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Interest Rate and Price Linkages between the USA and Japan: Evidence from the Post-Bretton Woods Period

Katarina Juselius
Ronald MacDonald

Studiestræde 6, DK-1455 Copenhagen K., Denmark
Tel. +45 35 32 30 82 - Fax +45 35 32 30 00
http://www.econ.ku.dk

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Katarina Juselius and Ronald MacDonald
Institute of Economics, University of Copenhagen and Department of Economics, University of Strathclyde

Abstract

In this paper we investigate the extent to which a number of key parity conditions hold within and between the USA and Japan. Previous research has demonstrated that the nonstationarity of the 'simple' parity conditions was related to the nonstationarity of the real exchange rate, reflecting the very slow adjustment to fundamental real exchange rates. The need to finance the resulting trade deficits seemed to have caused similar nonstationary movements in the long-term bond differential. Support for this proposition is also given in this paper. Furthermore, our results point to a reversal of the linkages partly in the term structure from the long to the short end of the market, partly in the Fisher parities from the nominal interest rate to inflation rate. These results might be important for the conduct of monetary policy which works on the economy through short-term interest rates.

JEL Classifications: E31, E43, F31, F32.

Keywords: International Parity Conditions.

1 Introduction

The introduction of a common European currency in 1999 is likely to have significantly changed the working of the international monetary system. But to be able to evaluate the future impact of the new currency area it is important first to have a good empirical understanding of how the international transmission mechanisms worked in the preceding system, i.e. in the post- Bretton Woods system. In this context it is of particular interest to study how the external float of the EMS area worked relative to the dollar and yen areas. Within the EMS area the role of the German mark as the leading currency is widely accepted, as is the choice of Germany to represent the European Economic Community in a variety of empirical studies. This was also the choice in Juselius and MacDonald (2000), hereafter JM, for studying international parity conditions - purchasing power parity (PPP), uncovered interest rate parity (UIP), the term structure (TS), and the Fisher parities - between the USA and Europe from the mid-seventies up to 1998, using cointegration analysis based on monthly data.

Altogether, the results in JM strongly suggested that the international price mechanisms between the USA and Germany in the post Bretton Woods period were quite different from what is usually assumed. An important finding was that the very slow, though significant, price adjustment towards sustainable levels of real exchange rates, had been compensated by corresponding changes in the spread of the long-term bond rates. Related to this was the strong empirical support for the weak exogeneity of the long-term bond rates, signifying the importance of the large US trade deficits (i.e. the low levels of US savings) and, hence, their linkage to international finance. The results pointed to a reversal of the linkages partly in the term structure from the long end of the market to the short end; partly in the Fisher parities from the nominal interest rates to inflation rate.

These results should be of concern for the conduct of monetary policy as it is supposed to be transmitted onto the economy through the short-term interest rates. In the words of Alan S. Blinder, (1997, p. 242):

"*Unfortunately the model miserably fails a variety of empirical tests (see Campbell, 1995). Economists are thus in desperate need of a better model of the term structure. More than academic completeness is at stake here, for the absence of a usable term structure severely handicaps the conduct of monetary policy, which works its will on the economy through short-term rates of interest rates.*"

Clearly if one of the important linkages has changed, prompting for a modified theoretical explanation, then it seems likely that we would find that other relationships do not work according to theory. Since the theoretical basis of international macroeconomics strongly build on these parity conditions, one would expect the international transmission mechanisms (nominal as well as real) to be significantly different from standard theoretical assumptions.

This was also the conclusion in JM where the empirical findings generated a number of new hypotheses that should be confronted against new data. This is the purpose of the present study, where we repeat the basic empirical design adopted in JM for the US-Japanese case. Since the trade between the US and Japan has been more important than the trade between Germany and the US, we believe this case is of particular interest.

By choosing a similar empirical design as in JM we achieve the additional advantage of testing the empirical findings in JM as prior hypotheses. Hence, we will frequently refer to the results of JM when interpreting the empirical results. Nevertheless, we will also try to empirically isolate those relationships which seem unique for the US-Japanese case. As we shall see, there are some important differences between the results for the Japanese-US and the German-US systems.

The outline of the remainder of this paper is as follows. In the next section we present a motivational overview of some international parity relationships which feature prominently in our empirical analysis. Section 3 defines the data and discusses the 'general-to-specific' approach in econometric modelling versus the 'specific-to-general' approach in the choice of information. Section 4 discusses some econometric impactions of analyzing inflation rates and real exchange rates instead of prices and spot exchange rates. Section 5 presents the empirical model of inflation rates, real exchange rates, and bond yields and presents a careful cointegration analysis of the theoretically motivated parities as well as empirically acceptable modifications of them. Section 6 adds the libor rates to the analysis, presents a fully identified structure of empirically acceptable long-run parity relations, and discusses the role of the short-term interest rates in the large model. Section 7 provides an empirical investigation of the common driving trends and the final impact of permanent shocks to the variables of the system. Section 8 presents a parsimoniously parameterized short-run adjustment model for the full system. Section 9 summarizes the findings and concludes.

2 Price adjustment, interest rates and exchange rates: A motivational discussion

The results in JM strongly rejected the stationarity hypothesis of the individual parity conditions, but when allowing the conditions to be interdependent, stationarity was recovered. The important finding was that the nonstationarity of the 'simple' parity relationships was primarily related to the nonstationarity of the real exchange rate and the long-term bond rate differential. A hypothetical explanation is that the lack of (or very, very slow) adjustment to a stationary PPP (Purchasing Power Parity) steady state has contributed to the large trade deficits between US and Europe versus Japan. The financing of the latter has caused the long-term bond yield differential to move in a corresponding nonstationary manner. Thus, the theoretical assumption of stationary parity conditions appeared to be a special case of a more general formulation allowing for persistent deviations from steady-state and, hence, market failure in a simple model framework.

JM demonstrated that the stationarity assumption underlying the theoretical parity conditions is consistent with two common stochastic trends, a general price trend reflecting permanent shocks to demand and supply, and a trend reflecting differences in policy between the two countries. This simple model did not obtain any empirical support and the theoretical assumption of two common driving trends had to be replaced by the empirically consistent finding of four common trends. Hypothetically the two additional trends were assumed to be: (1) a trend describing relative national savings behavior and (2) a 'safe haven' trend capturing the role of the dollar as a world reserve currency.

