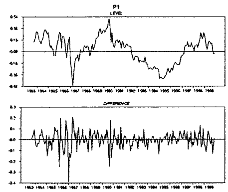

Figure 27 p_t (P1).

⊂ 4 ⊃ **Testing the adequacy of the (unrestricted) model (6)**
Note that the (unrestricted) VAR model (6) is equivalent to (7) with $r = 3$ (i.e., $\mathbf{\Pi}$ of full rank). What follows Table 12 is a set of various time-series plots of three (potentially) endogenous variables in $\boldsymbol{y}_t$. Each of Figures 25 through 27 draws a pair of a (logged) time series and its first difference (monthly growth rate). Each of Figures 28 through 30 draws a set of four plots: the actual and fitted difference, the standardized residuals, a histogram of the standardized residuals with the standard Normality histogram, and the residuals correlogram. Figure 31 plots the residuals cross-correlogram, along with their correlogram as drawn in Figures 28 through 30. (For the notation used in these figures, see Table 1.)

elements. Moreover, the number of roots close to unity represents the number of common stochastic trends (Hansen and Juselius, p.28).

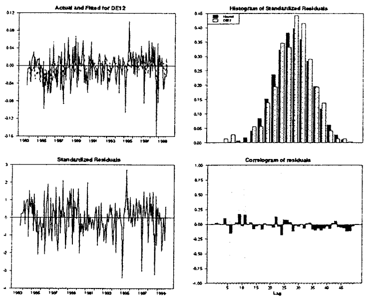

Figure 28 Δs_t (DE12). Those autocorrelations outside the confidence bands $\pm 2/\sqrt{T}$ are shaded.

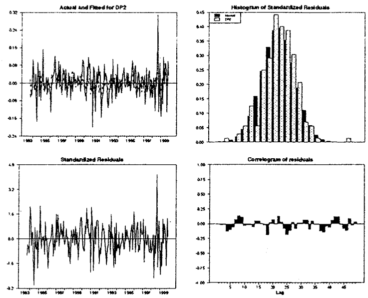

Figure 29 Δp_t^* (DP2).

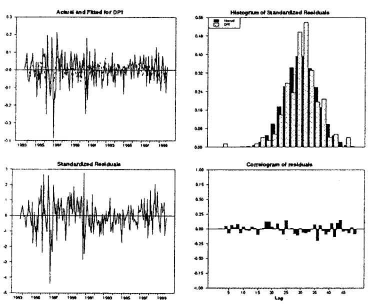

Figure 30 Δp_t (DP1).

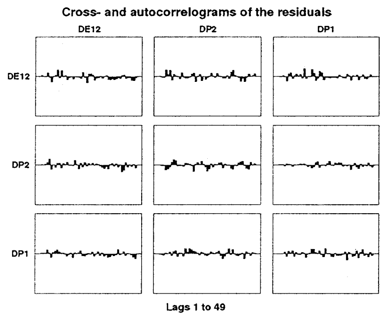

Figure 31 Cross- and Autocorrelograms for $\Delta s_t, \Delta p_t^*, \Delta p_t$ (DE12, DP2, DP1). See Hansen and Juselius (p.22) for the formula for cross- and autocorrelograms.

Another set of graphs for testing the adequacy of the model comprises Figures 32 through 34 plotting the 3 cointegration relations/eigenvectors

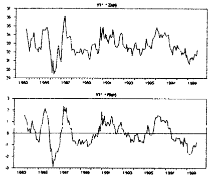

Figure 32 s_t.

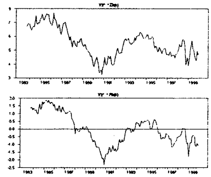

Figure 33 p_t^*.

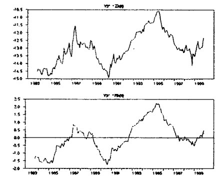

Figure 34 p_t.

without any restrictions. How to interpret the figures is as follows (Hansen and Juselius, pp.22-23): The upper half "pictures the actual disequilibrium as function of all short-run dynamics including seasonal and other dummies" and the lower half "is corrected for the short-run effects, and pictures the 'clean' disequilibrium. It is the series in the lower graph that is actually tested for stationarity and thus determines r in the maximum likelihood procedure. In case both look radically different, in particular, if the upper looks $I(1)$ whereas the latter looks stationary, it is a good idea to check whether your data vector is second order instead of first order nonstationary."

Since Figures 32 through 34 show that no such radical differences are present, the order of integration of the variables is first, which will be found consistent with the unit-root test results later in Table 15.

Still another model adequacy check is a battery of residual diagnostic checks for the (unrestricted) model (7) with $r = 3$: see Table 13. In the table, a prefix 'Δ' is attached to reflect that the model is in error-correction form (7), here still with $\mathbf{\Pi}$ of full rank. Later in Table 21, the residual analysis for the *restricted* model (7) with $r = 1$ (i.e., $\mathbf{\Pi}$ of less than full rank) will be seen to also lead to almost the same results as those in Table 13 here for the unrestricted model. See section 4 for the interpretation of the results.

$\subset$ 5 $\supset$ **Preliminary analysis for the restricted model: Setting** $r = 1$ From Tables 11 and 12 it was inferred that r=1; hence, $\boldsymbol{\beta}$ is a column vector. Setting now $r = 1$ and normalizing on element 1 (s_t) of the cointegration vector, we will get the estimated results in Table 14; note that, at this point, these parameters are still *not* under any imposed restrictions.

Table 13 Cointegration Analysis [3]:[a] Residual Analysis; Unrestricted Model (7) with $r = 3$

Correlation Matrix						
	Δs_t	Δp^*_t	Δp_t			
	1.000					
	-0.168	1.000				
	0.123	-0.003	1.000			
Standard Deviations ofResiduals	0.032	0.060	0.072			
Multivariate Statistics[b]						
Log (Det (Sigma))	-17.769					
Information Criteria: SC	-15.830					
HQ	-16.547					
Trace Correlation	0.178					
Test for Autocorrelation[c]						
L-B(49), χ^2 (405)	401.661	(0.54)[d]				
LM(1), χ^2 (9)	9.447	(0.40)				
LM(4), χ^2 (9)	14.377	(0.11)				
Test for Normality						
χ^2 (6)	41.473	(0.00)				
Univariate Statistics[e]						
	Mean	Std. Dev.	Skewness	Kurtosis[f]	Maximum	Minimum
	0.000	0.032	-0.437	3.718	0.089	-0.111
	0.000	0.060	0.066	4.167	0.254	-0.182
	-0.000	0.072	-0.421	4.925	0.201	-0.320
	ARCH(4)[g]	Normality[h]	R^2			
	2.169	7.109	0.162			
	8.631	11.906	0.189			
	24.603	21.015	0.177			

[a] A prefix 'Δ' is attached to reflect that the model is in error-correction form (7), here with $\mathbf{\Pi}$ of full rank.

[b] See Hansen and Juselius (pp.26-27).

[c] L-B stands for Ljung-Box, and LM Lagrange Multiplier.

[d] P-value.

[e] Computed are the statistics for the residuals from each equation; see Hansen and Juselius (p.27).

[f] This is not an excess kurtosis.

[g] Presented is an univariate LM-type test for ARCH of order 4 in the residuals.

[h] The modified version of the Shenton-Bowman test of normality is reported.

Panel "$\boldsymbol{\alpha}$" in Table 14 indicates that α_{11} for Δs_t is not statistically significant, suggesting the weak exogeneity of s_t. A restriction is then to be imposed on $\boldsymbol{\alpha}$ and tested; this hypothesis testing will be carried out subsequently, after testing for the restriction on $\boldsymbol{\beta}$. (The order of testing does not, of course, affect the results in any way.)

3.2 Multivariate cointegration testing of PPP

With all these preliminary results in the preceding section, we are now ready to test for unit roots of each of the three variables by the Johansen procedure.

Table 14 Cointegration Analysis [4]: Estimated Results of Unrestricted Model (7) with $r = 1$

Re-normalisation of the eigenvectors			
Eigenvector(s) (transposed)	s_t	p_t^*	p_t
	6.328	5.921	-6.590
The matrices based on 1 cointegration vectors			
β'	s_t	p_t^*	p_t
	1.000	0.936	-1.042
α			
Δs_t	-0.005		
Δp_t^*	-0.091		
Δp_t	0.102		
T-values for α			
	-0.349		
	-3.320		
	3.116		
$\boldsymbol{\Pi}$	s_t	p_t^*	p_t
Δs_t	-0.005	-0.005	0.005
Δp_t^*	-0.091	-0.085	0.095
Δp_t	0.102	0.096	-0.107
T-values for $\boldsymbol{\Pi}$			
	-0.349	-0.349	0.349
	-3.320	-3.320	3.320
	3.116	3.116	-3.116

3.2.1 Testing for unit roots with the Johansen procedure

The null hypothesis to be tested in the Johansen procedure is the *stationarity* of a variable, not its nonstationarity. The stationarity null for the ith variable is formulated as the null that e_i, whose elements are all zero except for the ith element being unity, is in the cointegrating space:[28]

$$\boldsymbol{\beta} = (\mathbf{H}_i, \varphi) \tag{10}$$

where a $(3 \times r_1)$ matrix $\mathbf{H}_i = \boldsymbol{e}_i$, to test for a unit root for the ith variable, and φ is a $(3\times r_2)$ matrix, with $r = r_1 + r_2$ (Harris, p.106). In

[28] For $r = 1 = i$, for example, in eq. (7), $\boldsymbol{\Pi y}_{t-1} = \boldsymbol{\alpha\beta' y}_{t-1} = \begin{bmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{bmatrix} \begin{pmatrix} 1 & 0 & 0 \end{pmatrix} \begin{bmatrix} s_{t-1} \\ p_{t-1}^* \\ p_{t-1} \end{bmatrix} = \begin{bmatrix} \alpha_{11} \\ \alpha_{21} \\ \alpha_{31} \end{bmatrix} s_{t-1}$. Notice that eq. (7) is indeed seen as a multivariate form of the univariate ADF regression equation in Table 6 and thus the Johansen approach is essentially "a particular type of unit-root test using a multivariate form of the ADF test with the null of stationarity" (Harris, p.107).

our case, $r = r_1 = 1$ with $r_2 = 0$, for there exists only one cointegrating vector:

$$\mathbf{H}_0 : \boldsymbol{\beta} = (\mathbf{H}_i). \tag{11}$$

Table 15 Cointegration Analysis [5]:[a] The Johansen Approach to Testing for Unit Roots; Unrestricted Model (7) with $r = 1$

In $\mathbf{H}_0$ in eq. (11):	$i = 1\,(s_t)$			$i = 2\,(p_t^*)$			$i = 3\,(p_t)$		
Re-normalisation of the eigenvectors									
Eigenvector(s) (transposed)									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
	3.696	0.000	0.000	0.000	4.698	0.000	0.000	0.000	5.058
The LR test:[b]									
χ^2 (2)	17.08 (0.00)[c]			14.94 (0.00)			16.06 (0.00)		
$\boldsymbol{\beta}'$									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
	1.000	0.000	0.000	0.000	1.000	0.000	0.000	0.000	1.000
$\boldsymbol{\alpha}$									
Δs_t	-0.017			0.020			-0.007		
Δp_t^*	-0.004			-0.043			0.004		
Δp_t	-0.003			0.007			-0.061		
T-values for $\boldsymbol{\alpha}$									
	-2.016			1.820			-0.609		
	-0.250			-2.065			0.174		
	-0.170			0.290			-2.308		
$\boldsymbol{\Pi}$									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
Δs_t	-0.017	0.000	0.000	0.000	0.020	0.000	0.000	0.000	-0.007
Δp_t^*	-0.004	0.000	0.000	0.000	-0.043	0.000	0.000	0.000	0.004
Δp_t	-0.003	0.000	0.000	0.000	0.007	0.000	0.000	0.000	-0.061
T-values for $\boldsymbol{\Pi}$									
	-2.016	NA	NA	NA	1.820	NA	NA	NA	-0.609
	-0.250	NA	NA	NA	-2.065	NA	NA	NA	0.174
	-0.170	NA	NA	NA	0.290	NA	NA	NA	-2.308

[a] Neither $\boldsymbol{\beta}$ nor $\boldsymbol{\alpha}$ is under any restrictions yet.

[b] The null of stationarity of each (potentially) endogenous variable is tested.

[c] P-value.

Tests of a null of type $\mathbf{H}_0$ are performed for each variable i, whose likelihood ratio (LR) test results, with degrees of freedom equal to $(3 - r)r_1 = (3-1)1 = 2$, are as reported in Table 15, in which no restrictions are yet imposed on either $\boldsymbol{\beta}$ or $\boldsymbol{\alpha}$: as desired, the null of stationarity is rejected strongly at any conventional level of significance, for every variable. This is consistent with the earlier observation that, for Japan, all the roots of the companion matrix are inside the unit circle, which is in turn consistent with each variable being $I(1)$. (Moreover, the result is consistent with the unit-root tests in Step 1 of the earlier Engle-Granger single-equation two-step cointegration test.)

We are now ready to go on to testing for restrictions on cointegration relation(s) $\boldsymbol{\beta}$ and on $\boldsymbol{\alpha}$ in the VEC system (7).

3.2.2 Restricting cointegration relation $\boldsymbol{\beta}$ and the speed of adjustment $\boldsymbol{\alpha}$

Restricting only the cointegrating vector(s) in $\boldsymbol{\beta}$ We first test for linear hypotheses on cointegration relation(s). Since we only have ($r =$) one cointegration vector, no identification problem occurs[29] and the (transposed) vector

$$\boldsymbol{\beta}' = \begin{pmatrix} \beta_{11} & \beta_{21} & \beta_{31} \end{pmatrix}. \tag{12}$$

A general type of restriction we impose here is

$$\mathbf{R}'\boldsymbol{\beta} = \mathbf{0}, \tag{13}$$

and a particular restriction motivated by our economic arguments is the (strong) *PPP restriction* on $\boldsymbol{\beta}'$ in (12)

$$\begin{pmatrix} 1 & 1 & -1 \end{pmatrix}.^{30} \tag{14}$$

This is in fact a vector of coefficients on the right-hand side of the real exchange rate's definition (in section 1) and implies two homogeneity restrictions:

$$\beta_{11} + \beta_{31} = 0 \tag{15}$$

$$\beta_{21} + \beta_{31} = 0. \tag{16}$$

Thus, in eq. (13),

$$\mathbf{R}' = \begin{bmatrix} 1 & 0 & 1 \\ 0 & 1 & 1 \end{bmatrix}. \tag{17}$$

[29] When $r > 1$, any linear combination of two cointegration relations will preserve the stationarity property, in which case one must go on to test whether the vectors of $\boldsymbol{\beta}$ are identified. See, for example, Harris (p.104) and Hansen and Juselius (p.3).

[30] Weaker PPP restrictions are as follows (MacDonald and Marsh, Table 4, p.35):
$\begin{pmatrix} 1 & \beta_{21} & -\beta_{21} \end{pmatrix}$ where β_{21} is free with its sign being constrained and, in eq. (13), one-row $\mathbf{R}' = \begin{pmatrix} 0 & 1 & 1 \end{pmatrix}$ (with element 1, s_t, being normalized on);
$\begin{pmatrix} 1 & \beta_{21} & -1 \end{pmatrix}$ where β_{21} is free and, in eq. (13), one-row $\mathbf{R}' = \begin{pmatrix} 1 & 0 & 1 \end{pmatrix}$; and
$\begin{pmatrix} 1 & 1 & \beta_{31} \end{pmatrix}$ where β_{31} is free and, in eq. (13), one-row $\mathbf{R}' = \begin{pmatrix} 1 & -1 & 0 \end{pmatrix}$.