The purpose of the graphical illustration of the parity conditions below is to demonstrate visually the co-movements (or the lack of them) between the theoretical determinants of these relationships. For example, if real exchange rates and the spread of the bond yield move closely together over time, then this is an indication that they share one common stochastic trend. If not, then they are probably influenced by one or several other stochastic trends, for example a trend captured by the inflation spread or the spread between short-term interest rates. For a discussion of the connection between common stochastic trends and cointegration, see Juselius (1999).

The strong-form of purchasing parity condition (PPP) is captured by:

$$p_t - p_t^* - s_t = ppp_t, \tag{1}$$

where p_t is the log of the domestic price level (CPI), p_t^* is the log of the foreign price level (CPI), and s_t denotes the log of the spot exchange rate measured as Yen/\$. Strong-form PPP requires that the ppp_t term (or, equivalently, the real exchange rate) is a stationary steady state relation. This implies that the spot exchange rate should mirror the development of relative prices. Figure 1, upper panel, illustrates the development of relative prices as compared to the spot exchange rate.

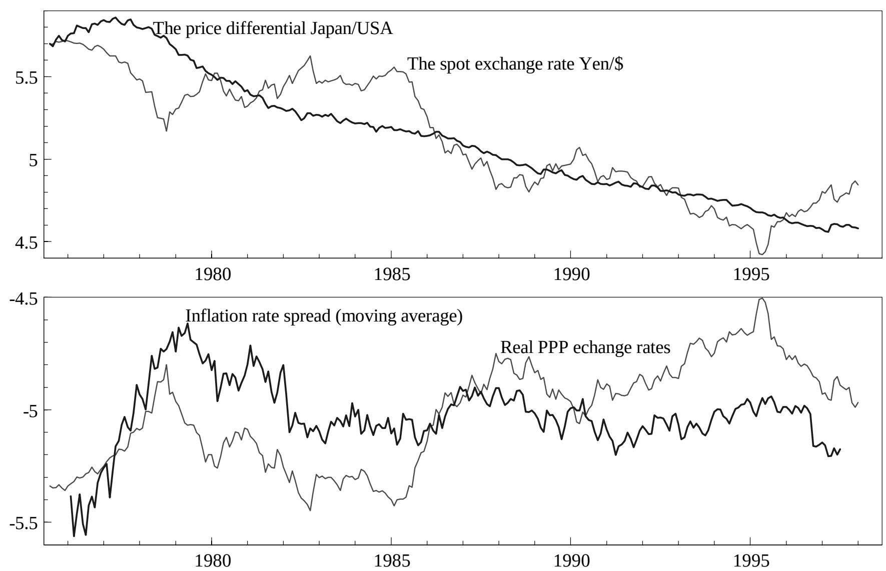

Figure 1: The Jp-US price differential and the spot exchange rate (upper panel) and the ppp term and the inflation rate differential (lower panel.

It appears that the spot exchange rate is much more volatile than the price differential and, hence, is more likely to reflect speculative behavior in the capital market rather than relative price behavior in the goods market. The subsequent empirical results will support this interpretation. See also Krugman (1993) for a theoretical explanation. The *ppp* term graphed in Figure 1, lower panel, exhibits distinctly nonstationary behavior, suggesting insufficient market adjustment. If adjustment towards fundamental PPP takes place exclusively in the goods market we would have the following adjustment relations, either in home inflation:

$$\Delta p_t = \omega_1 \Delta p_t^* + \omega_2 \Delta s_t - \omega_3 ppp_{t-1} + v_t, \tag{2}$$

or in foreign inflation:

$$\Delta p_t^* = \omega_4 \Delta p_t + \omega_5 \Delta s_t + \omega_6 ppp_{t-1} + v_t, \tag{3}$$

or in the spot exchange rate:

$$\Delta s_t = \omega_7 \Delta(p_t - p_t^*) + \omega_8 ppp_{t-1} + v_t. \tag{4}$$

A connection between PPP, from the goods market, and uncovered interest rate parity (UIP), from the capital market, can be derived through the expected exchange rate:

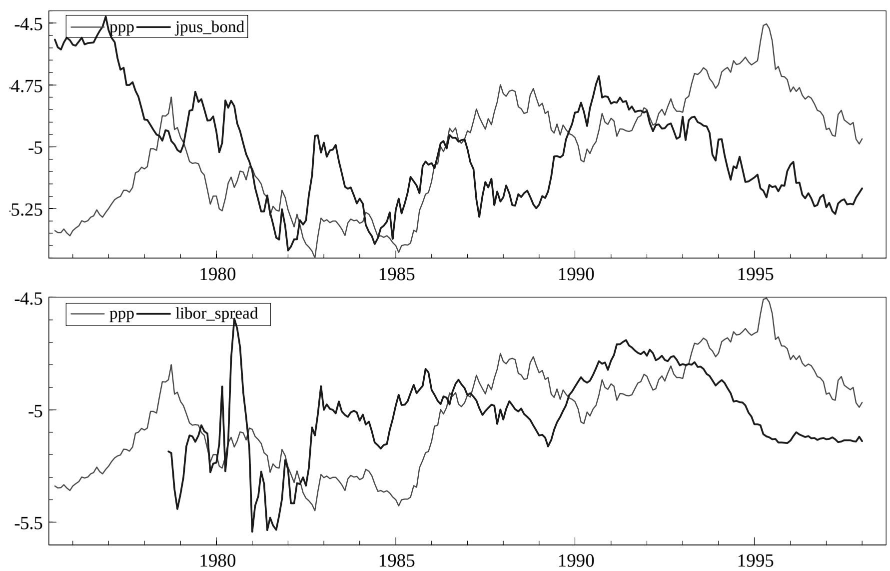

Figure 2:

$$E_t \Delta_l s_{t+l} = i_t^l + i_t^{l*}. \tag{5}$$

By assuming that (4) is used as a predictor of the future change in exchange rate and that $E_t(\Delta_l p - \Delta_l p^*)_{t+1} = (\Delta_l p - \Delta_l p^*)_t$, we get:

$$E_t \Delta_l s_{t+l} = \omega_7 \Delta_l (p_t - p_t^*) + \omega_8 (p_t - p_t^* - s_t), \tag{6}$$

i.e. the expected change in the future spot exchange rate is a function of the present deviation from the *ppp* term and the inflation rate differential at time t. By inserting (6) in (5) we get an expression combining the two parity conditions:

$$(i_t^l - i_t^{l*}) = \omega_7 \Delta_l (p_t - p_t^*) + \omega_8 (p_t - p_t^* - s_t) + v_t, \tag{7}$$

where $\omega_7 = 1$ delivers a relationship between real interest rate parity and the *ppp* term.

The condition for v_t to be stationary in (7), when the *ppp* term is nonstationary, is that the stochastic trend movements in *ppp* are counteracted by similar movement in either the bond rate differential, the inflation rate differential, or both. To illustrate this we have graphed the *ppp* term with the inflation rate differential in Figure 1, lower panel and with the 10 year bond rate differential in Figure 2, upper panel. From (2) or (3) we see that the *ppp* term and the JP-US inflation rate

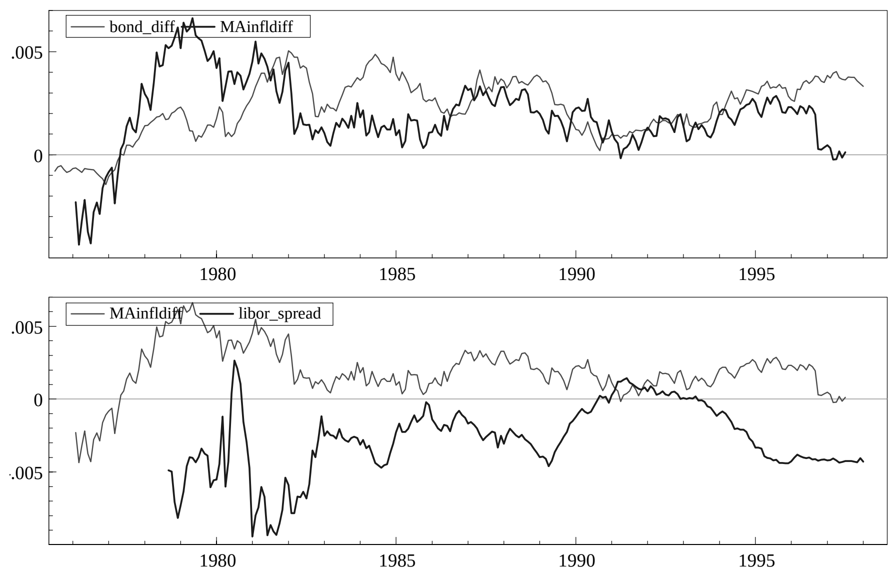

Figure 3: The monthly bond spread and inflation differential (upper panel) and the libor spread and inflation spread (lower pnel). The inflation differential is given as a yearly moving average.

differential should move in opposite directions to secure steady-state behavior. When this is not the case (7) shows that one would expect the JP-US bond rate differential to move in the same direction as the *ppp* term. The graphical display demonstrates that this is obviously not the case: the inflation rate differential seems to move much more in line with the *ppp* term, instead of in the opposite direction. Similarly, the *ppp* term and the bond rate differential do not seem to move as closely together as in the US-German case. In particular, at the beginning of the sample period the bond spread and the *ppp* term seem to have moved dramatically apart. The subsequent empirical analysis will pick it up as a transition towards a more market determined steady-state (see Section 6.1, figure 7).

From a visual inspection, the market adjustment of price inflation and long-term interest rate between the USA and Japan does not seem to have been very effective. This is especially so when compared with the US-German system reported in JM.

With similar arguments as for the long-term bond rate we get the following relationship between the short spread, the inflation rate spread and the *ppp* term:

$$i_t^s - i_t^{s*} = \omega_9(\Delta p_t - \Delta p_t^*) + \omega_{10}(p_t - p_t^* - s_t) + v_t.$$

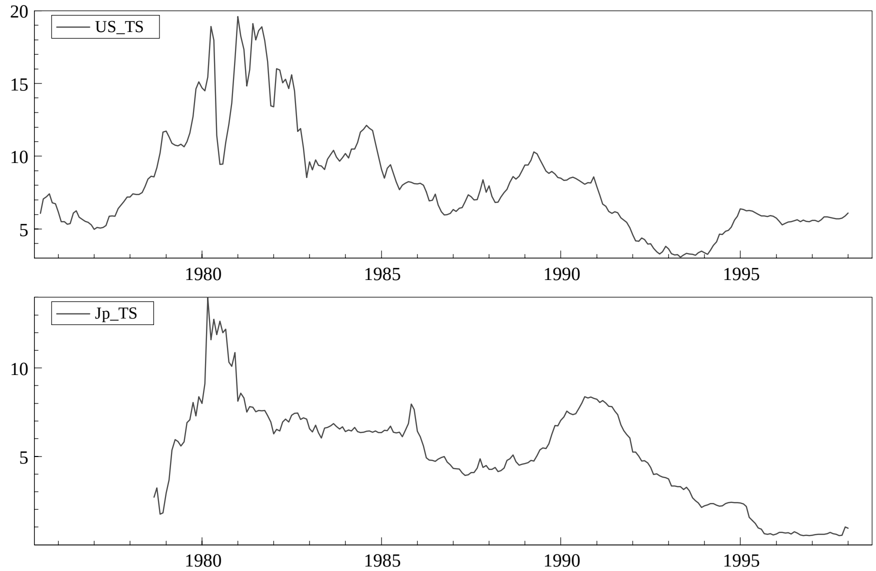

Figure 4: The term structure (libor-bond rate) for US (upper panel) and for Japan (lower panel).

In figure 2, lower panel, we have graphed the *ppp* term together with the Japanese-US libor spread starting from November 1978, the first available date. We notice the large fluctuations in the spread in the first few years. But even after the first turbulent years, the graphs do not exhibit strong co-movements.

The term structure relationship between short- and long-term interest rates predicts that the short rates 'drive' long rates. The standard expectations model of the term structure implies that the term spread (TS), or yield gap, should be stationary. The TS is defined as:

$$i_t^l - i_t^s = v_t. \tag{8}$$

However, in the context of an open economy, the stationarity assumption of (8) is based on the *ceteris paribus* assumption of stationary *ppp* exchange rates (see, for example, JM). This is not the case here and the stationarity of the TS may not hold empirically for our sample period as figure 4 seems to suggest.