Note that normalizing on β_{11} leads readily to the PPP restriction (14) if β' in (12) has either of signs (+ + −) and (− − +).

Under column "Restricting Only β" in Table 16, note the italic figures which indicate (i) two homogeneity restrictions (15) and (16) implied by the (strong) PPP restriction on $\boldsymbol{\beta}$, in panel "Eigenvector(s) (transposed)," and (ii) normalizing on the first element of the cointegration vector (12), in panel "$\boldsymbol{\beta}'$."

Table 16 Cointegration Analysis [6]: Hypothesis Testing of the Restrictions; Restricted Model (7) with $r = 1$

	Restricting Only β (PPP Restriction)[a]			Restricting Only α			Restricting Both		
Re-normalisation of the eigenvectors									
Eigenvector(s) (transposed)									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
	-6.262	*-6.262*	*6.262*	6.242	6.001	-6.564	*-6.262*	*-6.262*	*6.262*
The LR test:									
χ^2 (2)[b]	0.16 (0.92)[c]			0.10 (0.76)			0.21 (0.98)		
$\boldsymbol{\beta}'$									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
	1.000	*1.000*	*-1.000*	1.000	0.961	-1.052	*1.000*	*1.000*	*-1.000*
$\boldsymbol{\alpha}$									
Δs_t	-0.003			*0.000*			*0.000*		
Δp_t^*	-0.094			-0.092			-0.095		
Δp_t	0.097			0.102			0.098		
T-values for $\boldsymbol{\alpha}$									
	-0.222			0.000			0.000		
	-3.464			-3.397			-3.502		
	2.983			3.145			3.010		
$\boldsymbol{\Pi}$									
	s_t	p_t^*	p_t	s_t	p_t^*	p_t	s_t	p_t^*	p_t
Δs_t	-0.003	-0.003	0.003	0.000	0.000	0.000	0.000	0.000	0.000
Δp_t^*	-0.094	-0.094	0.094	-0.092	-0.088	0.097	-0.095	-0.095	0.095
Δp_t	0.097	0.097	-0.097	0.102	0.098	-0.107	0.098	0.098	-0.098
T-values for $\boldsymbol{\Pi}$									
	-0.222	-0.222	0.222	NA	NA	NA	NA	NA	NA
	-3.464	-3.464	3.464	-3.448	-3.448	3.448	-3.555	-3.555	3.555
	2.983	2.983	-2.983	3.168	3.168	-3.168	3.031	3.031	-3.031

[a] See the italic figures under this column; the same applies to the remaining two restrictions.

[b] Degrees of freedom= $(3 - r)r_1 = (3 - 1)1 = 2$, where $r = r_1 + r_2$ with $r_2 = 0$ (Hansen and Juselius, p.40).

[c] P-value.

The table reports the LR test results for the null hypothesis of the PPP restriction: the null of the PPP restriction is not rejected (or is readily accepted). The VEC model (7) under the PPP restriction which yields *the equilibrium error in the form of a real exchange rate* may now be formally called a *PPP-based* system of exchange rate and prices. C-R-X's estimated inflation rates have thus been shown to have desired theoretical content, here in the VEC model of the PPP relationship; this adds to C-R-X's evidence of the same theoretical content documented in their single equation-based analysis of the PPP relationship.

Restricting only the speed of adjustment $\boldsymbol{\alpha}$ Table 14 at the end of the preliminary analysis shows that α_{11} for the first element s_t is not statistically significant, suggesting the weak exogeneity of s_t. A restriction is now imposed on $\boldsymbol{\alpha}$ and tested, without imposing any restrictions on $\boldsymbol{\beta}$.

For one restriction on the first element, one-row

$$\mathbf{B}' = (\ 1 \quad 0 \quad 0\ ), \tag{18}$$

and

$$\mathbf{B}'\boldsymbol{\alpha} = \mathbf{0}.^{31} \tag{19}$$

Under column "Restricting Only α" in Table 16, note (i) no restrictions on $\boldsymbol{\beta}$, in panel "Eigenvector(s) (transposed)," (ii) normalizing on the first element of the cointegration vector (12), in panel "$\boldsymbol{\beta}'$," and (iii) the italic figures indicating the zero restriction on the first element of $\boldsymbol{\alpha}$, in panel "$\boldsymbol{\alpha}$." The LR test results reported in the table show that the null of the weak exogeneity is not rejected. We will later estimate the conditional version (9), which is conditioned on the weakly exogenous variable s_t.[32]

Restricting, jointly, $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ Under column "Restricting Both" in Table 16, notice the italic figures indicating (i) two homogeneity restrictions (15) and (16) implied by the (strong) PPP restriction on $\boldsymbol{\beta}$, in panel "Eigenvector(s) (transposed)," (ii) normalizing on the first element of the cointegration vector (12), in panel "$\boldsymbol{\beta}'$," and (iii) a zero restriction on the first element of $\boldsymbol{\alpha}$, in panel "$\boldsymbol{\alpha}$." The LR test results in the table show that the null of the joint restrictions on $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ is not rejected (or is readily accepted).

---

[31] For a restriction on each of two elements 1 and 3, e.g., we would have a two-row

$$\mathbf{B}' = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}.$$

[32] Harris (pp.103-104) presents similar weak-exogeneity evidence for the U.K. (effective) exchange rate in the PPP and UIP model of Johansen and Juselius (1992).

Shown in Table 17 is a list of the eigenvalues of the companion matrix computed under the joint restrictions: two unit roots are obtained; this is, as desired, consistent with two eigenvalues close to unity under column $L = 4$ in Table 12 constructed for a preliminary purpose.

Table 17 Cointegration Analysis [7]: The Eigenvalues of the Companion Matrix; (Jointly) Restricted Model (7) with $r = 1$

real	complex	modulus	argument
1.000	-0.000	1.000	-0.000
1.000	0.000	1.000	0.000
0.722	0.144	0.736	0.196
0.722	-0.144	0.736	-0.196
0.357	0.350	0.500	0.775
0.357	-0.350	0.500	-0.775
0.082	0.489	0.496	1.404
0.082	-0.489	0.496	-1.404
-0.259	0.408	0.483	2.137
-0.259	-0.408	0.483	-2.137
-0.436	-0.163	0.466	-2.784
-0.436	0.163	0.466	2.784

3.2.3 Summary: Long-run structure, real exchange rate, weak exogeneity, and causality direction

Long-run structure and real exchange rate The estimated results under column "Restricting Both" in Table 16 strongly suggest the following features of the long-run structure of the yen per dollar rate and the prices: in the context of PPP-based VEC model (7) with $y_t = (s_t, p_t^*, p_t)'$, there is *not* observed any long-run cointegration relation in the short-run equation for Δs_t, while there is for two other (price change) equations.

That is, the short-run change in the nominal exchange rate Δs_t will *not* adjust to the equilibrium error (i.e., a previous real exchange rate $s_{t-1} + p_{t-1}^* - p_{t-1}$), while two short-run price changes Δp_t^* and Δp_t do adjust, respectively, in a negative and positive direction, but in almost the same speed, to the previous real exchange rate.

Weak exogeneity and causality direction s_t is found weakly exogenous to the system of equations under study. Does the weak exogeneity of s_t make sense in the PPP context? Ito (2005) argues that the direction of causality *from prices to exchange rates* is consistent with the PPP hypothesis. Also, in a cointegrated system, $\{y_t\}$ (e.g., price) does not Granger cause $\{z_t\}$ (e.g., exchange rate) if (i) lagged values Δy_{t-l} do not enter the short-run Δz_t equation *and* if (ii) $\{z_t\}$ is weakly exogenous to the system of the equations (i.e., does not respond to the deviation from long-run equilibrium) (Enders, p.334). As will be shown in the next section,[33] in our present empirical results, (i) does not apply, although (ii) does; therefore, no such Granger *non*-causality from prices to exchange rate seems to be detected.

The weak-exogeneity evidence of the yen per dollar rate alone would suggest no direction of causality. We do not test for the Granger non-causality null for our cointegrated system, however; rather, impulse response functions of exchange rate and prices will be computed and studied within the short-run structure of the cointegrated system, in the subsequent section.

4 Short-run Structure and Conditional Model

We now accept the long-run structure as estimated under column "Restricting Both" (including the strong PPP restriction) in Table 16, as a reasonable yen against dollar behavior during the sample time period; this in turn means that "the modeling of the long-run structure is done." With the long-run structure on hand, we now move on to the short-run study (Hansen and Juselius, p.49). Later, we will turn to the conditional version of the model, to attempt another multivariate PPP analysis based on eq. (9), assuming explicitly the weak exogeneity of the exchange rate.

4.1 Short-run structure

4.1.1 The complete PPP-based VEC model and the short-run effects

With the long-run structure as estimated under column "Restricting Both" in Table 16 in the preceding section, the short-run matrices are estimated and shown in eq. (20) which is a numerically written, complete PPP-based VEC model (7) with (8) being substituted. The correspond-

[33]See the Δs_t equation in eq. (20) or the Δs_t row in its summary Table 19 there.

ing t-values are reported in Table 18,[34] and the statisitically significant short-run effects as asterisked in (20) are summarized in Table 19.[35]

Table 18 Cointegration Analysis [8]: T-values of the Estimated Short-run Matrices and the Deterministic Variables in Eq. (20)

The short-run matrices Φ_l^{Δ}:

Time: $t-1$			Time: $t-2$			Time: $t-3$		
0.379	-2.921	0.014	2.183	0.570	0.319	0.315	0.125	-2.825
0.745	0.907	-1.377	0.825	-0.754	-0.275	1.378	0.612	-0.941
0.084	-2.212	0.452	0.144	-1.510	-1.017	0.145	-1.180	-0.734

Ψ for the deterministic variables:

SEA(1)	SEA(2)	SEA(3)	SEA(4)	SEA(5)	SEA(6)	SEA(7)	SEA(8)	SEA(9)
-1.086	-1.482	-0.673	-0.533	0.110	-1.387	-0.671	-1.195	-0.331
0.189	-2.388	-0.925	-0.871	-2.348	-2.634	-1.051	-2.713	-0.555
1.081	0.775	0.739	1.642	-0.408	-0.420	-1.295	1.291	0.664

SEA(10)	SEA(11)	CONST
-0.751	-1.007	-0.044
-2.418	-0.648	3.518
1.550	0.831	-3.004

Table 19 Summary Results: Statistically Significant Effects in Eq. (20)[a]

Short-run equation	Short-run effects of[b]	Long-run cointegration relation
Δs_t	$\Delta s_{t-2}(+)^{***}$, $\Delta p^*_{t-1}(-)^{***}$, $\Delta p_{t-3}(-)^{***}$	Not observed
Δp^*_t	$\Delta s_{t-3}(+)^{*}$, $\Delta p_{t-1}(-)^{*}$; large but insignificant[c] $\Delta s_{t-2}(+)$	Observed
Δp_t	$\Delta p^*_{t-1}(-)^{***}$, $\Delta p^*_{t-2}(-)^{*}$; large but insignificant $\Delta p^*_{t-3}(-)$	Observed

[a]*** and * (attached to the right paranthesis) denote significance at 1 and 10% levels, respectively. The sign in parantheses indicates the sign of the effect.

[b]The statistically significant "short-run effects" here coincide, in sign though not in magnitude, with those kinks in the impulse response functions observed in Figure 36: the positive and negative short-run effects detected there are graphically displayed, respectively, as upward and downward kinks in the impulse response functions; for example, $\Delta s_{t-2}(+)$ in the short-run Δs_t equation corresponds to the upward kink in the top left impulse response function in Figure 36.

[c]See the underlined figures in eq. (20).

[34]Hansen and Juselius' *CATS in RATS* does not compute p-values here.

[35]The statistical significance for $\boldsymbol{\alpha}$ is reported under column "Restricting Both" in Table 16; recall that the (strong) PPP restriction (14) is imposed on $\boldsymbol{\beta}'$.

$$\begin{aligned}
\begin{bmatrix}\Delta s_t\\ \Delta p_t^*\\ \Delta p_t\end{bmatrix} &= \begin{bmatrix} 0.028 & -0.112^{***} & 0.000\\ 0.100 & 0.064 & -0.082^{*}\\ 0.014 & -0.189^{***} & 0.033\end{bmatrix}\begin{bmatrix}\Delta s_{t-1}\\ \Delta p^*_{t-1}\\ \Delta p_{t-1}\end{bmatrix}\\
&+\begin{bmatrix} 0.157^{***} & 0.023 & 0.010\\ \underline{0.109} & -0.055 & -0.016\\ 0.023 & -0.133^{*} & -0.073\end{bmatrix}\begin{bmatrix}\Delta s_{t-2}\\ \Delta p^*_{t-2}\\ \Delta p_{t-2}\end{bmatrix}\\
&+\begin{bmatrix} 0.022 & 0.005 & -0.091^{***}\\ 0.180^{*} & 0.045 & -0.056\\ 0.023 & \underline{-0.105} & -0.053\end{bmatrix}\begin{bmatrix}\Delta s_{t-3}\\ \Delta p^*_{t-3}\\ \Delta p_{t-3}\end{bmatrix}\\
&+\begin{bmatrix} 0.000\\ -0.095^{***}\\ 0.098^{***}\end{bmatrix}\begin{pmatrix} 1 & 1 & -1\end{pmatrix}\begin{bmatrix} s_{t-1}\\ p^*_{t-1}\\ p_{t-1}\end{bmatrix}+\begin{bmatrix} -0.003\\ 0.497^{***}\\ -0.510^{***}\end{bmatrix}\\
&+\begin{bmatrix} -0.013 & -0.018^{*} & -0.008 & -0.006 & 0.001 & -0.016 & -0.008 & -0.014 & -0.004 & -0.009 & -0.012\\ 0.004 & -0.052^{***} & -0.020 & -0.019 & -0.052^{***} & -0.058^{***} & -0.023 & -0.061^{***} & -0.013 & -0.053^{***} & -0.014\\ 0.028 & 0.020 & 0.019 & 0.042^{*} & -0.011 & -0.011 & -0.034^{*} & 0.035^{*} & 0.018 & 0.041^{*} & 0.022\end{bmatrix} \boldsymbol{sdum} \qquad (20)
\end{aligned}$$

where: *** and * denote significance at 1 and 10% levels, respectively; the underlined figures are large in absolute value but statistically insignificant (see Tables 18 and 19); and the column vector

$$\boldsymbol{sdum} = \begin{pmatrix} sdum_2 & sdum_3 & sdum_4 & sdum_5 & sdum_6 & sdum_7 & sdum_8 & sdum_9 & sdum_{10} & sdum_{11} & sdum_{12}\end{pmatrix}'. \qquad (21)$$

The short-run effects As also seen from Table 19, the short-run Δs_t equation in the estimated system (20) contains *strongly* statistically significant short-run effects of $\Delta s_{t-2}(+)$,[36] $\Delta p^*_{t-1}(-)$ and $\Delta p_{t-3}(-)$, while s_t itself is weakly exogenous to the system. Then the responses of exchange rate to prices appear to be detected here.

Notice, too, that neither short-run effects of one-month lagged Δs_{t-1} and Δp_{t-1} are found significant in the Δs_t equation.[37] This may appear puzzling, and some refinement of the model could be required. Instead, here, we estimate a univariate, 3rd-order autoregressive model, AR[3], to check if Δs_{t-1} would become statistically significant in a simple time series model:

$$\Delta s_t = c + \sum_{i=1}^{3} \phi_i \Delta s_{t-i} + u_t \qquad (22)$$

[36]The sign in parantheses indicates the sign of the effect.