Finally, the Fisher condition states that the real interest rates r_t are constant or at least stationary:

$$r_t = i_t - E_t \Delta p_{t+1} + v_t. \tag{9}$$

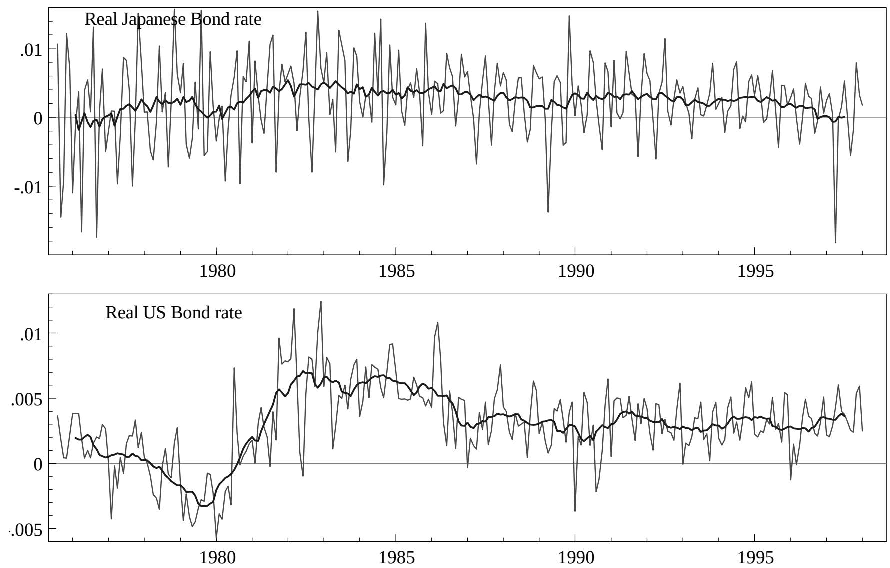

Figure 5:

Again, the stationarity of (9) is based on the *ceteris paribus* assumption of a stationary real exchange rate. When this is not the case real interest rates may not be stationary and it is an empirical question whether real interest rate parity holds between the home and foreign country or whether the parity condition is related to the *ppp* term:

$$(i_t - E_t\Delta p_{t+1}) - (i_t^* - E_t\Delta p_{t+1}^*) + \omega_{11}ppp + v_t. \tag{10}$$

Figure 5 show the graphs of the real US and Japanese long-term bond rates. To see the trend movements more clearly we have also graphed the yearly moving average component. The real rates of the two countries exhibit quite different trending behavior over the sample period and it seems very unlikely that v_t in (10) could be stationary unless $\omega_{11} \neq 0$.

Altogether the graphical presentation contained in this section suggests that the international parity relationships between the USA and Japan are not immediately transparent. The econometric analysis of the parities contained in succeeding sections should shed further light on these relationships.

3 The econometric approach

The interdependence of the parities requires a joint modelling approach of all the relevant variables, x_t, altogether seven in the final 'large' model.

The statistical approach is based on the 'general-to-specific' principle discussed in Hendry and Mizon (1993) and Juselius (1993) and starts from an unrestricted VAR model with a constant term, μ, seasonal dummies, S_t, and intervention dummies, D_t:

$$\begin{gathered}\Delta^2 x_t = \Gamma_1 \Delta^2 x_{t-1} + \Gamma \Delta x_{t-1} + \Pi x_{t-2} + \mu + \Phi_1 S_t + \Phi_2 D_t + \varepsilon_t, \\ \varepsilon_t \sim N_p(0, \Sigma\ ),\ t = 1, ..., T\end{gathered} \tag{11}$$

where $\{\Gamma_1, \Gamma, \Pi, \mu, \Phi, \Sigma\}$ are unrestricted. In this form the model is heavily overparametrized and represents a convenient way of describing the covariances of the data rather than a meaningful economic model. The idea is to sequentially test and impose data consistent restrictions on the VAR (such as reduced rank restrictions, zero parameter restrictions, and other linear or nonlinear parameter restrictions) so that the end result is a more parsimonious model with economically interpretable coefficients. If the VAR can be shown to be an adequate description of the chosen data, then the final model will decompose the covariances of the data into a systematic part, i.e. the part that can be anticipated given the information set, and an unsystematic part, i.e. the part that is unanticipated given the information set. Hence, the strength of the final conclusions rely crucially on the adequacy of the VAR to satisfactorily describe the data in terms of constant parameters, innovation errors, etc.

In the present case x_t is a vector of monthly variables observed for $t =$ *1975:07-1998:1*[1] except for the Japanese libor rate which is observable for 1978:9-1998:1. It is defined by:

p_t = the Japanese, or 'home', price index (CPI),
p_t^*= the US, or 'foreign', price index (CPI),
i_t^l = the Japanese 10 year bond yield,
$i_t^{l,*}$= the US 10 year bond yield,
s_t= the spot exchange rate, defined as Yen/\$,
i_t^s = the Japanese libor rate[2],
$i_t^{s,*}$= the US libor rate.

All of the data used in this study have been extracted from the International Monetary Funds CD-Rom disc (December 1998). Both price series are Consumer Prices (line 64), the long interest rates are 10 year bond yields (line 61), the short rates are libor bill rates (line ??), and the exchange rate is the end of period rate (line ae). All variables, apart from the interest rates (which appear as fractions) are in natural logarithms.

[1] The sample period starts a few years later than for the US-German case, because of the unavailability of Japanese libor rate for the first three years.

[2] The libor rates are chosen here, instead of the 3 months treasury bill rates in JM, because they were available over a longer sample period.

The graphs of the variables in levels and in differences are given in Appendix B. The normality assumption underlying (11) is clearly not satisfied for many of the marginal processes, as the graphs of the differenced variables demonstrate. This is particularly the case for short-term interest rates, signifying the many monetary interventions in this period. To secure valid statistical inference we need to control for intervention effects that fall outside the normality confidence bands. If a residual larger than $|3.3\sigma_{\varepsilon}|$ corresponds to a known intervention, we include it in the information set as a dummy variable.