[37]A similar observation (with regard to Δs_{t-1}, in particular) is documented in MacDonald and Marsh (Table 5, pp.36-39) for the U.K. and Germany.

Table 20 Box-Jenkins - Estimation by Gauss-Newton: Two AR[3] Models

Dependent Variable	Eq. (22): Δs_t	Eq. (23): s_t
Constant	-0.005 (0.131)[a]	4.702 (0.000)
Δs_{t-1}	0.037 (0.613)	
Δs_{t-2}	0.134 (0.062)	
Δs_{t-3}	-0.018 (0.808)	
s_{t-1}		1.020 (0.000)
s_{t-2}		0.010 (0.334)
s_{t-3}		-0.135 (0.057)
Monthly Data	1983:09 To 1999:12	1983:08 To 1999:12
Usable Observations	196	197
Degrees of Freedom	192	193
Adjusted R^2	0.983	0.983
Residual Standard Deviation	0.035	0.035
Regression F(3,192); F(3,193)	3704.201 (0.000)	3873.769 (0.000)
Durbin-Watson Statistic	1.984	1.998
Q(36-3)	29.595 (0.637)	30.369 (0.599)
Ljung-Box Q-Statistics for Residuals SACF[b]	17.607 (0.347)	18.334 (0.305)

[a]P-value.

[b]Sample autocorrelation function.

where c and u_t are, respectively, a constant and a white noise. As readily seen from Table 20, Δs_{t-1} is still not statistically significant.[38] The statistical *in*significance of Δs_{t-1} (and Δp_{t-1} as well) could be then due to the estimated VEC model (20) consisting of *growth-rate* (or first-difference) equations rather than level equations: the statistical sig-

[38]Examining the regression (22) with a GARCH error process, Ito (2005, Table 2, pp.12-13) also obtains similar evidence for the sample period from April 1991 through December 2003 that *none* of the three autoregressive parameters are statistically significant for monthly data. His result indeed evidences the random-walk nature of the nominal exchange rate level s_t. (See MacDonald and Marsh, pp.25-28 for three approaches to testing a random walk hypothesis, one of which estimates such a model as (22).)

From Table 6 we already infer that s_t itself has a unit root. Only for reference purposes, however, Table 20 also reports the estimated AR[3] model for s_t

$$s_t = c + \sum_{i=1}^{3} \phi_i s_{t-i} + u_t; \tag{23}$$

the table shows that s_{t-1} is statistically significant with its estimated ϕ_1 being slightly greater than unity, which violates the sationarity condition of the AR model (and thus is consistent with the non-rejection of the unit-root null for s_t in Table 6). The random walk hypothesis is studied by Kojima (2006c).

nificance/insignificance in the model should be interpreted in terms of growth rate. We may thus take the statistical insignificance of Δs_{t-1} in VEC model (20) to be reasonable in the growth-rate form and consistent with the widely documented random-walk nature of s_t.

The short-run Δp_t^* (U.S. price change) equation contains weakly statistically significant short-run effects of $\Delta s_{t-3}(+)$ and $\Delta p_{t-1}(-)$, along with large, though insignificant, effects from the previous change in exchange rate $\Delta s_{t-2}(+)$: some degree of exchange rate effect and response of U.S. price to Japanese price appear to be detected here.

On the other hand, the short-run Δp_t (Japanese price change) equation contains statistically significant short-run effects only of $\Delta p_{t-l}^*(-)$, $l = 1, 2$: no exchange rate effects but only the responses of Japanese price to U.S. price appear to be present.

The analysis here is only preliminary; those positive and negative short-run effects detected here will be graphically displayed, respectively, as upward and downward kinks in the impulse response functions, later in section 5.

Another set of short-run effects comprises seasonal dummies (Harris, p.83): notice in eq. (20) that both price change equations contain statistically significant effects of seasonal dummies: several, strongly statistically significant effects are observed in the Δp_t^* equation, while weakly statistically significant effects in the Δp_t equation.

4.1.2 Residual analysis for the jointly restricted model

Restricted model Shown in Table 21 is the residual analysis for the jointly restricted model (20).[39] The tests in Table 21 may be divided into three parts, (i) Correlation Matrix and Standard Deviations of Residuals, (ii) Multivariate Statistics and (iii) Univariate Statistics:[40]

In (i), the (contemporaneous) cross-correlations of the residuals are reasonably small. (They need not to be small, however, since, for our underlying VAR model (6), the white noise $\boldsymbol{u}_t \sim IN(\mathbf{0}, \Sigma)$ and the dependence is allowed among the white-noise disturbance terms $u_{1t_1}, ..., u_{nt_n}$ for any t_i, $i = 1, ..., n$.)

(ii) contains some 'goodness of fit' measures in the multivariate con-

[39]The residual analysis for the unrestricted model (7) with $r = 3$ is provided in Table 13, to which we will turn later.

[40]For the brief description of the associated formulas, see Hansen and Juselius (pp.26-27, 72-76).

Table 21 Cointegration Analysis [9] (Continued from Tables 18 and 19):[a] Residual Analysis; Jointly Restricted Model (20)

Correlation Matrix						
	Δs_t	Δp^*_t	Δp_t			
	1.000					
	-0.172	1.000				
	0.119	0.007	1.000			
Standard Deviations of Residuals	0.033	0.061	0.073			
Multivariate Statistics						
Log (Det (Sigma))	-17.723					
Information Criteria: SC	-15.892					
HQ	-16.569					
Trace Correlation	0.164					
Test for Autocorrelation						
L-B(49), χ^2 (411)	409.524	(0.51)[b]				
LM(1), χ^2 (9)	8.277	(0.51)				
LM(4), χ^2 (9)	14.615	(0.10)				
Test for Normality						
χ^2 (6)	39.059	(0.00)				
Univariate Statistics						
	Mean	Std. Dev.	Skewness	Kurtosis	Maximum	Minimum
	-0.000	0.033	-0.418	3.599	0.093	-0.108
	0.000	0.061	0.013	4.169	0.252	-0.179
	-0.000	0.073	-0.468	4.862	0.200	-0.314
	ARCH(4)	Normality	R^2			
	1.743	6.267	0.141			
	10.707	12.002	0.183			
	26.228	19.269	0.164			

[a]See notes in Table 13.

[b]P-value.

text, to determine whether the residuals of the error-correction equaitons approximate white noise. The null of no autocorrelations is not rejected, while the null of normality is. If the residuals were serially correlated, lag lengths L might be too short. That is not the case here; $L = 4$ has been found long enough a lag length.

In (iii), a univariate test for ARCH of order 4 ($=L$) in the residuals is included. R^2 and the coefficients of determination are all less than 20% but greater than 14%. The small magnitude of the coefficients of determination is reasonable, for the growth-rate variables entering the short-run equations are stationary.

Restricted versus unrestricted models Does the restricted model (with $r = 1$, i.e., $\mathbf{\Pi}$ of reduced rank.) here in Table 21 have better in-sample properties, as compared to the unrestricted model (with $r = 3$, i.e., $\mathbf{\Pi}$ of full rank.) earlier in Table 13? Contrasting the residual analyses for the two models, one readily sees that (i), (ii) and (iii) remain almost the same. This would imply that restricting itself would not improve (or worsen) the residuals behavior.

4.2 Conditional version of the model

With the weak exogeneity of s_t, we now turn to and estimate eq. (9), a conditional version of the model, where $z_t = (p^*_t, p_t)'$ and $y_t = (p^*_t, p_t, s_t)$. The estimated results are shown in Tables 22 through 25.

Table 22 Cointegration Analysis [10a]: Conditional, Unrestricted Model (9)

Endogeneous series	p^*_t p_t					
Exogeneous series	Non stationary : s_t					
Deterministic series	Unrestricted constant and trend in coint. space; 11 centered seasonal dummies					
Effective sample	1983:09 TO 1999:12					
Lag(s) in VAR-model	4					
No. of observations	196					
Obs.- no.of variables	171					
$I(1)$ Analysis						
Cointegration Rank Tests	Eigenv.	L-max	Trace	H_0: r	$p-r$	
	0.103	21.35	25.57	0	2	
	0.021	4.21	4.21	1	1	
β'	p^*_t	p_t	s_t			
	-6.001	6.564	-6.242			
	3.445	3.546	0.743			
α						
p^*_t	0.015	-0.006				
p_t	-0.016	-0.008				
Π	p^*_t	p_t	s_t			
	-0.109	0.075	-0.096			
	0.071	-0.135	0.096			
Residual Analysis:[a]						
Correlation Matrix	Δp^*_t	Δp_t				
	1.000					
	0.018	1.000				
Standard Deviations of Residuals	0.059	0.072				
Multivariate Statistics						
Log (Det (Sigma))	-10.911					
Information Criteria: SC	-9.565					
HQ	-10.062					
Trace Correlation	0.199					
Test for Autocorrelation						
L-B(49), χ^2 (180)	157.541	(0.89)[b]				
LM(1), χ^2 (4)	4.579	(0.33)				
LM(4), χ^2 (4)	2.831	(0.59)				
Test for Normality						
χ^2 (4)	31.407	(0.00)				
Univariate Statistics	Mean	Std. Dev.	Skewness	Kurtosis	Maximum	Minimum
	-0.0000	0.059	0.015	4.111	0.245	-0.172
	-0.0000	0.072	-0.382	4.850	0.211	-0.315
	ARCH(4)	Normality	R^2			
	13.123	11.098	0.212			
	22.894	20.433	0.189			

[a]See notes in Table 13.

[b]P-value.

Skipping the residual analysis for the unrestricted model in Table 22 and looking at the LR test result of the PPP-restriction null in Table 23,[41] one cannot reject, or can readily accept, the null. This is exactly the same decision we reached for the (unconditional) model (7).

Also, the short-run structure as well remains unchanged, except that

[41]The PPP-restriction null is shown in italic figures in panel "β'" of the table.

there is no longer a Δs_t equation; as is clear from eq. (9), Δs_t appears as a separate term in each endogenous variable equation, whose short-run structure is shown separately in Table 24.

Table 23 Cointegration Analysis [10b] (Continued from the Preceding Table): Conditional, Restricted Model (9)

Re-normalisation of the eigenvectors			
Eigenvector(s) (transposed)	p_t^*	p_t	s_t
	-6.001	6.564	-6.242
The matrices based on 1 cointegration vectors			
β'	p_t^*	p_t	s_t
	1.000	-1.094	1.040
α			
Δp_t^*	-0.088		
Δp_t	0.098		
T-values for α			
	-3.448		
	3.168		
Π	p_t^*	p_t	s_t
Δp_t^*	-0.088	0.097	-0.092
Δp_t	0.098	-0.107	0.102
T-values for Π			
	-3.448	3.448	-3.448
	3.168	-3.168	3.168
Re-normalisation of the eigenvectors			
Eigenvector(s) (transposed)	p_t^*	p_t	s_t
	6.262	-6.262	6.262
The LR test:			
χ^2 (2)	0.11 (0.94)[a]		
β'[b]	p_t^*	p_t	s_t
	1.000	*-1.000*	*1.000*
α			
Δp_t^*	-0.095		
Δp_t	0.098		
T-values for α			
Δp_t^*	-3.554		
Δp_t	3.031		
Π	p_t^*	p_t	s_t
Δp_t^*	-0.095	0.095	-0.095
Δp_t	0.098	-0.098	0.098
T-values for Π			
	-3.554	3.554	-3.554
	3.031	-3.031	3.031

[a] P-value.

[b] The italic figures indicate the PPP-restriction null to be tested.

Moreover, compared with the diagnostic tests of the residuals in Table 21 for the unconditional model (7), the diagnostic tests for the conditional model in Table 25 indicate only slight or little change[42] in the standard deviations of residuals, the trace correlation, the p-values of the test of autocorrelation for L-B(49) and LM(1), the test statistic for the test of normality, and the coefficients of determination.

Table 24 Cointegration Analysis [10c] (Continued from the Preceding Table): Conditional, Restricted Model (9)

The short-run matrices:

$\tilde{\Phi}_l^{\Delta}$ for the lagged endogenous variables

Time: $t-1$		Time: $t-2$		Time: $t-3$	
Δp_t^*	Δp_t	Δp_t^*	Δp_t	Δp_t^*	Δp_t
0.029	-0.082	-0.048	-0.013	0.047	-0.085
-0.159	0.032	-0.139	-0.075	0.047	-0.085
t-values					
0.403	-1.395	-0.666	-0.223	0.643	-1.418
-1.839	0.453	-1.588	-1.063	-1.204	-0.392

Θ_l^{Δ} for the differences of the exogenous I(1) variables

Time: $t-0$	Time: $t-1$	Time: $t-2$	Time: $t-3$
Δs_t	Δs_t	Δs_t	Δs_t
-0.317	0.109	0.159	0.187
0.265	0.006	-0.019	0.017
t-values			
-2.438	0.822	1.202	1.453
1.679	0.039	-0.116	0.108

Ψ for the deterministic variables:

SEA(1)	SEA(2)	SEA(3)	SEA(4)	SEA(5)	SEA(6)
0.000	-0.058	-0.023	-0.021	-0.051	-0.063
0.032	0.025	0.021	0.044	-0.011	-0.007
t-values					
0.003	-2.667	-1.055	-0.976	-2.365	-2.902
1.216	0.953	0.824	1.717	-0.424	-0.256
SEA(7)	SEA(8)	SEA(9)	SEA(10)	SEA(11)	CONST
-0.026	-0.065	-0.014	-0.056	-0.018	0.496
-0.032	0.039	0.019	0.04	0.025	-0.509
t-values					
-1.182	-2.951	-0.621	-2.582	-0.831	3.563
-1.223	1.439	0.709	1.648	0.956	-3.020

One may then prefer the unconditional model (7) which will provide valuable information on the short-run dynamics of the exchange rate behavior, which is missing in the conditional/partial system (9).

[42] Harris (pp.115-117) illustrates similar evidence for the U.K. PPP and UIP model, suggesting both unconditional and conditional models (7) and (9) exhibit very similar residual behavior.

Table 25 Cointegration Analysis [10d] (Continued from the Preceding Table): Residual Analysis;[a] Conditional, Restricted Model (9)

Correlation Matrix	Δp_t^*	Δp_t				
	1.000					
	0.028	1.000				
Standard Deviations of Residuals	0.060	0.072				
Multivariate Statistics						
Log (Det (Sigma))	-10.889					
Information Criteria: SC	-9.597					
HQ	-10.074					
Trace Correlation	0.189					
Test for Autocorrelation						
L-B(49), χ^2 (182)	158.582	(0.89)[b]				
LM(1), χ^2 (4)	3.264	(0.51)				
LM(4), χ^2 (4)	2.884	(0.58)				
Test for Normality						
χ^2 (4)	30.000	(0.00)				
Univariate Statistics	Mean	Std. Dev.	Skewness	Kurtosis	Maximum	Minimum
	-0.000	0.060	-0.036	4.136	0.244	-0.173
	0.000	0.072	-0.427	4.782	0.207	-0.309
	ARCH(4)	Normality	R^2			
	14.866	11.472	0.207			
	25.598	18.667	0.176			

[a]See notes in Table 13.

[b]P-value.

5 Impulse Responses in the Cointegrated System

The previous section has investigated the responses of exchange rate and prices only by examining the short-run parameters' statistical significance in the estimated PPP-based VEC model (20). Here we will study the same problem, by employing more refined statistical methods of variance decomposition and impulse response function.