The analysis of a system of seven variables is econometrically quite demanding and, as in JM a specific-to-general approach in the choice of variables will be adopted: real exchange rates, inflation rates, and long-term bond rates will first be analyzed in a five-dimensional system. When the smaller system is understood, it will be extended by the inclusion of three months treasury bill rates and this larger system will then be thoroughly analyzed. The motivation for first including the long-term bond rates rather than the short-term treasury bill rates is because the long-term movements in real exchange rates are likely to be more informative about the long-term than the short-term interest rates.

The advantage of this approach is twofold. First, the identification of long-run relations is facilitated by building upon the cointegration results from the smaller model. Since the cointegration property is invariant to extensions of the information set, the finding of a particular cointegration relationship in a small set of variables, should also hold in an extended data set. Second, the gradual expansion of the information set facilitates an analysis of the '*ceteris paribus*' assumption and its importance for the empirical analysis of the smaller set of variables. By this procedure the impact of the short-term rates on the system is likely to be much more transparent. The change in the cointegration rank and in weak exogeneity status are particularly informative in this respect.

4 Prices versus inflation rates

The first step of the empirical analysis should preferably involve an examination of prices, the spot exchange rate, and long-term bond rates, i.e. a VAR analysis based on the data vector:

$$x_t' = [(s_t, p_t, p_t^*, i_t^l, i_t^{l*}]. \tag{12}$$

As appears from the graphical display in the Appendix, the inflation rates seem to have behaved more like I(1) variables, suggesting that the price variables and, possibly, the spot exchange rates, should be treated as I(2) variables. The formal testing of the hypothesis that x_t is $I(2)$ can be formulated within the VAR models as two reduced rank hypotheses:

$$\Pi = \alpha\beta' \tag{13}$$

$$\alpha'_{\perp}\Gamma\beta_{\perp} = \zeta\eta' \tag{14}$$

where α, β are $p \times r$ and ζ, η are $p - r \times s_1$ matrices (see Johansen (1991) for further details). By solving the first reduced rank problem one gets information about the number of stationary cointegrating relations and by solving the second about the number of second order stochastic trends. The trace tests (see Paolo, 1996, and Rahbek et. al., 1999) showed mixed evidence of I(2) components in the data. However, the roots of the characteristic polynomial provided clear evidence of near unit roots in the differenced part of the process. The latter finding implies that an empirical analysis of (12) using the I(1) procedure would leave at least one, possibly two, (near) unit roots in the model. Because usual χ^2 inference becomes very unreliable in a model with a near unit root, one should either use more appropriate (Dickey-Fuller type) distributions or transform the model to get rid of the extra near unit roots. Since the appropriate software for inference in the I(2) model is not yet available we have chosen the latter approach. This solution was also adopted in JM.

In the case of I(2) prices and I(1) spot exchange rates, the obvious transformation assuming long-run price homogeneity is to use the vector $x'_t = [(p - p^*)_t, s_t, \Delta p_t, i^l_t, i^{l*}_t]$. However, the spot exchange rate was found to be approximately I(2) and long-run price homogeneity was clearly rejected. Similar results were obtained in the US-German case, though the violation of price homogeneity is more serious in the present data set. Considering the persistently large trade imbalances between the USA and Japan, as well as the USA and Europe, these findings are quite interesting and deserve a detailed analysis in their own right. Such an analysis has to be performed in the I(2) model and we leave this for future research. Instead, following JM, we base our subsequent empirical analyses exclusively on inflation rates and the real exchange rates rather than on prices and the spot exchange rate. However, the transformation $p_t - p^*_t - s_t = ppp_t$ can be econometrically problematic if it does not annihilate all I(2) components in the data. Nevertheless, we have chosen to proceed with this choice, but with due attention to potential econometric and interpretational problems associated with the *ppp* transformation.

Table 1: Misspecification tests and characteristic roots

Multivariate tests:					
Residual autocorr. LM_1	$\chi^2(25)$	$=$	17.3	p-val.	0.87
Normality: LM	$\chi^2(10)$	$=$	58.6	p-val.	0.00
Univariate tests:	Δp_t	Δp_t^*	Δi_t^l	Δi_t^{l*}	Δppp_t
ARCH(2)	0.7	**8.6**	**6.7**	**11.0**	0.2
Jarq.Bera(2)	**6.6**	**8.1**	**19.0**	**16.2**	**8.5**
Skewness	0.3	-0.1	0.3	0.3	0.4
Ex. Kurtosis	3.7	3.8	4.4	4.3	3.5
$\hat{\sigma}_\varepsilon \times 100$	0.30	0.10	0.02	0.02	3.00
R^2	0.79	0.43	0.21	0.38	0.22
Eigenvalues of the trace-test	0.46	0.16	0.04	0.02	0.00
The trace test	232 (69)	64 (47)	17 (29)	6 (15)	0 (4)
Modulus of 5 largest roots:					
Unrestricted model	1.00	0.97	0.97	0.76	0.35
$r = 3$	1.00	1.00	0.97	0.76	0.35
$r = 2$	1.00	1.00	1.00	0.77	0.35

5 An empirical model for inflation rates, bond rates and *ppp*

The VAR model (11) is first estimated based on:

$$x_t' = [(\Delta p_t, \Delta p_t^*, ppp_t, i_t^l, i_t^{l*}] \tag{15}$$

To control for the largest intervention outliers we needed to include the following dummies in the model:

$$D_t' = [D78.11_t, DI80.02_t, D80.05_t, D80.07_t, D81.11_t, DI82.01_t] \tag{16}$$

where $Dxx.yy_t = 1$ in 19xx:yy, 0 otherwise, $DIxx.yy_t = 1$ in $19xx.yy$, -1 in $19xx.yy + 1$, 0 otherwise.

5.1 Specification tests and the choice of rank

Table 1 reports some multivariate and univariate misspecification tests as a first check whether the VAR model conditional on (16) is able to adequately describe the variation in the data. A significant test statistic is given in bold face. We also report the estimated eigenvalues of the trace test, as well as the five largest roots of the characteristic polynomial.

The multivariate LM test for first order residual autocorrelation is not significant, whereas multivariate normality is clearly violated. Normality can be rejected as a result of skewness (third moment) or excess kurtosis (fourth moment). Since the properties of the cointegration estimators are more sensitive to deviations from normality due to skewness than to excess kurtosis we report the univariate Jarque-Bera test statistics together with the third and fourth moments around the mean. As is evident, the rejection of normality is essentially due to excess kurtosis, and hence not so serious for the estimation results. The $ARCH(2)$ is a test for second order autoregressive heteroscedastisity and is rejected for all equations except for Japanese inflation and the US bond rate. Again cointegration estimates are not very sensitive to $ARCH$ effects (Rahbek, 1999). The R^2 measures the improvement in explanatory power relative to the random walk hypothesis, i.e. $\Delta x_t = \varepsilon_t$. They show that with this information set we can explain quite a large proportion of the variation in the inflation rates, but to a much lesser extent the variation in the bond rates and the real exchange rate.

The cointegration rank can be seen as an indication of how well markets adjust and, therefore, of market barriers. Both the trace test and the roots of the characteristic polynomial support the choice of $r = 2$ and, hence, $p - r = 3$ common trends. As a sensitivity check, the roots under the choice $r = 3$ are also reported in Table 1. In this case a large root remained in the model supporting the choice of $r = 2$. A similar result was found in JM and the common trends were interpreted to represent a cumulated shocks to demand and supply, a real trend associated with differences in national savings rates, and finally a safe haven trend associated with the reserve currency role of the US dollar.

5.2 Long-run exclusion, stationarity, and weak exogeneity

Table 2 reports three different tests related to the relevance and the role of the individual variables in the model. The test of long-run exclusion (Johansen and Juselius, 1992) investigates whether any of the variables can be excluded from the cointegration space, implying no long-run relationship with the remaining variables. None of the variables can be excluded.

The test of stationarity (Johansen and Juselius, 1992) investigates whether any of the variables are individually stationary by testing if they correspond to a unit vector in the cointegration space. Accepting the hypothesis implies that the variable in question can be considered $I(0)$ for the choice of $r = 2$. None of the variables can be considered stationary over the sample period.

Finally, the test of long-run weak exogeneity (Johansen and Juselius, 1992) investigates the absence of long-run levels feed-back and is formulated as a zero row of α, i.e. the hypothesis that the variable x_i, $i =$

Table 2: Tests of hypothesis about some properties of the system variables

	Δp_t	Δp_t^*	i_t^l	i_t^{l*}	ppp	$\chi^2(\nu)$
Long-run exclusion:	152.9	61.7	21.3	34.3	11.7	$\chi^2(2) = 5.99$
Stationarity:	39.2	39.4	43.9	45.2	44.2	$\chi^2(3) = 7.81$
Long-run weak exogeneity	141.6	25.9	**2.7**	12.4	**3.7**	$\chi^2(2) = 5.99$

$1, ..., p$, does not adjust to the equilibrium errors $\beta_i' x_t$, $i = 1, ..., r$. If accepted, the variable in question can be considered a driving variable in the system: it 'pushes' the system, but is not being 'pushed' by it.

It appears that the Japanese long-term bond rate and the *ppp* exchange rate can be assumed weakly exogenous, but not the US bond rate. This is an interesting result, partly because it differs from the analysis of the US-German data where the US bond rate was clearly found to be weakly exogenous, partly because the US is generally believed to dominate international capital markets. A possible explanation is that the large US trade deficits and the low levels of national savings relative to Japan have made the US economy more dependent on cheap capital imports from Japan than the other way around.

The strong rejection of weak exogeneity for the inflation rates, (similarly rejected in JM) suggests that prices are primarily adjusting to deviations from long-run steady-states, rather than exchange rates[3]. This might seem puzzling considering the float of the Yen/Dollar rate and the general price stickiness in this period. The results suggest that the large fluctuations in exchanges rates cannot be interpreted as movements towards fundamental *ppp* equilibrium, but probably as speculative (or policy induced) movements with little relationship to the fundamentals of the two countries.

Weak exogeneity is not invariant to changes in the information set and more conclusive test results are presented in Section 7 with the extension of the data set. But before moving to the investigation of feedback and dynamic adjustment effects, we will first test the stationarity of the theoretical parities discussed in Section 2.

5.3 Testing the theoretical parities

The hypotheses reported in Table 3 are of the form $\beta = \{H\phi_1, \psi_1\}$, i.e. they test whether a single restricted relation is in $sp(\beta)$, leaving the other two relations unrestricted (Johansen and Juselius, 1992). The table has been divided into four parts: the first part reports tests on the

[3]Because the residual covariances are very small in this model as demonstrated later in Section 9, the conclusion is robust to linear transformations of the model.

Table 3: Cointegration properties of small Japanese model

	ppp	Δp_1	Δp_2	i_b	i_b^*	$\chi^2(v)$	$p.val.$
$\mathcal{H}_1$	0	1	-1	0	0	10.0 (3)	0.02
$\mathcal{H}_2$	0	0	0	1	-1	34.1(3)	0.00
$\mathcal{H}_3$	0	1	0	-1	0	21.6(3)	0.00
$\mathcal{H}_4$	0	0	1	0	-1	18.1(3)	0.00
$\mathcal{H}_5$	0	1	0.1	-1	-0.1	12.6(2)	0.00
$\mathcal{H}_6$	0	1	-1	-0.2	0.2	3.6(2)	**0.16**
$\mathcal{H}_7$	0.148	1	-1	0	0	8.9(2)	0.01
$\mathcal{H}_8$	-0.046	0	0	-1	1	34.1(2)	0.00
$\mathcal{H}_9$	-0.095	1	0	-1	0	21.1(2)	0.00
$\mathcal{H}_{10}$	-0.010	0	1	0	-1	17.2(2)	0.00
$\mathcal{H}_{11}$	-0.004	1	0.16	-1	-0.16	5.07(1)	0.02
$\mathcal{H}_{12}$	0.008	1	-1	-0.4	0.4	6.0(1)	0.01
$\mathcal{H}_{13}$	-0.224	1	3.0	-4.0	1	1.0(1)	**0.32**
$\mathcal{H}_{14}$	-0.214	-0.25	1	-0.75	0	0.3(1)	**0.61**

stationarity of the hypothetical parities as such, the second part report combination of the parities, the third part includes the *ppp* term to the previous tests, and finally the fourth reports slight modifications of the previous tests. $\mathcal{H}_1-$ $\mathcal{H}_{12}$ have been tested for the German case, whereas $\mathcal{H}_{13} - \mathcal{H}_{14}$ are specific for the Japanese data.