The estimated short-run effects as already given in eq. (20) may be directly used to perform the analyses of variance decomposition and impulse response function in the cointegrating system (7), with (8) rewritten as

$$\mathbf{\Pi} = (\ \alpha_{11} \quad \alpha_{21} \quad -\alpha_{31}\ )' \tag{24}$$

under the (strong) PPP restriction $(\ 1 \quad 1 \quad -1\ )$ on β' being supported by the data in the earlier section. For a purely technical programming reason, however, we will estimate once again the short-run effects for (7).

5.1 Setting a lag length and estimation

We again start out with selecting a lag length L for the underlying VAR model (6). At least a year's worth of lags is usually recommended (Doan, *UG*, p.332). Earlier in Table 12 we attempted to select L, based on the observation of the roots of the companion matrix. Here, we consider several pairs of lags to test a null (longer) length of lags, based on Tables 26 and 27.[43]

Table 26 Setting Lag Length [1]:[a] Testing the Null of l versus the Alternative of $l-1$

Lags l	AIC	SBC	LR Test	P-Value
1	-17.059*	-16.285*		
2	-17.039	-16.1100	14.227	0.115
3	-16.988	-15.903	8.380	0.496
4	-16.961	-15.722	12.929	0.166
5	-16.953	-15.558	16.420	0.059
6	-16.920	-15.370	11.804	0.225
7	-16.925	-15.221	19.094	0.024
8	-16.868	-15.009	7.220	0.614
9	-16.805	-14.790	6.062	0.734
10	-16.801	-14.632	17.291	0.044
11	-16.7690	-14.445	12.003	0.213
12	-16.703	-14.224	5.575	0.782

[a]See Doan (*UG*, p.336) for the test statistics used.

Table 27 Setting Lag Length [2]: Test of l_0 versus l_1 Lags

l_0	l_1	χ^2 Statisitc
12	4	71.093[a] (0.508)[b]
5	4	13.705[c] (0.133)
4	3	11.604 (0.237)
3	2	7.836 (0.551)

[a]The degrees of freedom is 72.

[b]P-value.

[c]The degrees of freedom is 9, which also applies to the tests below.

Although the sequential likelihood ratio tests in Table 26 show that there are indeed several possible lags that can be appropriate for the model, we only concentrate on $L \leq 5$ for the reason based on the roots of the companion matrix given in Table 12 in the previous section . Also, based on the first set of tests in Table 27, the null of $L = 12$ is not rejceted and yet it is not our chosen lag length for the same reason.

It is clear from Table 26 that the null of $L = 5$ is rejected at a 10% significance level (p-value=0.059), with the alternative of $L - 1 = 4$ being accepted. The decision here does support the earlier one made based on Table 12 . The second test in Table 27, however, could lead to an opposite decision; we will ignore this test result, for both Table 26 and the third set in Table 27 do lead to a decision to set $L = 4$.

By Table 26 and the third and the fourth sets in Table 27 combined

[43]See Doan (*UG*, p.336) for the test statistics used in the tables.

together, one could make L even shorter, such as three months. $L = 4$ and $L - 1 = 3$ will, however, be our final choice, respectively, for the underlying VAR model (6) and the VEC model (7), since too short a length could cause a serial correlation of the residuals of the model.

Table 28 VAR/System - Estimation by Cointegrated Least Squares: Model (7) with (24)

Dependent Variable	s_t	p_t^*	p_t
Δs_{t-1}	0.030 (0.693)[a]	0.099 (0.487)	0.014 (0.933)
Δs_{t-2}	0.160 (0.037)	0.109 (0.442)	0.024 (0.888)
Δs_{t-3}	0.025 (0.744)	0.179 (0.197)	0.023 (0.888)
Δp_{t-1}^*	-0.111 (0.007)	0.064 (0.397)	-0.188 (0.039)
Δp_{t-2}^*	0.024 (0.566)	-0.056 (0.474)	-0.132 (0.157)
Δp_{t-3}^*	0.007 (0.878)	0.045 (0.568)	-0.105 (0.269)
Δp_{t-1}	-0.001 (0.970)	-0.082 (0.199)	0.032 (0.675)
Δp_{t-2}	0.009 (0.801)	-0.016 (0.802)	-0.073 (0.336)
Δp_{t-3}	-0.093 (0.007)	-0.056 (0.380)	-0.053 (0.486)
Constant	0.014 (0.867)	0.491 (0.001)	-0.506 (0.005)
C_SEASON{-10}[b]	-0.005 (0.710)	-0.041 (0.095)	0.023 (0.434)
C_SEASON{-9}	-0.008 (0.524)	-0.002 (0.937)	0.004 (0.882)
C_SEASON{-8}	0.004 (0.760)	0.013 (0.600)	-0.018 (0.531)
C_SEASON{-7}	-0.009 (0.500)	0.017 (0.484)	0.010 (0.719)
C_SEASON{-6}	-0.013 (0.290)	-0.040 (0.091)	0.002 (0.933)
C_SEASON{-5}	-0.004 (0.760)	-0.008 (0.746)	0.001 (0.963)
C_SEASON{-4}	-0.002 (0.874)	-0.006 (0.792)	0.024 (0.391)
C_SEASON{-3}	0.005 (0.664)	-0.039 (0.094)	-0.029 (0.304)
C_SEASON{-2}	-0.012 (0.337)	-0.045 (0.058)	-0.029 (0.309)
C_SEASON{-1}	-0.004 (0.758)	-0.011 (0.651)	-0.052 (0.066)
C_SEASON	-0.010 (0.423)	-0.048 (0.044)	0.017 (0.559)
EC1{1}[c]	-0.003 (0.834)	-0.094 (0.001)	0.097 (0.005)
Monthly Data	1983:09 To 1999:12	1983:09 To 1999:12	1983:09 To 1999:12
Usable Observations	196	196	196
Degrees of Freedom	174	174	174
Adjusted R^2	0.038	0.084	0.063
Residual Standard Deviation	0.035	0.064	0.077
Durbin-Watson Statistic	1.952	2.005	2.001

[a]P-value.

[b]Following Doan (*UG*, p.46 and *RM*, p.368), the 11 centered seasonal dummies C_SEASON{-10 to 0} correspond, respectively, to $sdum_2$ (February dummy) to $sdum_{12}$ (December dummy) in eq. (20).

[c]EC1{1}$\equiv \boldsymbol{y}_{t-1} = (\; s_{t-1} \quad p_{t-1}^* \quad p_{t-1} \;)'$ and their corresponding coefficients are those α's in the right-hand side of (24).

We now again estimate the short-run effects for the cointegrating system (7) with (24); those newly estimated short-run matrices/effects are reported in Table 28.[44] Note, in the table, that the dependent variable indicated is a level variable (such as s_t) and yet the model actually estimated by the cointegrated least squares is a cointegrated system (7).

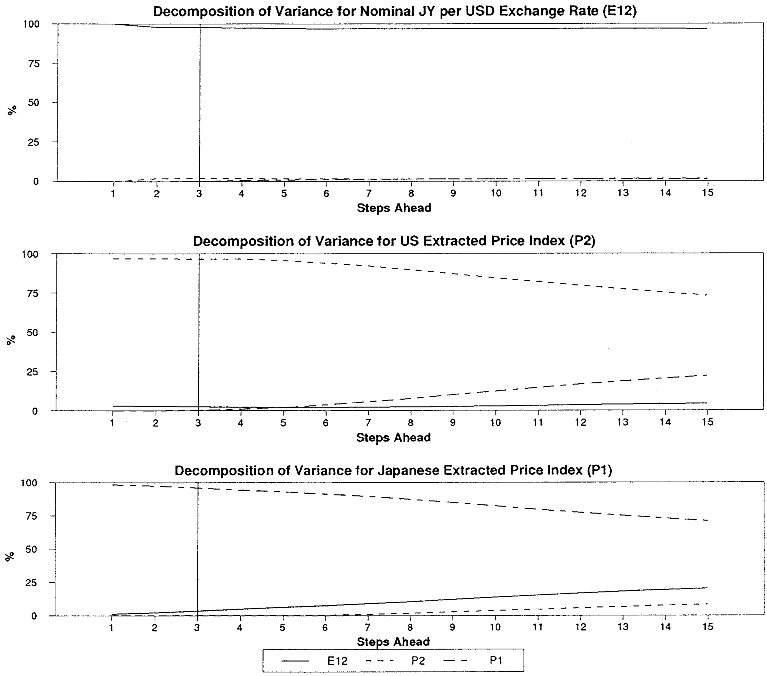

Figure 35 Variance Decomposition for s_t (E12), p_t^* (P2), and p_t (P1). The vertical grid line is drawn at $L-1=3$, the lag length of model (7); see Table 29. See Table 1 for notation of the symbols in the graph.

[44]Those estimates of short-run matrices are indeed close in magnitude to, and exactly the same in terms of sign and statistical siginificance as, those estimates in eq. (20) and t-values in Table 18. This does not apply to the deterministic terms (including a constant). This does not in any way affect the estimates of impulse response functions, however, as will be confirmed by comparing Figure 36 and Table 19.

5.2 Variance decomposition analysis

Information on interactions among the variables is now studied. In variance decomposition and impulse response function analyses, different orderings in the vector $\boldsymbol{y}_t = (s_t, p_t^*, p_t)'$ could yield different results on the interactions.[45] One criterion for ordering, as proposed by Sims (1980), is that contemporaneous causes come first. For example, the ordering (s_t, p_t^*, p_t) can be taken as a causality from exchange rate to prices, whereas its reversed ordering (p_t, p_t^*, s_t) as a causality from prices to exchange rate. Ito (2005) may prefer the latter (as quoted in section 1), but we will continue with the former that has been analyzed in the previous sections, since s_t is found to be weakly exogenous.[46]

Table 29 Decomposition of Variance[a]

	s_t				p_t^*				p_t			
Step	Std Error[b]	s_t	p_t^*	p_t	Std Error	s_t	p_t^*	p_t	Std Error	s_t	p_t^*	p_t
1	0.033	100.000	0.000	0.000	0.061	2.944	97.056	0.000	0.073	1.418	0.075	98.507
2	0.048	98.024	1.975	0.001	0.084	2.870	97.120	0.010	0.100	2.357	0.166	97.477
3	0.063	98.044	1.935	0.021	0.098	2.641	96.921	0.438	0.116	3.554	0.412	96.034
4	0.075	97.370	1.926	0.704	0.109	2.231	96.817	0.952	0.126	5.022	0.550	94.427
5	0.086	97.125	1.852	1.022	0.117	2.062	95.886	2.052	0.134	6.346	0.495	93.159
6	0.095	96.984	1.742	1.274	0.124	2.026	94.228	3.746	0.140	7.662	0.705	91.633
7	0.103	96.961	1.644	1.394	0.131	2.118	92.268	5.614	0.146	9.059	1.212	89.729
8	0.111	96.952	1.595	1.453	0.136	2.312	89.916	7.772	0.152	10.593	1.953	87.453
9	0.118	96.941	1.580	1.478	0.142	2.578	87.347	10.076	0.157	12.189	2.850	84.961
10	0.125	96.930	1.588	1.482	0.148	2.889	84.716	12.395	0.162	13.775	3.830	82.395
11	0.131	96.919	1.610	1.472	0.153	3.222	82.136	14.642	0.167	15.307	4.829	79.864
12	0.138	96.906	1.639	1.455	0.158	3.559	79.683	16.758	0.172	16.756	5.802	77.442
13	0.143	96.892	1.672	1.435	0.163	3.886	77.401	18.713	0.177	18.108	6.722	75.170
14	0.149	96.879	1.707	1.414	0.168	4.197	75.304	20.499	0.182	19.356	7.575	73.069
15	0.154	96.866	1.741	1.393	0.173	4.487	73.395	22.118	0.186	20.502	8.357	71.141

[a]See Figure 35.

[b]The standard error of forecast (or, more precisely, the variance of the one-step forecast error) for the "Series" indicated.

The decomposition of variance for a level variable s_t in Table 29, which is plotted in Figure 35, is consistent with the weak exogeneity of the yen per dollar exchange rate as detected in section 3.2: the variance of the one-step forecast error for s_t is accounted for nearly by own innovations. Notice, in Table 29, that the variance of the one-step forecast error for the U.S. extracted price index p_t^*, at the longer steps, is less explained by own innovations and more by the Japanese extracted price index p_t, but

[45]This will be most likely in the general case where the disturbance terms in the VEC model (7) are contemporaneously correlated. See, for example, Kojima (1996, section 3.3.2).

[46]Doan (*UG*, p. 353) recommends that an exogenous variable be put first in the ordering, and this is indeed satisfied by our chosen ordering (s_t, p_t^*, p_t). For weak exogeneity of s_t, see section 3.2.2.

only slightly by the yen per dollar exchange rate s_t. On the other hand, the variance of the one-step forecast error for the Japanese extracted price index p_t, at the longer steps, is less explained by own innovations and more by the yen per dollar exchange rate s_t, but only moderately by the U.S. extracted price index p_t^*.

The observed variance decomposition here will be combined later with impulse response functions, to infer robust responses of prices and exchange rate.

5.3 Impulse response functions

The impulse response functions and their two-standard-deviation confidence bands are computed and drawn in Figure 36, to complement the above variance decomposition analysis. Those kinks in the impulse response functions observed in the figure coincide, in sign though not in magnitude, with the statistically significant "short-run effects" in Table 19 (or in eq. (20)) earlier: the positive and negative short-run effects detected there are graphically displayed, respectively, as upward and downward kinks in the impulse response functions; for example, $\Delta s_{t-2}(+)$ in the short-run Δs_t equation corresponds to the upward kink in the top left impulse response function in Figure 36.

We now interpret the impulse responses in the figure, following Sims and Zha (1999, esp. p.1148, 1150-1153) and noting that the confidence bands are interpreted as indicating the degree of uncertainty about the shape of impulse responses estimated.

One remark is in order on impulse responses plotted in Figure 36. Impulse response functions are computed in such a way that their contemporaneous (i.e., time-zero) values in the *lower off-diagonal* of Figure 36 are all set equal to zero.[47] Therefore, in the following sections, no interpretations are given to these contemporaneously zero-valued responses in the lower off-diagonal of the figure. Even so, we will be able to draw economic insights from the remaining, later impulse responses.

First, common to all the shocks in Figure 36 is the horizontal impulse responses estimated beyond a certain month (at the earliest, around the second month) of the forecast horizon. This reflects the short-run nature ($\Delta \boldsymbol{y}_t$) of the system under study. The remaining interpretations of Figure 36 are given in the following two subsections.

[47] See Kojima (1996, section 3.3.2).