$\mathcal{H}_1 - \mathcal{H}_4$ are hypotheses tests on pairs of variables, such as relative inflation ($\mathcal{H}_1$), relative interest rates ($\mathcal{H}_2$) and Fisher parity conditions ($\mathcal{H}_3$ and $\mathcal{H}_4$). These test, therefore, seek to determine if certain key parity conditions introduced in Section 2 are empirically verifiable on their own. Since all of the p-values are less than the 5% critical value, this is not the case. The same result was obtained for the German data.

$\mathcal{H}_5$ and $\mathcal{H}_6$ are tests of variants of real interest rate parity in which full proportionality has not been imposed. Restricting the Japanese nominal interest rate / inflation combination to conform to a strict Fisher parity condition which, in turn, equals a proportion of the US Fisher condition ($\mathcal{H}_5$) does not prove to be a stationary combination. However, restricting the two inflation rates to have unitary coefficients and the nominal interest rates to have equal and opposite signs ($\mathcal{H}_6$) is not rejected, but the estimated coefficient to the interest rate spread is quite small.

$\mathcal{H}_7-$ $\mathcal{H}_{12}$ adds the *ppp* term first to the key parities, then to the combined inflation and interest rate spread. None of the hypotheses can be accepted as stationary. This is contrary to the US-German case where the real interest rate parities when combined with the *ppp* term were accepted as stationary. Combining the five variables in a manner sug-

Table 4: A structural representation of the cointegrating space.

	Eigenvectors β		Weights α		
	$\hat{\beta}_1$	$\hat{\beta}_2$	Eq.	$\hat{\alpha}_1$	$\hat{\alpha}_2$
Δp_t	1.0	$-$ 0.26 (3.0)	$\Delta^2 p_t$	**−1.21** (13.2)	**0.33** (3.2)
Δp_t^*	−0.17 (2.0)	1.0	$\Delta^2 p_t^*$	−0.05 (1.1)	**-0.28** (5.3)
i_t^l	−1.0	-0.74	Δi_t^l	0.00 (0.3)	0.01 (1.6)
i_t^{l*}	1.0	0	Δi_t^{l*}	−0.01 (1.6)	**0.02** (3.1)
$ppp_t^{1)}$	0.41 (4.1)	−0.21 (2.0)	Δppp_t	-0.01 (1.7)	-0.01 (1.5)
$cnst$	-0.002	-0.001			

1) *ppp* has been divided by 100 to avoid very small coefficients

gested by the parity conditions, therefore, does not seem to be sufficient for the US-Japanese case. This might suggests important differences in the market behavior between the USA and the two major currency blocks. This point is underscored in $\mathcal{H}_{13}$ and $\mathcal{H}_{14}$ where we are able to identify inflation relationships with the *ppp* term embedded in a modified parity relationship. We discuss these relationships further in the next section

5.4 Fully specified cointegrating relations

As Table 3 showed we can recover two stationary relationships ($\mathcal{H}_{13}$ and $\mathcal{H}_{14}$) that seem to capture elements of a reaction function of US and Japanese price inflation to deviations from PPP and movements in the long-term bond rates. Having established this we test whether they can be jointly accepted as hypothesis $\mathcal{H}_{15}$:

$$\mathcal{H}_{15}: \quad \beta = \{H_1\varphi_1, H_2\varphi_2\}$$

where the design matrices H_1 and H_2 corresponding to $\mathcal{H}_{13}$ and $\mathcal{H}_{14}$ are defined as:

$$H_1 = \begin{bmatrix} 1\,0\,0\,0 \\ 0\,1\,0\,0 \\ -1\,0\,0\,0 \\ 1\,0\,0\,0 \\ 0\,0\,1\,0 \\ 0\,0\,0\,1 \end{bmatrix}, \; H_2 = \begin{bmatrix} -1 & 0\,0\,0 \\ 1 & 1\,0\,0 \\ 0 & -1\,0\,0 \\ 0 & 0\,0\,0 \\ 0 & 0\,1\,0 \\ 0 & 0\,0\,1 \end{bmatrix}$$

The three overidentifying restrictions were accepted based on a LR test statistic, aymptotically distributed as $\chi^2(2)$, of 0.27 and a p-value of 0.87. The two stationary relations are reported in Table 4. The first vector has been normalized on the Japanese inflation rate and the second on the US inflation rate. The estimates of the β_{ij} coefficients and their asymptotic t-values indicate that all of the freely estimated coefficients are strongly significant, thereby implying that the suggested structure is also empirically identified. The interpretation of the first relation primarily as a Japanese inflation relation is supported by the strongly significant adjustment coefficient exclusively in the Japanese inflation equation. It essentially shows that Japanese inflation is positively related to US inflation with a coefficient of approximately 0.2 and that the adjustment towards the *ppp* term takes place relative to the long-term bond spread. This seems to suggest that the very slow adjustment to fundamental real exchange rates has been facilitated by a corresponding widening of the bond spread. Cf. the graphical display in Figure 2.

The second relation is describing US inflation as adjusting homogeneously towards JP inflation and JP bond rate, and additionally adjusting towards fundamental *ppp*. It is strongly significant in the US inflation equation and in the Japanese inflation and US bond rate equations, demonstrating its importance in particular for the US variables.

Consistent with the weak exogeneity results in Table 2, neither the *ppp* term nor the Japanese bond rate adjusts to the two cointegration relations. A joint test of the weak exogeneity of the *ppp* term and the Japanese bond rate produced a $\chi^2(4)$ statistic of 5.58 with an associated p-value of 0.23. Hence, permanent shocks to these two variables seem to have a long-run impact on the two inflation rates and the US bond rate without similarly being 'pushed' by them.

Finally, we subject our chosen model to a set of Hansen-Johansen recursive stability tests. The graphs in the Appendix C, Figures C1-C3, are based on the recursively calculated $\alpha-$estimates for β fixed at the values given in Table 4. The recursions start from the model estimates based on the first ten years of the sample (1979-1988), which are then reestimated adding one new observation. They seem to indicate a remarkable degree of stability for the coefficients in the cointegrating vectors.