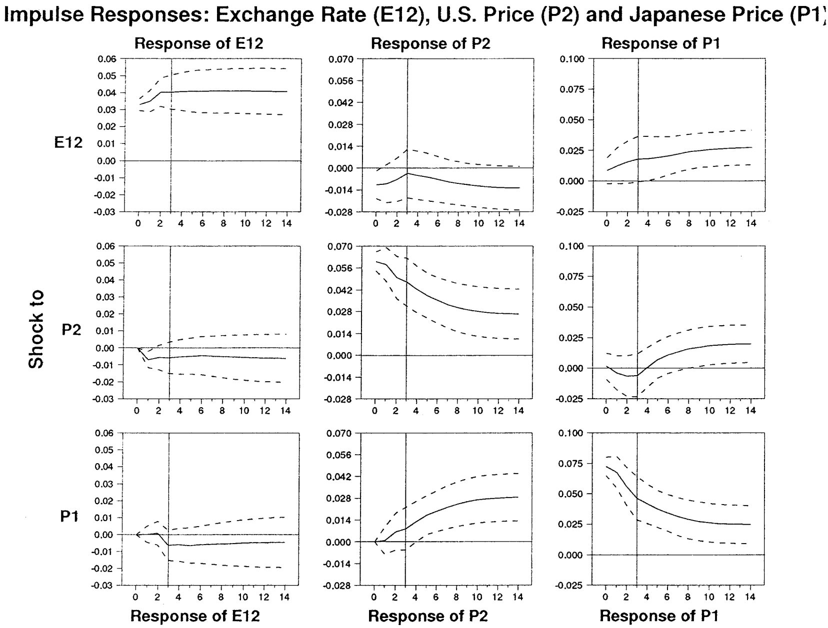

Figure 36 Impulse Responses for s_t (E12), p_t^* (P2), and p_t (P1). Forecast origin and horizon are, respectively, 0 and from 1 to 14 months. See Doan (*UG*, pp.350-355) for technicalities of computing the impulse response functions. The two-standard-deviation confidence bands (the dotted lines) for impulse responses are computed by Monte Carlo integration; see Doan (*UG*, ps.351, 472, 486) and Sims and Zha (p.1127). The vertical grid line is drawn at $L-1=3$, the lag length of model (7). For details on the kink(s) in each plot, see column "Short-run effects of" in Table 19. See Table 1 for notation of the symbols in the graph.

5.3.1 Exchange rate effects on prices

Positive shocks to the monthly exchange rate level s_t have the following effects (as shown in the first row of Figure 36):

- the *negative* response of U.S. price index p_t^*, contemporaneously and one to two months later, with gradually smaller absolute magnitude due to the (marginally) statistically significant kink $\Delta s_{t-3}(+)$ in the Δp_t^* equation in Table 19; and

- contemporaneously and up to three months later a *strong positive* response of Japanese price index p_t,[48] though without any statistically significant kink $\Delta s_{t-l}, l > 0$ in the Δp_t equation in Table 19.

Those exchange rate effects on the price changes are in line with those documented in the recent empirical study of exchange rate pass-through by Landon and Smith (2006),[49] and accord with our casual anticipation of the *Japanese yen depreciation* shocks leading to the *lower* U.S. dollar price of the Japanese goods exported to the U.S. and the *higher* Japanese yen price of the U.S. goods imported to Japan. The exchange rate effects here may be simply presented in the framework of PPP as row (a) of Table 30 which is constructed from the absolute PPP relation, $s_t + p_t^* = p_t$.

Two possible implications may be derived from Table 30 for a corporate traded-goods pricing policy:

- With a *negative* shock from s_t, row (a) of Table 30 may be interpreted as a yen-appreciation situation where Japanese exporters' exchange rate pass-through would tend toward *complete*. That is, as yen appreciation is more fully passed through by the Japanese exporters, the Japanese goods exported to the U.S. will be priced more highly in the U.S., which in turn would, in part, cause a positive response of p_t^*. This could suggest a possibility of such Japanese export pricing behavior during the 1990's, the later part of the sample period;

- Row (b) of Table 30 is consistent with, and thus may be interpreted as, the well-documented phenomenon during the later 1980's (the earlier part of the sample period) of Japanese exporters' more *incomplete* exchange rate pass-through.[50] That is, as yen appreciation is much less passed through, a negligible response of p_t^* is

[48] The lower confidence band is above zero almost throughout the horizon.

[49] See section 1 for the literature on exchange rate pass-through.

[50] It is puzzling, however, that during the second half of the 1980's period the Japanese export firms passed through only a little of yen exchange rate variation into their dollar-denominated export prices, whereas the degree of the U.S. firms' pass-through was almost complete.

expected possibly, in part, due to the only a little changing U.S. dollar price of the Japanese exports to the U.S.

The above implications derived from the observed exchange rate effects are consistent with the variance decomposition for p_t due to the exchange rate's innovation gradually *rising* (above that due to the p_t^* innovation) (see Figure 35).

Whether there occurred a (structural) shift in the Japanese export pricing strategy between the 1980's and the 1990's, as suggested above, requires a further investigation, which is beyond of the scope of the present paper.

Table 30 Shock-response Based on Absolute PPP: Exchange Rate Effects; Three-Month Horizon

	Shock	Signs of responses[a]	
	s_t	p_t^*	p_t
(a)	+	−	+
(b)	+	Toward 0, three months later	+

[a]See the first row of Figure 36.

Table 31 Shock-response Based on Absolute PPP: Response of Exchange Rate to U.S. Price; Three-Month Horizon

	Shock	Signs of responses[a]	
	p_t^*	s_t	p_t
(a)	+	−, initially	$0 \to -$

[a]See the second row of Figure 36.

Table 32 Shock-response Based on Absolute PPP: Response of Exchange Rate to Japanese Price; Three-Month Horizon

	Shock	Signs of responses[a]	
	p_t	s_t	p_t^*
(a)	+	$0(\text{or} \pm) \to -$	$0 \to +$

[a]See the third row of Figure 36.

5.3.2 Responses of exchange rate to prices and responses between prices

Stock and Watson (2003, p.798) note that "Exchange rates are a channel through which inflation can be imported in open economies." The role of exchange rate as such a channel may be drawn, based on absolute PPP, as simple flows of impact: $p_t^* \to s_t \to p_t$ and $p_t \to s_t \to p_t^*$. Clearly, the first half of the flows of impact here is the causality direction from prices to exchange rate that is claimed to be in line with PPP by Ito (2005) (as quoted in section 1). The impulse responses of exchange rate to prices *and* those of U.S. [Japanese] price to Japanese [U.S.] price, *combined together*, would help draw these impact flows and thus study the role of exchange rate as such a channel.

Focusing on the three-month long horizon, the estimated responses of exchange rate to prices and between prices in Figure 36 are summarized as (i) and (ii):

(i) Positive shocks to the monthly U.S. price index p_t^* have the following effects (as shown in the second row of Figure 36):

- one to three months later a *negative* response of the exchange rate (i.e., an *appreciation* of the Japanese yen), with the (strongly) statistically significant kink $\Delta p_{t-1}^*(-)$ in the Δs_t equation in Table 19; and

- no contemporaneous response, and then gradually the *negative* response possibly over the three-month horizon of the Japanese price index p_t, with the strongly statistically significant kink $\Delta p_{t-1}^*(-)$ and marginally significant kink $\Delta p_{t-2}^*(-)$ in the Δp_t equation in Table 19.

The impulse response mechanism here is summarized in Table 31. The table suggests the flow of impact, over the three-month horizon:
positive shock to the U.S. price p_t^*
$\to$ negative response of the yen per dollar rate s_t
$\to$(initially zero, and) later negative response of the Japanese price p_t.
The flow of impact here is in line with our casual observation that, as a result of yen appreciation, the lower yen prices of goods imported from the U.S. would lower in part the Japanese price p_t. There is indeed observed an unambiguous flow over the three-month horizon of impact from the U.S. price to the Japanese price, channeled through exchange rate.

(ii) Positive shocks to the monthly Japanese price index p_t have the effects (as shown in the third row of Figure 36):

- one to two months later *no* (or, equally likely positive or negative) response and a *negative* response three months later of the exchange rate, the latter of which is due to the (strongly) statistically significant kink $\Delta p_{t-3}(-)$ in the Δs_t equation in Table 19; and

- one month later no response and the *positive* response two and three months later of the U.S. price index p_t^*, although with (marginally) statistically significant kink $\Delta p_{t-1}(-)$ in the Δp_t^* equation in Table 19.

The impulse response chain here is summarized in Table 32. The table suggests the flow over the three-month horizon of impact:

positive shock to the Japanese price p_t

→initially ambiguous but 3 months later negative response of the yen per dollar rate s_t

→(initially ambiguous, but) later positive response of the U.S. price p_t^*.

In contrast to the unambiguous flow in (i), the (opposite) flow of impact from the Japanese price to the U.S. price, via exchange rate, is somewhat ambiguous to the extent that, initially, no or ambiguous Japanese price effects are observed on the exchange rate. Later, though, the yen appreciates (i.e., U.S. dollar depreciates), and then the U.S. price will rise, possibly in part, due to higher dollar prices of Japanese goods imported to the U.S.

Analyzing two impact flows, (i) and (ii), we thus find that the U.S. price and Japanese price effects on the yen per dollar rate turn out *both* negative (after three months). How could they be both negative? Under relative PPP, $\Delta s_t = \Delta p_t - \Delta p_t^*$, positive shock to the Japanese price would be expected to *contemporaneously* have a *positive* impact on Δs_t (through Δp_t), while that to the U.S. price a negative impact (through $-\Delta p_t^*$). It is not immediately clear how, with these two contemporaneous, opposing impacts derived under relative PPP, "both negative" effects could still result three months after the shocks to the prices. An attempt to resolve the apparent puzzling conflict may require refinements in testing. One obvious refinement is to test for the Granger noncausality null hypothesis in our cointegrated system (7) (as remarked in section 3.2.3). This will be a future extension of the present study.

5.3.3 Summary

The U.S. price, initially, appears to have more definitive (negative) effects on the yen per dollar exchange rate than the Japanese price would have on the exchange rate;[51] but, later (over the horizon of three months), the observed exchange rate changes seem to unambiguously channel inflation of one country into another country. The observed exchange rate effects on prices are found to be in accord with both our casual anticipation and the recent empirical result in the exchange rate pass-through literature.

6 Concluding Remarks

Relying on the estimates of Japanese and U.S. inflation rates extracted from stock returns by Chowdhry, Roll and Xia (2005), our exploration starts out with the cointegration analysis of the yen per U.S. dollar nominal exchange rate (s_t) and Japanese and U.S. prices (p_t, p_t^*) to study their long-run structure (i.e., their PPP relationship), and then, based on the estimated PPP-based vector error-correction model, we conduct analyses of variance decomposition and impulse response functions to examine the short-run structure (i.e., the impulse responses of exchange rate and prices). Strong evidence is documented in support of (i) the PPP restriction which yields the equilibrium error in the form of a real exchange rate. Evidenced further under the PPP relationship so documented are (ii) the impulse responses of exchange rate to prices and between prices, combined together, that would imply exchange rates channeling inflations into countries and (iii) the impulse responses of prices to exchange rate (i.e., exchange rate effects on prices) that would usefully indicate the degree of exchange rate pass-through by Japanese exporters.

Several main findings that lead to the results (i) through (iii) are summarized as follows:

First, in the cointegration analysis with $\boldsymbol{y}_t = (s_t, p_t^*, p_t)'$, the null of a (strong) PPP restriction $(\ 1 \quad 1 \quad -1\ )$ on the cointegration vector $\boldsymbol{\beta}'$

[51] It should be noted that the possible responses of exchange rate to prices as detected in (i) and (ii) may be somewhat ambiguous, since more than 90% of the exchange rate's variance decomposition is accounted for by *own* innovation (see Figure 35). The ambiguity here is consistent, in particular, with two confidence bands having zero in between for the Japanese price index (as shown in the left most in the third row of Figure 36).

is readily accepted by the LR test. C-R-X's estimated inflation rates are thus shown to have desired PPP-theoretic content, here in the VEC context. This adds to C-R-X's evidence of the same theoretical content documented in their single equation-based analysis of the PPP relationship. The same inference is derived for the conditional version of the cointegrating system which takes into account the detected weak exogeneity of the exchange rate.

Second, the estimated speed of adjustment $\boldsymbol{\alpha}$ shows that s_t appears weakly exogenous to the system of equations (7). That is, the short-run change Δs_t will not adjust to the previous equilibrium error (i.e., the previous real exchange rate $s_{t-1} + p^*_{t-1} - p_{t-1}$), while two short-run price changes Δp^*_t and Δp_t do adjust, respectively, in a negative and positive direction, but in almost the same speed, to the equilibrium error.

Third, the short-run Δs_t equation in the estimated PPP-based system (20) contains strongly statistically significant short-run effects of Δs_{t-2}, Δp^*_{t-1} and Δp_{t-3}. Thus, the responses of exchange rate to prices appear to be detected here. Inferred at the same time are some responses of prices to exchange rate (i.e., exchange rate effects on prices) in the Δp^*_t equation (but not in the Δp_t equation): exchange rate as a financial variable may lead changes in U.S. goods prices.

Finally, from the impulse response functions and their confidence bands computed, we infer two flows over the three-month long horizon of impact from one price to exchange rate and, further *via exchange rate*, on to another price: (i) positive shock to the U.S. price p^*_t → negative response of the yen per dollar rate s_t → negative response of the Japanese price p_t; and (ii) positive shock to the Japanese price p_t → initially ambiguous but 3 months later negative response of the yen per dollar rate s_t → (initially ambiguous but) later positive response of the U.S. price p^*_t. That is, initially, the U.S. price appears to have more definitive (negative) effects on the yen per dollar exchange rate than the Japanese price would have on the exchange rate, but, later (over the horizon of three months), the observed exchange rate changes seem to unambiguously channel inflation of one country into another country.

The exchange rate effects on prices are also detected over the three month horizon: positive shocks to the monthly exchange rate level s_t leads to the negative response of U.S. price index p^*_t, and the strong positive response of Japanese price index p_t. Two possible implications for a corporate traded-goods pricing policy are that a yen-appreciation situation where Japanese exporters' exchange rate pass-through would tend toward complete is observed possibly during the 1990's, the latter part of the sample period, and that the well-documented phenomenon that Japanese exporters exhibited more incomplete exchange rate pass-through during the later 1980's, the early part of the sample period, is also observed.

C-R-X also extracted the U.K. and German inflation rates from the assoicated stock returns. Would the (strong) PPP restriction be satisfied by the VEC models for the two countries as well? What would the impulse responses look like? Would testing for the Granger non-causality null hypothesis in our cointegrated system (7) lead to results consistent with the impulse responses? These are the topics that deserve further research, in an attempt to provide additional international evidence on PPP, impulse responses and Granger causality in the VEC framework.

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A Data Appendix

This appendix tabulates all the data used, or simply referred to, in the paper: Tables 33 and 34 are, respectively, for Japanese and U.S. data, May 1983 through December 1999. The data sources are detailed in section 2.

Table 33 Japanese Data (Percent per Month, except for ipi and jpusfxr)[a]

	jpmkt	jpsmb	jphml	jptb	jpcpi	jpinfl[b]	ipi_jpcpi	ipi_jpinfl	jpusfx	Date	Day	jpusfxr
mean	0.749	0.033	0.128	0.297	0.100	0.297			-0.424			
std dev	5.916	3.342	3.069	0.208	0.468	7.790	1.000[c]	1.000	3.520	1983.4.29	Fri	238.31
198305	1.518	-1.639	2.627	0.505	0.972	9.564	1.010	1.096	0.231	1983.5.31	Tue	238.86
198306	3.163	-0.413	0.375	0.524	-0.601	0.885	1.004	1.105	0.176	1983.6.30	Thu	239.28
198307	1.557	-1.142	-5.753	0.529	-0.364	1.955	1.000	1.127	1.023	1983.7.29	Fri	241.74
198308	2.083	-0.932	2.126	0.525	-0.243	1.758	0.998	1.147	1.930	1983.8.31	Wed	246.45
198309	1.973	-3.091	7.515	0.537	1.097	0.516	1.009	1.153	-4.477	1983.9.30	Fri	235.66
198310	-0.690	2.283	-0.660	0.538	0.843	2.717	1.017	1.184	-0.690	1983.10.31	Mon	234.04
198311	0.460	-0.749	1.706	0.509	-0.478	4.775	1.012	1.241	-0.664	1983.11.30	Wed	232.49
198312	6.515	-1.385	-7.894	0.530	-0.360	10.308	1.009	1.368	-0.336	1983.12.30	Fri	231.71
198401	6.300	0.196	-2.070	0.535	0.361	-0.158	1.012	1.366	1.274	1984.1.31	Tue	234.68
198402	-0.016	3.741	-1.815	0.510	0.601	-5.656	1.018	1.289	-0.534	1984.2.29	Wed	233.43
198403	11.798	-1.202	-4.106	0.530	0.239	-6.398	1.021	1.207	-3.789	1984.3.30	Fri	224.75
198404	-0.561	0.816	7.168	0.520	0.238	-2.458	1.023	1.177	0.996	1984.4.30	Mon	227
198405	-9.870	0.445	0.442	0.501	0.712	-3.967	1.030	1.130	2.002	1984.5.31	Thu	231.59
198406	2.405	-3.708	-3.762	0.516	-0.707	5.425	1.023	1.191	2.444	1984.6.29	Fri	237.32
198407	-3.613	5.789	1.625	0.510	0.119	3.363	1.024	1.232	3.307	1984.7.31	Tue	245.3
198408	7.071	-2.021	-3.132	0.512	-0.831	10.297	1.016	1.358	-1.400	1984.8.31	Fri	241.89
198409	1.072	-1.364	-5.619	0.515	1.555	3.395	1.032	1.404	2.046	1984.9.28	Fri	246.89
198410	4.452	3.043	-3.142	0.516	0.707	0.857	1.039	1.417	-0.491	1984.10.31	Wed	245.68
198411	1.887	5.143	2.714	0.510	-0.468	-4.194	1.034	1.357	0.738	1984.11.30	Fri	247.5
198412	4.385	-1.327	1.769	0.526	0.235	-3.713	1.036	1.307	1.643	1984.12.31	Mon	251.6
198501	2.035	0.886	-0.878	0.516	0.234	4.884	1.039	1.371	1.256	1985.1.31	Thu	254.78
198502	5.053	1.327	-1.494	0.512	-0.351	-14.462	1.035	1.172	1.816	1985.2.28	Thu	259.45
198503	2.345	1.169	3.561	0.521	0.353	-9.450	1.039	1.062	-3.315	1985.3.29	Fri	250.99
198504	-3.127	0.459	-1.242	0.510	0.584	4.608	1.045	1.110	0.207	1985.4.30	Tue	251.51
198505	3.172	2.834	5.978	0.510	0.233	-5.600	1.047	1.048	-0.115	1985.5.31	Fri	251.22
198506	3.028	0.697	4.559	0.517	0.000	-0.513	1.047	1.043	-1.129	1985.6.28	Fri	248.4
198507	-3.252	3.734	2.570	0.515	0.232	-9.925	1.050	0.939	-4.939	1985.7.31	Wed	236.43
198508	2.748	0.879	0.812	0.516	-0.231	2.432	1.047	0.962	1.085	1985.8.30	Fri	239.01
198509	0.841	0.705	2.173	0.519	0.117	-2.499	1.049	0.938	-9.896	1985.9.30	Mon	216.49
198510	0.105	-2.837	-6.313	0.523	0.927	13.401	1.058	1.064	-2.304	1985.10.31	Thu	211.56
198511	-1.844	2.130	-1.316	0.583	-0.804	8.028	1.050	1.149	-4.565	1985.11.29	Fri	202.12
198512	4.118	-1.153	-0.114	0.619	0.116	7.373	1.051	1.234	-0.934	1985.12.31	Tue	200.24

(Continued on next page)

[a]Symbols "jpmkt, jpsmb, jphml, jptb, jpcpi, jpinfl, jpusfx" correspond, respectively, to those in C-R-X (Table 1 for the three Fama-French factor returns, p.262) as follows: Japan market, Japan SMB, Japan HML, Japan Tbill, Japanese CPI inflation, Japanese extracted risk-free rate $\hat{R}^J_{ft}$, Japanese yen-US$. Symbols "ipi_jpcpi, ipi_jpinfl, jpusfxr" correspond, respectively, to those in Table 1 in the present paper as follows: $P^J_{CPI,t}$, $P^J_{R,t}$, (unlogged, raw) yen per dollar exchange rate.

[b]This represents "extracted Rf."

[c]Initial value of price index at time 0.

Table 33 (Continued)

	jpmkt	jpsmb	jphml	jptb	jpcpi	jpinfl	ipi_jpcpi	ipi_jpinfl	jpusfx	Date	Day	jpusfxr
198601	-0.428	3.532	-1.087	0.599	0.346	10.165	1.055	1.360	-4.046	1986.1.31	Fri	192.3
198602	4.792	3.152	1.569	0.506	-0.115	-1.095	1.053	1.345	-6.299	1986.2.28	Fri	180.56
198603	16.154	-5.399	-1.488	0.500	0.000	-22.381	1.053	1.044	-1.647	1986.3.31	Mon	177.61
198604	-1.036	-0.990	0.777	0.438	0.231	22.184	1.056	1.275	-5.795	1986.4.30	Wed	167.61
198605	3.929	4.087	1.225	0.378	0.345	-1.072	1.060	1.262	4.028	1986.5.30	Fri	174.5
198606	4.616	5.177	3.979	0.387	-0.573	0.825	1.053	1.272	-6.328	1986.6.30	Mon	163.8
198607	5.198	-1.423	-3.330	0.385	-0.231	-7.788	1.051	1.173	-6.260	1986.7.31	Thu	153.86
198608	6.906	-7.791	0.310	0.386	-0.231	-13.783	1.049	1.011	0.447	1986.8.29	Fri	154.55
198609	-0.558	-4.382	2.619	0.392	0.463	-10.824	1.053	0.902	-0.123	1986.9.30	Tue	154.36
198610	-7.143	1.344	0.280	0.400	0.115	20.517	1.055	1.087	5.685	1986.10.31	Fri	163.39
198611	5.951	2.075	-0.513	0.371	-0.461	10.922	1.050	1.206	-0.817	1986.11.28	Fri	162.06
198612	4.871	-2.913	-0.093	0.365	-0.231	5.864	1.047	1.276	-2.341	1986.12.31	Wed	158.31
198701	12.100	0.945	-1.726	0.365	-0.347	-32.130	1.044	0.866	-2.981	1987.1.30	Fri	153.66
198702	1.464	-1.675	-1.423	0.334	0.000	-11.103	1.044	0.770	-0.261	1987.2.27	Fri	153.26
198703	5.441	-0.489	-4.261	0.330	0.349	-15.573	1.047	0.650	-5.065	1987.3.31	Tue	145.69
198704	11.124	-1.260	2.608	0.334	0.928	-9.828	1.057	0.586	-3.506	1987.4.30	Thu	140.67
198705	2.744	2.332	-0.773	0.319	0.230	24.284	1.060	0.729	2.409	1987.5.29	Fri	144.1
198706	-4.442	3.693	0.562	0.310	-0.229	17.633	1.057	0.857	1.856	1987.6.30	Tue	146.8
198707	-1.131	3.098	3.639	0.309	-0.460	5.731	1.052	0.906	2.096	1987.7.31	Fri	149.91
198708	6.947	2.801	-2.631	0.306	0.116	-0.707	1.053	0.900	-5.597	1987.8.31	Mon	141.75
198709	-0.777	-0.512	3.932	0.300	0.807	7.063	1.062	0.963	3.282	1987.9.30	Wed	146.48
198710	-12.574	2.894	-2.809	0.314	0.000	-10.857	1.062	0.859	-5.681	1987.10.30	Fri	138.39
198711	-1.036	0.824	1.004	0.316	-0.458	8.074	1.057	0.928	-4.417	1987.11.30	Mon	132.41
198712	-6.517	2.552	1.977	0.318	-0.114	12.583	1.056	1.045	-8.805	1987.12.31	Thu	121.25
198801	11.111	2.039	-0.656	0.323	-0.230	-5.126	1.053	0.991	5.222	1988.1.29	Fri	127.75
198802	8.462	-1.940	-0.198	0.315	-0.231	0.415	1.051	0.995	0.546	1988.2.29	Mon	128.45
198803	3.406	1.041	3.115	0.315	0.346	0.459	1.055	1.000	-3.437	1988.3.31	Thu	124.11
198804	1.972	2.754	0.951	0.320	0.577	11.146	1.061	1.111	0.659	1988.4.29	Fri	124.93
198805	-2.738	6.676	2.335	0.312	0.000	-2.548	1.061	1.083	0.144	1988.5.31	Tue	125.11
198806	2.604	0.967	5.266	0.315	-0.115	-3.067	1.060	1.050	6.513	1988.6.30	Thu	133.53
198807	1.939	-3.887	5.667	0.315	-0.229	4.258	1.057	1.094	-0.360	1988.7.29	Fri	133.05
198808	-4.299	0.348	-2.800	0.316	0.345	0.299	1.061	1.098	2.575	1988.8.31	Wed	136.52
198809	1.138	-4.030	-6.265	0.331	0.687	-4.119	1.068	1.053	-1.938	1988.9.30	Fri	133.9
198810	0.278	-0.194	0.172	0.350	0.569	-11.322	1.074	0.933	-6.487	1988.10.31	Mon	125.49
198811	6.043	2.271	2.042	0.347	-0.452	3.468	1.069	0.966	-2.903	1988.11.30	Wed	121.9
198812	3.161	-2.188	-2.956	0.338	-0.228	11.529	1.067	1.077	2.551	1988.12.30	Fri	125.05
198901	4.583	6.009	6.749	0.347	-0.228	5.150	1.064	1.133	4.266	1989.1.31	Tue	130.5
198902	-0.705	-0.094	1.519	0.342	-0.342	-9.788	1.061	1.022	-2.821	1989.2.28	Tue	126.87
198903	1.146	0.820	1.930	0.343	0.572	4.361	1.067	1.066	4.538	1989.3.31	Fri	132.76
198904	0.787	2.215	1.818	0.350	1.709	4.761	1.085	1.117	0.068	1989.4.28	Fri	132.85
198905	1.958	0.849	3.571	0.346	0.560	9.849	1.091	1.227	6.900	1989.5.31	Wed	142.34
198906	-3.443	0.754	-1.041	0.358	0.000	1.346	1.091	1.244	1.159	1989.6.30	Fri	144
198907	7.331	0.819	0.218	0.369	-0.223	0.455	1.089	1.249	-5.020	1989.7.31	Mon	136.95
198908	-0.966	4.620	3.338	0.388	-0.112	10.068	1.087	1.375	5.442	1989.8.31	Thu	144.61
198909	3.973	2.246	1.256	0.395	0.894	5.292	1.097	1.448	-3.519	1989.9.29	Fri	139.61
198910	-0.352	2.556	-0.643	0.413	0.776	-7.283	1.106	1.342	2.217	1989.10.31	Tue	142.74
198911	5.090	0.277	-2.499	0.438	-1.099	-2.592	1.094	1.308	0.126	1989.11.30	Thu	142.92
198912	1.851	1.556	0.088	0.469	0.111	5.366	1.095	1.378	0.607	1989.12.29	Fri	143.79
199001	-4.987	3.923	0.527	0.488	0.444	2.804	1.100	1.416	0.527	1990.1.31	Wed	144.55
199002	-6.276	0.489	-5.496	0.504	0.111	-4.604	1.101	1.351	2.918	1990.2.28	Wed	148.83
199003	-12.923	2.258	8.347	0.510	0.441	0.710	1.106	1.361	5.871	1990.3.30	Fri	157.83
199004	-0.964	-6.281	-3.473	0.538	0.770	5.908	1.114	1.441	0.644	1990.4.30	Mon	158.85
199005	10.422	0.238	-3.717	0.550	0.545	0.007	1.120	1.441	-3.988	1990.5.31	Thu	152.64
199006	-3.775	5.944	7.754	0.563	-0.434	10.208	1.115	1.588	-0.197	1990.6.29	Fri	152.34
199007	-3.873	3.553	3.997	0.568	-0.218	5.861	1.113	1.681	-4.162	1990.7.31	Tue	146.13
199008	-12.361	-1.967	1.406	0.581	0.437	-12.968	1.118	1.463	-1.517	1990.8.31	Fri	143.93
199009	-20.206	0.259	0.261	0.604	0.761	-22.162	1.126	1.139	-4.012	1990.9.28	Fri	138.27
199010	18.156	-2.097	-4.360	0.614	1.079	20.253	1.139	1.370	-6.206	1990.10.31	Wed	129.95
199011	-10.989	-1.317	2.739	0.618	-0.107	-15.744	1.137	1.154	2.395	1990.11.30	Fri	133.1
199012	4.985	-4.284	-1.543	0.623	-0.106	-6.213	1.136	1.082	1.971	1990.12.31	Mon	135.75
199101	-1.317	-3.993	0.716	0.640	0.642	-2.035	1.143	1.060	-3.264	1991.1.31	Thu	131.39
199102	14.591	4.929	0.668	0.632	-0.319	11.527	1.140	1.183	1.180	1991.2.28	Thu	132.95
199103	0.880	2.613	-0.592	0.633	0.534	0.713	1.146	1.191	5.602	1991.3.29	Fri	140.61
199104	-0.368	0.729	0.194	0.643	0.530	1.414	1.152	1.208	-3.121	1991.4.30	Tue	136.292
199105	0.075	-0.431	0.625	0.638	0.528	-5.579	1.158	1.140	1.529	1991.5.31	Fri	138.393
199106	-7.398	1.490	5.573	0.628	-0.420	5.613	1.153	1.204	-0.311	1991.6.28	Fri	137.955
199107	2.207	-2.378	-0.400	0.625	-0.106	-4.505	1.152	1.150	-0.407	1991.7.31	Wed	137.402
199108	-6.823	-1.650	-0.692	0.588	0.211	-8.905	1.154	1.048	-0.401	1991.8.30	Fri	136.848
199109	6.060	-0.717	0.463	0.584	0.212	12.954	1.157	1.184	-2.966	1991.9.30	Mon	132.848
199110	3.018	0.660	-1.019	0.547	1.050	-2.095	1.169	1.159	-1.739	1991.10.31	Thu	130.555
199111	-8.266	-0.699	-0.750	0.519	0.208	-9.089	1.171	1.053	-0.391	1991.11.29	Fri	130.054
199112	-0.927	-1.457	-0.940	0.501	-0.519	0.964	1.165	1.064	-4.081	1991.12.31	Tue	124.849
199201	-4.878	-2.939	0.858	0.493	-0.104	3.816	1.164	1.104	0.594	1992.1.31	Fri	125.593
199202	-4.672	1.019	-0.502	0.422	-0.105	3.088	1.163	1.138	2.829	1992.2.28	Fri	129.197
199203	-8.329	-1.485	1.602	0.423	0.418	-4.325	1.168	1.089	2.829	1992.3.31	Tue	132.904
199204	-7.121	-1.917	8.772	0.394	1.040	0.043	1.180	1.089	0.370	1992.4.30	Thu	133.397
199205	4.476	1.701	-4.977	0.376	0.103	0.038	1.181	1.090	-4.365	1992.5.29	Fri	127.699
199206	-10.148	0.723	1.239	0.373	-0.103	-0.294	1.180	1.087	-1.463	1992.6.30	Tue	125.844
199207	-1.367	-2.205	-1.822	0.370	-0.720	-10.248	1.171	0.975	1.187	1992.7.31	Fri	127.347
199208	13.650	-2.655	-2.277	0.323	0.311	-4.724	1.175	0.929	-3.514	1992.8.31	Mon	122.95
199209	-5.090	3.312	-2.275	0.304	0.414	-6.123	1.180	0.872	-2.429	1992.9.30	Wed	120
199210	-2.415	-2.088	2.898	0.309	0.103	1.459	1.181	0.885	2.681	1992.10.30	Fri	123.261
199211	3.482	-3.083	-0.059	0.300	-0.206	5.838	1.179	0.937	0.881	1992.11.30	Mon	124.352
199212	-1.143	3.125	1.883	0.297	0.000	-1.991	1.179	0.918	0.311	1992.12.31	Thu	124.739

(Continued on next page)

Table 33 (Continued)

	jpmkt	jpsmb	jphml	jptb	jpcpi	jpinfl	ipi_jpcpi	ipi_jpinfl	jpusfx	Date	Day	jpusfxr
199301	-0.664	-2.150	0.256	0.298	0.000	1.251	1.179	0.930	0.020	1993.1.29	Fri	124.764
199302	-1.110	1.620	-2.941	0.284	0.000	-0.956	1.179	0.921	-5.567	1993.2.26	Fri	118.008
199303	11.998	1.114	0.545	0.242	0.309	-8.632	1.182	0.841	-2.653	1993.3.31	Wed	114.918
199304	13.196	1.858	4.511	0.246	0.720	4.302	1.191	0.877	-3.281	1993.4.30	Fri	111.209
199305	0.981	10.201	-6.617	0.247	0.103	-1.482	1.192	0.864	-3.831	1993.5.31	Mon	107.029
199306	-3.416	-0.137	-0.235	0.249	-0.102	-1.450	1.191	0.852	-0.326	1993.6.30	Wed	106.681
199307	5.044	-3.148	2.255	0.245	0.204	2.595	1.193	0.874	-1.593	1993.7.30	Fri	104.995
199308	2.009	-0.600	1.066	0.245	0.305	-3.350	1.197	0.845	-0.321	1993.8.31	Tue	104.658
199309	-3.705	-1.919	0.534	0.231	0.099	10.705	1.198	0.935	1.288	1993.9.30	Thu	106.015
199310	0.269	-4.493	2.260	0.187	-0.099	-6.160	1.197	0.878	2.231	1993.10.29	Fri	108.407
199311	-15.728	-1.281	2.627	0.184	-0.507	-6.662	1.191	0.819	0.655	1993.11.30	Tue	109.119
199312	4.789	-1.569	0.214	0.175	0.103	1.788	1.192	0.834	2.280	1993.12.31	Fri	111.635
199401	13.201	1.838	3.424	0.154	0.101	-9.142	1.193	0.758	-2.659	1994.1.31	Mon	108.706
199402	0.168	0.661	2.889	0.157	0.000	1.734	1.193	0.771	-4.271	1994.2.28	Mon	104.161
199403	-3.845	4.538	0.387	0.158	0.404	0.614	1.198	0.775	-1.686	1994.3.31	Thu	102.42
199404	2.569	4.313	-0.161	0.158	0.210	0.069	1.201	0.776	-0.381	1994.4.29	Fri	102.031
199405	4.942	-0.099	2.052	0.159	0.100	-3.103	1.202	0.752	2.623	1994.5.31	Tue	104.743
199406	-0.520	3.111	0.772	0.155	-0.407	2.779	1.197	0.773	-6.263	1994.6.30	Thu	98.384
199407	-2.144	-0.436	-1.428	0.157	-0.405	3.334	1.192	0.798	1.644	1994.7.29	Fri	100.015
199408	0.192	-1.029	0.808	0.158	0.407	-0.798	1.197	0.792	-0.008	1994.8.31	Wed	100.007
199409	-3.621	-1.118	-0.431	0.162	0.309	-3.517	1.201	0.764	-0.987	1994.9.30	Fri	99.025
199410	0.495	1.330	2.118	0.166	0.499	0.732	1.207	0.770	-2.186	1994.10.31	Mon	96.884
199411	-4.050	-1.003	0.105	0.169	-0.298	0.517	1.203	0.774	2.094	1994.11.30	Wed	98.934
199412	2.572	0.869	1.538	0.170	-0.299	-2.712	1.199	0.753	0.631	1994.12.30	Fri	99.56
199501	-6.103	0.511	-0.568	0.171	0.000	-12.934	1.199	0.655	-0.083	1995.1.31	Tue	99.477
199502	-7.871	0.297	-2.724	0.169	-0.309	2.804	1.196	0.674	-2.967	1995.2.28	Tue	96.569
199503	-2.531	-0.398	1.885	0.170	-0.102	0.365	1.194	0.676	-10.680	1995.3.31	Fri	86.787
199504	1.831	-1.002	-0.816	0.152	0.303	-3.067	1.198	0.656	-3.290	1995.4.28	Fri	83.978
199505	-5.830	-2.336	-1.897	0.104	0.210	0.518	1.201	0.659	0.733	1995.5.31	Wed	84.596
199506	-4.538	-3.036	0.345	0.097	-0.100	1.110	1.199	0.666	0.152	1995.6.30	Fri	84.725
199507	11.625	-0.781	-0.495	0.093	-0.512	3.163	1.193	0.687	4.145	1995.7.31	Mon	88.311
199508	6.860	3.361	0.351	0.059	0.101	0.850	1.194	0.693	10.221	1995.8.31	Thu	97.815
199509	1.049	-1.831	-0.193	0.061	0.613	5.596	1.202	0.732	1.231	1995.9.29	Fri	99.027
199510	-1.878	1.316	-0.069	0.029	-0.309	6.970	1.198	0.783	3.205	1995.10.31	Tue	102.252
199511	5.042	0.794	-2.511	0.026	-0.302	-5.365	1.194	0.741	-0.490	1995.11.30	Thu	101.752
199512	6.462	1.510	1.519	0.018	0.000	-3.010	1.194	0.719	1.479	1995.12.29	Fri	103.268
199601	2.249	3.307	1.466	0.020	-0.101	5.468	1.193	0.758	3.503	1996.1.31	Wed	106.95
199602	-3.247	-0.871	0.912	0.021	-0.204	2.631	1.191	0.778	-1.774	1996.2.29	Thu	105.069
199603	5.311	-0.091	-1.693	0.022	0.204	-4.541	1.193	0.743	1.744	1996.3.29	Fri	106.917
199604	4.615	5.270	0.245	0.022	0.615	9.547	1.201	0.814	-2.151	1996.4.30	Tue	104.642
199605	-1.857	0.559	0.418	0.028	0.199	1.346	1.203	0.824	3.221	1996.5.31	Fri	108.067
199606	1.910	-2.084	0.913	0.029	-0.299	3.562	1.199	0.854	1.290	1996.6.28	Fri	109.47
199607	-7.474	-1.854	-0.672	0.027	-0.110	-0.882	1.198	0.846	-2.526	1996.7.31	Wed	106.739
199608	-2.573	1.486	1.341	0.027	-0.099	-4.235	1.197	0.810	1.810	1996.8.30	Fri	108.689
199609	5.742	-3.585	0.985	0.027	0.409	1.561	1.202	0.823	2.692	1996.9.30	Mon	111.655
199610	-4.730	1.060	0.077	0.026	0.199	-2.115	1.204	0.806	1.921	1996.10.31	Thu	113.821
199611	0.793	-3.847	3.454	0.026	-0.298	8.723	1.201	0.876	0.001	1996.11.29	Fri	113.822
199612	-5.859	-4.959	1.030	0.022	0.100	6.458	1.202	0.933	1.700	1996.12.31	Tue	115.774
199701	-6.688	-0.886	3.116	0.023	-0.100	1.206	1.201	0.944	4.636	1997.1.31	Fri	121.268
199702	1.339	-1.982	-1.373	0.022	-0.209	0.448	1.198	0.948	-0.369	1997.2.28	Fri	120.821
199703	-0.779	-2.789	-1.200	0.020	0.110	4.713	1.199	0.993	2.362	1997.3.31	Mon	123.709
199704	4.947	-4.790	2.409	0.022	2.026	9.617	1.224	1.088	2.535	1997.4.30	Wed	126.885
199705	3.174	4.491	-1.732	0.027	0.196	1.182	1.226	1.101	-8.597	1997.5.30	Fri	116.433
199706	4.515	-0.143	-0.093	0.027	0.000	-0.653	1.226	1.094	-1.574	1997.6.30	Mon	114.615
199707	-0.626	-7.306	-3.973	0.029	-0.400	7.964	1.221	1.181	3.226	1997.7.31	Thu	118.373
199708	-7.505	0.540	-2.256	0.031	0.098	-8.861	1.222	1.076	1.980	1997.8.29	Fri	120.74
199709	-2.470	-12.180	-4.416	0.030	0.695	2.598	1.231	1.104	-0.035	1997.9.30	Tue	120.698
199710	-8.009	7.050	2.699	0.027	0.301	-3.045	1.235	1.071	-0.295	1997.10.31	Fri	120.342
199711	-1.946	-9.750	-5.927	0.028	-0.688	8.532	1.226	1.162	5.870	1997.11.28	Fri	127.617
199712	-6.139	-5.185	-4.836	0.022	-0.195	-7.300	1.224	1.077	2.242	1997.12.31	Wed	130.511
199801	7.876	15.699	6.044	0.026	-0.108	-1.532	1.222	1.061	-2.751	1998.1.30	Fri	126.966
199802	0.410	5.130	2.794	0.028	-0.098	-5.468	1.221	1.003	-0.695	1998.2.27	Fri	126.086
199803	-1.115	-2.863	-0.635	0.028	0.402	9.034	1.226	1.093	5.531	1998.3.31	Tue	133.263
199804	-2.294	-2.336	-0.226	0.033	0.195	0.713	1.228	1.101	-0.723	1998.4.30	Thu	132.304
199805	-0.118	-2.103	1.382	0.029	0.302	2.219	1.232	1.125	4.587	1998.5.29	Fri	138.509
199806	0.746	0.849	0.022	0.032	-0.398	-5.848	1.227	1.060	-0.144	1998.6.30	Tue	138.309
199807	2.577	-0.678	0.352	0.030	-0.595	7.301	1.220	1.137	4.447	1998.7.31	Fri	144.603
199808	-12.314	-3.874	-3.680	0.030	-0.098	4.540	1.219	1.189	-2.606	1998.8.31	Mon	140.879
199809	-5.332	-3.071	-0.719	0.014	0.796	14.278	1.228	1.358	-3.078	1998.9.30	Wed	136.608
199810	-0.762	0.041	0.478	0.009	0.692	-7.351	1.237	1.259	-15.847	1998.10.30	Fri	116.591
199811	10.423	4.088	0.764	0.013	-0.097	-3.570	1.236	1.214	5.311	1998.11.30	Mon	122.954
199812	-4.914	-2.202	-3.534	0.021	-0.397	11.031	1.231	1.347	-8.483	1998.12.31	Thu	112.952
199901	3.527	-0.395	1.305	0.019	-0.486	-3.761	1.225	1.297	2.621	1999.1.29	Fri	115.954
199902	-0.444	1.527	0.752	0.010	-0.401	-6.000	1.220	1.219	2.260	1999.2.26	Fri	118.6
199903	13.663	2.367	0.161	0.008	0.098	-11.074	1.221	1.084	-0.194	1999.3.31	Wed	118.369
199904	5.517	7.379	8.005	0.004	0.500	-5.119	1.227	1.028	0.850	1999.4.30	Fri	119.38
199905	-2.985	-0.317	0.218	0.003	0.000	8.628	1.227	1.117	1.563	1999.5.31	Mon	121.26
199906	9.178	-6.212	-9.642	0.003	-0.293	5.780	1.224	1.182	-0.256	1999.6.30	Wed	120.95
199907	4.432	-2.154	-1.145	0.002	-0.401	-1.236	1.219	1.167	-5.367	1999.7.30	Fri	114.63
199908	-1.472	0.010	1.718	0.002	0.295	-4.361	1.222	1.116	-4.798	1999.8.31	Tue	109.26
199909	3.687	-3.410	-2.525	0.002	0.304	2.882	1.226	1.148	-2.249	1999.9.30	Thu	106.83
199910	3.788	-3.411	-5.104	0.002	0.195	-12.319	1.228	1.007	-2.454	1999.10.29	Fri	104.24
199911	4.968	-7.294	-7.403	0.003	-0.594	-4.282	1.221	0.964	-2.388	1999.11.30	Tue	101.78
199912	4.934	-7.220	-7.352	0.005	-0.294	1.247	1.218	0.976	0.353	1999.12.31	Fri	102.14

Table 34 U.S.Data (Percent per Month, except for ipi)[a]

	usmkt	ussmb	ushml	ustb	uscpi	usinfl[b]	ipi_uscpi	ipi_usinfl
mean	1.357	0.200	0.153	0.482	0.266	0.482		
std dev	4.240	2.674	2.722	0.156	0.202	6.802	1.000	1.000
198305	1.310	6.300	1.650	0.690	0.302	5.353	1.003	1.054
198306	3.800	0.990	3.860	0.670	0.402	3.797	1.007	1.014
198307	3.170	1.540	5.550	0.740	0.300	1.013	1.010	1.024
198308	0.360	4.390	5.220	0.760	0.499	11.871	1.015	1.145
198309	1.630	0.630	1.110	0.760	0.298	4.583	1.018	1.093
198310	2.770	3.570	4.960	0.760	0.198	6.300	1.020	1.024
198311	2.930	1.890	0.820	0.700	0.099	3.374	1.021	0.989
198312	1.040	0.310	1.750	0.730	0.592	5.552	1.027	0.935
198401	1.290	0.470	7.770	0.760	0.491	11.543	1.032	1.042
198402	3.920	1.690	3.420	0.710	0.195	7.280	1.034	1.118
198403	1.350	0.040	0.220	0.730	0.487	2.984	1.039	1.085
198404	0.270	1.050	1.360	0.810	0.291	0.743	1.042	1.077
198405	5.240	0.040	0.520	0.780	0.290	4.365	1.045	1.124
198406	2.360	0.160	2.320	0.750	0.386	10.208	1.049	1.009
198407	2.040	2.410	0.240	0.820	0.384	16.112	1.053	0.847
198408	11.260	0.180	1.630	0.830	0.478	10.295	1.058	0.934
198409	0.040	0.150	5.220	0.860	0.286	2.743	1.061	0.959
198410	0.010	1.120	0.520	1.000	0.000	6.052	1.061	0.901
198411	1.060	0.590	3.980	0.730	0.000	5.978	1.061	0.955
198412	2.380	0.570	0.270	0.640	0.190	7.729	1.064	0.881
198501	8.570	3.200	4.970	0.650	0.474	0.527	1.069	0.877
198502	1.700	0.870	0.050	0.580	0.377	0.309	1.073	0.874
198503	0.190	1.000	3.880	0.620	0.470	16.854	1.078	1.021
198504	0.220	0.050	3.580	0.720	0.374	1.281	1.082	1.034
198505	5.590	2.490	0.910	0.660	0.280	4.029	1.085	0.993
198506	1.720	0.580	0.430	0.550	0.186	2.074	1.087	0.972
198507	0.050	2.930	1.960	0.620	0.186	0.597	1.089	0.966
198508	0.480	0.420	2.080	0.550	0.278	2.983	1.092	0.995
198509	3.970	1.840	1.130	0.600	0.369	5.078	1.096	0.945
198510	4.460	1.610	0.550	0.650	0.276	2.751	1.099	0.971
198511	6.930	0.150	2.920	0.610	0.275	0.982	1.102	0.961
198512	4.310	0.690	1.630	0.650	0.274	0.227	1.105	0.959
198601	0.990	1.170	0.110	0.560	0.274	5.321	1.102	0.908
198602	7.280	0.590	0.690	0.530	0.457	12.679	1.097	0.793
198603	5.390	0.530	0.190	0.600	0.184	0.178	1.095	0.794
198604	0.800	2.910	2.740	0.520	0.276	0.970	1.098	0.786
198605	5.080	1.180	0.070	0.490	0.551	3.108	1.104	0.811
198606	1.430	0.990	1.430	0.520	0.000	4.063	1.104	0.844
198607	5.970	3.400	4.450	0.520	0.183	11.518	1.106	0.941
198608	6.640	4.290	3.460	0.460	0.456	13.402	1.111	1.067
198609	7.910	2.520	3.420	0.450	0.091	1.941	1.112	1.046
198610	4.940	2.490	1.460	0.460	0.091	2.927	1.113	1.016
198611	1.520	2.140	0.420	0.390	0.091	4.060	1.114	1.057
198612	2.650	0.060	0.210	0.490	0.633	0.086	1.121	1.056

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[a]Symbols "usmkt, ussmb, ushml, ustb, uscpi, usinfl" correspond, respectively, to those in C-R-X (Table 1 for the three Fama-French factor returns, p.262) as follows: US market, US SMB, US HML, US Tbill, US CPI inflation, US extracted risk-free rate $\hat{R}^U_{ft}$. Symbols "ipi_uscpi, ipi_usinfl" correspond, respectively, to those in Table 1 in the present paper as follows: $P^U_{CPI,t}$, $P^U_{R,t}$.

[b]This represents "extracted Rf."

Table 34 (Continued)

	usmkt	ussmb	ushml	ustb	uscpi	usinfl	ipi_uscpi	ipi_usinfl
198701	12.840	1.670	2.950	0.420	0.360	10.781	1.125	1.170
198702	4.760	3.700	5.740	0.430	0.448	1.926	1.130	1.147
198703	2.320	0.410	1.720	0.470	0.535	9.054	1.136	1.251
198704	1.710	1.660	0.430	0.440	0.355	1.849	1.140	1.274
198705	0.520	0.450	0.420	0.380	0.354	4.545	1.144	1.216
198706	4.380	2.130	1.080	0.480	0.264	3.573	1.147	1.173
198707	4.410	0.520	1.180	0.460	0.527	6.995	1.153	1.255
198708	3.720	0.860	1.150	0.470	0.524	3.884	1.159	1.304
198709	2.070	0.530	0.060	0.450	0.261	6.074	1.162	1.383
198710	22.490	8.400	4.170	0.600	0.087	3.363	1.163	1.337
198711	7.290	2.650	3.090	0.350	0.000	6.822	1.163	1.428
198712	7.040	0.060	4.300	0.390	0.260	0.985	1.166	1.414
198801	4.530	0.680	5.390	0.290	0.259	6.455	1.169	1.322
198802	5.160	3.280	1.620	0.460	0.431	2.793	1.174	1.285
198803	1.710	6.210	0.930	0.440	0.515	4.982	1.180	1.349
198804	1.100	0.830	1.520	0.460	0.342	1.196	1.184	1.333
198805	0.090	2.700	2.400	0.510	0.426	3.342	1.190	1.378
198806	5.140	1.870	1.450	0.490	0.424	7.587	1.195	1.273
198807	0.720	0.240	2.270	0.510	0.422	1.844	1.200	1.297
198808	2.790	0.060	1.960	0.590	0.672	1.285	1.208	1.280
198809	3.730	1.250	0.750	0.620	0.334	3.973	1.212	1.229
198810	1.770	2.730	1.730	0.610	0.083	3.343	1.213	1.270
198811	1.640	1.620	1.350	0.570	0.166	5.026	1.215	1.334
198812	2.080	1.890	1.420	0.630	0.498	0.750	1.221	1.324
198901	6.590	2.210	0.470	0.550	0.413	2.079	1.226	1.352
198902	1.640	2.850	0.800	0.610	0.576	0.989	1.233	1.365
198903	2.150	0.590	0.330	0.670	0.654	0.178	1.241	1.363
198904	4.850	0.650	1.670	0.670	0.569	3.673	1.248	1.413
198905	3.960	0.080	1.050	0.790	0.242	0.769	1.251	1.424
198906	0.500	0.960	2.180	0.710	0.242	1.721	1.254	1.448
198907	7.770	4.190	2.960	0.700	0.161	0.518	1.256	1.456
198908	2.230	0.520	0.670	0.740	0.321	3.959	1.260	1.398
198909	0.180	0.340	1.320	0.650	0.480	4.372	1.266	1.459
198910	2.940	3.010	0.880	0.680	0.239	2.978	1.269	1.503
198911	1.790	1.280	1.230	0.690	0.159	11.370	1.271	1.673
198912	1.830	2.310	0.180	0.610	1.031	10.063	1.284	1.842
199001	7.010	1.000	1.000	0.570	0.471	0.409	1.290	1.849
199002	1.500	0.910	0.530	0.570	0.547	5.914	1.297	1.740
199003	2.420	1.250	2.850	0.640	0.155	3.956	1.299	1.671
199004	2.820	0.230	2.310	0.690	0.233	8.013	1.302	1.537
199005	8.890	2.900	3.870	0.680	0.542	7.786	1.309	1.657
199006	0.430	1.160	2.220	0.630	0.385	0.608	1.315	1.667
199007	0.940	2.900	0.360	0.680	0.920	12.618	1.327	1.877
199008	9.170	3.340	1.630	0.660	0.836	6.671	1.338	2.003
199009	5.390	3.620	0.560	0.600	0.603	9.654	1.346	2.196
199010	1.240	5.760	0.270	0.680	0.225	2.373	1.349	2.144
199011	6.600	0.270	2.870	0.570	0.000	3.458	1.349	2.070
199012	2.950	0.880	1.630	0.600	0.598	1.946	1.357	2.110
199101	4.900	3.580	1.650	0.520	0.149	18.246	1.359	1.725
199102	7.570	3.870	0.980	0.480	0.148	8.059	1.361	1.864
199103	2.880	3.900	1.220	0.440	0.148	5.923	1.363	1.974
199104	0.340	0.620	1.520	0.530	0.296	4.160	1.367	1.892
199105	4.060	0.370	0.560	0.470	0.295	12.359	1.371	1.658
199106	4.420	0.190	1.200	0.420	0.147	0.263	1.373	1.663
199107	4.680	0.720	1.000	0.490	0.294	5.524	1.377	1.755
199108	2.690	1.530	0.980	0.460	0.439	0.556	1.383	1.745
199109	1.110	1.570	1.130	0.460	0.146	8.386	1.385	1.891
199110	1.770	0.850	0.610	0.420	0.291	3.214	1.389	1.952
199111	3.730	0.470	1.800	0.390	0.073	5.618	1.390	2.062
199112	10.700	2.400	3.900	0.380	0.145	3.220	1.392	1.995

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Table 34 (Continued)

	usmkt	ussmb	ushml	ustb	uscpi	usinfl	ipi_uscpi	ipi_usinfl
199201	0.170	8.490	4.720	0.340	0.362	14.654	1.397	1.703
199202	1.330	0.780	6.620	0.280	0.505	7.622	1.404	1.573
199203	2.370	1.350	3.590	0.340	0.144	5.817	1.406	1.482
199204	1.380	6.030	4.370	0.320	0.143	3.641	1.408	1.428
199205	0.650	0.420	1.170	0.280	0.358	9.303	1.413	1.560
199206	1.920	3.160	3.280	0.320	0.214	1.627	1.416	1.586
199207	3.990	0.280	0.400	0.310	0.285	4.842	1.420	1.663
199208	2.080	0.040	1.070	0.260	0.284	3.073	1.424	1.714
199209	1.240	0.710	0.090	0.260	0.354	2.373	1.429	1.673
199210	1.090	2.220	2.060	0.230	0.141	10.332	1.431	1.500
199211	4.020	3.390	1.660	0.230	0.070	10.680	1.430	1.340
199212	1.760	1.620	2.430	0.280	0.493	2.512	1.438	1.306
199301	1.240	1.980	5.960	0.230	0.351	3.015	1.443	1.267
199302	0.550	3.380	6.430	0.220	0.349	3.616	1.448	1.313
199303	2.500	0.490	1.270	0.250	0.279	4.266	1.452	1.369
199304	2.550	0.490	2.640	0.240	0.139	8.830	1.454	1.248
199305	2.940	2.070	3.480	0.220	0.139	11.607	1.456	1.393
199306	0.510	0.230	2.540	0.250	0.000	0.654	1.456	1.384
199307	0.080	1.000	3.000	0.240	0.277	2.455	1.460	1.418
199308	3.930	0.120	0.040	0.250	0.207	1.499	1.463	1.396
199309	0.060	3.050	0.240	0.260	0.414	6.945	1.469	1.299
199310	1.810	1.780	2.300	0.220	0.069	3.882	1.470	1.350
199311	1.730	1.480	1.010	0.250	0.000	6.029	1.470	1.268
199312	1.940	1.230	0.270	0.230	0.274	9.189	1.474	1.385
199401	3.130	0.310	1.500	0.250	0.342	4.410	1.479	1.324
199402	2.410	2.660	1.460	0.210	0.341	4.005	1.484	1.271
199403	4.570	0.960	1.500	0.270	0.136	1.731	1.486	1.249
199404	0.980	0.860	1.480	0.270	0.068	0.838	1.487	1.259
199405	0.950	2.010	0.330	0.320	0.339	4.079	1.492	1.208
199406	2.740	0.440	1.990	0.310	0.270	4.464	1.496	1.154
199407	3.040	1.730	1.030	0.280	0.404	2.356	1.502	1.181
199408	4.280	1.170	2.850	0.370	0.268	4.386	1.506	1.233
199409	1.870	2.600	1.350	0.370	0.067	7.505	1.507	1.326
199410	1.490	2.420	1.450	0.380	0.134	0.904	1.509	1.314
199411	3.710	0.140	0.020	0.370	0.000	1.967	1.509	1.288
199412	1.280	0.100	0.000	0.440	0.401	0.393	1.515	1.283
199501	2.060	3.030	2.540	0.420	0.399	5.136	1.521	1.349
199502	3.960	0.340	0.110	0.400	0.331	1.827	1.526	1.324
199503	2.700	0.470	1.900	0.460	0.330	4.683	1.531	1.386
199504	2.490	0.460	2.130	0.440	0.197	0.850	1.534	1.374
199505	3.410	2.080	1.750	0.540	0.197	1.127	1.537	1.390
199506	3.080	2.930	2.910	0.470	0.000	3.705	1.537	1.338
199507	4.070	2.200	2.120	0.450	0.262	2.546	1.541	1.304
199508	0.930	1.710	2.040	0.470	0.196	5.945	1.544	1.382
199509	3.640	2.190	0.640	0.430	0.326	11.053	1.549	1.534
199510	1.110	4.120	0.350	0.470	0.065	5.109	1.548	1.613
199511	4.300	0.800	0.230	0.420	0.065	2.231	1.547	1.577
199512	1.540	0.480	1.000	0.490	0.586	6.415	1.556	1.678
199601	2.810	2.550	0.420	0.430	0.324	8.562	1.561	1.822
199602	1.610	2.060	1.970	0.390	0.516	3.064	1.570	1.766
199603	1.120	1.250	1.050	0.390	0.385	6.630	1.576	1.649
199604	2.510	5.020	4.290	0.460	0.192	2.574	1.579	1.606
199605	2.670	3.240	1.460	0.420	0.064	5.587	1.580	1.696
199606	0.770	3.990	2.530	0.400	0.191	3.665	1.583	1.758
199607	5.340	3.590	4.480	0.450	0.191	6.232	1.586	1.649
199608	3.220	2.290	0.620	0.410	0.318	0.178	1.591	1.646
199609	5.300	0.850	3.760	0.440	0.317	1.825	1.596	1.616
199610	1.390	4.030	4.620	0.420	0.190	3.016	1.599	1.664
199611	6.570	3.460	0.090	0.410	0.000	3.416	1.599	1.721
199612	1.140	3.140	1.080	0.460	0.315	0.998	1.604	1.738

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Table 34 (Continued)

	usmkt	ussmb	ushml	ustb	uscpi	usinfl	ipi_uscpi	ipi_usinfl
199701	5.300	1.580	2.540	0.450	0.314	1.580	1.609	1.711
199702	0.090	2.540	4.830	0.390	0.251	3.675	1.613	1.774
199703	4.440	0.320	3.850	0.430	0.125	9.411	1.615	1.607
199704	4.250	5.140	1.200	0.430	0.062	5.911	1.614	1.512
199705	7.130	4.620	4.090	0.490	0.125	7.469	1.616	1.625
199706	4.420	1.360	0.830	0.370	0.125	4.172	1.618	1.557
199707	7.630	2.370	0.690	0.430	0.187	2.141	1.621	1.524
199708	3.650	7.440	0.900	0.410	0.249	4.503	1.625	1.592
199709	5.800	2.580	0.380	0.440	0.248	6.312	1.629	1.693
199710	3.410	0.930	2.530	0.420	0.062	6.623	1.628	1.581
199711	3.110	5.050	1.050	0.390	0.124	3.806	1.626	1.520
199712	1.800	2.330	3.600	0.480	0.186	1.905	1.629	1.549
199801	0.440	1.010	1.670	0.430	0.186	6.519	1.632	1.448
199802	7.280	0.290	1.230	0.390	0.185	5.868	1.635	1.363
199803	5.140	1.450	1.920	0.390	0.185	2.671	1.638	1.400
199804	1.090	0.410	0.220	0.430	0.185	0.880	1.641	1.387
199805	2.550	3.620	4.290	0.400	0.123	12.026	1.643	1.221
199806	3.270	3.400	1.540	0.410	0.123	1.492	1.645	1.239
199807	2.320	4.510	1.790	0.400	0.123	1.379	1.647	1.222
199808	15.680	5.920	5.690	0.430	0.122	8.547	1.649	1.326
199809	6.410	0.020	3.760	0.460	0.244	36.413	1.653	1.809
199810	7.430	3.360	2.850	0.320	0.000	12.470	1.653	1.583
199811	6.170	1.360	3.680	0.310	0.061	1.429	1.652	1.561
199812	6.320	0.310	4.950	0.380	0.244	9.812	1.656	1.714
199901	3.820	1.150	6.160	0.350	0.122	10.532	1.658	1.533
199902	3.800	5.590	1.660	0.350	0.304	14.185	1.663	1.316
199903	3.750	3.820	3.040	0.430	0.727	1.703	1.675	1.338
199904	4.840	2.890	2.800	0.370	0.000	9.206	1.675	1.215
199905	2.050	3.460	3.080	0.340	0.000	12.838	1.675	1.371
199906	5.120	3.420	4.330	0.400	0.301	0.562	1.680	1.363
199907	3.070	2.010	0.700	0.380	0.240	9.380	1.684	1.491
199908	0.960	1.160	1.260	0.390	0.479	9.334	1.693	1.631
199909	2.290	3.230	3.180	0.390	0.179	1.957	1.696	1.662
199910	6.190	6.530	3.190	0.390	0.059	12.519	1.697	1.454
199911	3.560	7.710	8.090	0.360	0.000	10.867	1.697	1.612
199912	8.270	6.980	9.050	0.440	0.297	3.522	1.702	1.669