6 The extended Japanese model

We now apply our knowledge of the small model to an analysis of the full model structure as given by the following vector:

$$x_t' = [ppp, \Delta p_t, \Delta p_t^*, i_t^l, i_t^{l*}, i_t^s, i_t^{s*}]. \tag{17}$$

where $t = 1978.9 - 1998.1$[4]. The final parity condition considered in section 2, namely the term structure relationship, can now be analyzed.

As mentioned above, the short-term interest rates were subject to many major interventions in this period. In particular, the short period of monetary targeting in the beginning of the eighties witnessed wild fluctuations in the short-term rates. The following 14 dummy variables[5] were needed for the residuals to approximately satisfy the normality condition $\hat{\varepsilon}_t \leq 3.3\hat{\sigma}_\varepsilon$:

$$D_t' = [D78.11,\ D79.03, D79.08, DI80.02, D80.03, D8005, D80.11, D81.01, \\ DI81.03, D81.11, DI82.01, DI84.12, DI87.09, D95.08,\ D97.04]$$

where $Dxx.yy_t = 1$ in 19xx:yy, 0 otherwise, $DIxx.yy_t = 1$ in $19xx.yy$, -1 in $19xx.yy+1$, 0 otherwise. Nevertheless, even if the residuals were reasonably well-behaved when controlling for these interventions, normality was rejected, primarily due to excess kurtosis. Both libor rates exhibited residual ARCH. As discussed for the smaller system, excess kurtosis and heteroscedastic errors might affect efficiency but should not produce biased inference.

In Table 5 we report the estimated eigenvalues and trace statistics associated with this system. Adding the two treasury bill rates to the data set implies three possibilities regarding the effect on the rank:

- $r = 2$, i.e. the rank is unchanged and the stochastic trends would have increased to $p - r = 5$, implying that the two short-term rates are not cointegrated over the sample period, neither with themselves nor with the inflation rates, bond rates, or *ppp* term. This case does not seem very plausible *a priori*.

- $r = 3$, i.e. $p - r = 4$ and including the short-term interest rates would have introduced one additional stochastic trend. This means that the short-term interest rates can be jointly cointegrated or cointegrated with the remaining variables of the system.

- $r = 4$, i.e. the common stochastic trends remain unchanged, $p-r = 3$ and the short-term interest rates would be fully integrated with long-term interest rates, inflation rates, and the real exchange rates consistent with the theoretical foundations of the parities of Section 2.

The trace test suggests four common stochastic trends and, consequently, three cointegration relations. However, the trace statistic for

[4]Because the Japanese libor rate is only available from 1978.11 the sample period is shorter than for the bond rate model.

[5]In terms of parameters added to the model they correspond to the equivalent of two additional lags.

Table 5: Eigenvalues, trace tests, and characteristic roots .

The trace test:							
$p-r$	7	6	5	4	3	2	1
λ_i	0.48	0.27	0.17	0.08	0.05	0.02	0.01
Q(r)	314	160	87	**42**	21	9	3
Q_{95}	123	93	69	47	29	15	4
Characteristic roots:							
$r=7$	0.97	0.97	0.95	0.95	0.90	0.39	0.39
$r=4$	1.0	1.0	1.0	0.93	0.90	0.38	0.38
$r=3$	1.0	1.0	1.0	1.0	0.92	0.38	0.38
$r=2$	1.0	1.0	1.0	1.0	1.0	0.44	0.44

both $p-r=4$ and $p-r=3$ is quite close to the 95% quantile, suggesting that the theoretically most consistent case $r=4$ might be true. To check the sensitivity of the model to the choice of r we have also calculated the roots of the characteristic polynomial. There are approximately four 'near unit roots' in the unrestricted system plus a large root of 0.90. The choice of $r=4$ leaves two large roots in the model whereas $r=3$ removes the largest roots but, nevertheless, leaves one quite large root, 0.92, in the model. Only for $r=2$, i.e. for $p-r=5$, all large roots disappear from the model. Since this case is not very likely from a theoretical point of view we continue the analysis with $r=3$, although noting that the libor rates are only weakly cointegrated in this system.

Adding variables to the information set can change the previous findings of long-run weak exogeneity. In fact, a change of weak exogeneity status is a sign of changing long-run feedback and is, therefore, of particular interest. If, for example, the short-term rates are driving the long-term bond rates, then including the former in the analysis should change the previous finding of weakly exogenous bond rates and, instead, we would find that the short rates are weakly exogenous. In Table 6 the test results of weak exogeneity are reported. The test statistics, asymptotically distributed as $\chi^2(3)$, indicate that the weak exogeneity results is altered for the *ppp* term, but not for the Japanese bond rate. This result, together with the rejection of weak exogeneity of the short-term interest rates, suggests that monetary policy shocks are transmitted through the exchange market, hence influencing the real exchange rates. Similar results were found in JM, where the weak exogeneity of the *ppp* term in the small system disappeared when the three months treasury bill rates were added.

We interpret this as evidence of the importance of monetary shocks for short-term changes in the exchange rates. We also find it plausible

Table 6: Tests of long-run weak exogeneity

Single tests of:	Δp_t	Δp_t^*	i_t^l	i_t^{l*}	i_t^s,	i_t^{s*}	ppp
$\chi^2(3) = 7.8$	125.5	43.5	**2.5**	19.7	19.0	12.3	8.1

that the fourth stochastic trend describes the cumulative effect of relative monetary intervention shocks between the USA and Japan and that this effect is particularly important in the short-term capital market.

6.1 A fully specified cointegration structure

An advantage of the principle of 'specific-to-general' is that we can keep the two steady-state relations found in the previous section unaltered. Hence, the 'additional' impact of the two new variables on the system will essentially be described by the third cointegrating relation. We first estimated the partially restricted long-run structure $\beta = \{H_1\varphi_1, H_2\varphi_2, \psi\},$, where H_1 and H_2 are the design matrices of the long-run structure of the small model of Section 5 and ψ is an unrestricted cointegration vector. It turned out, however, that the first cointegration relation, the Japanese inflation relation, was much improved by replacing the Japanese long-term bond rate with the Japanese libor rate and we re-estimated β with $H_1\varphi_1$ defined by the design matrix reported below. Given this change in cointegration design, the 'new' unrestricted cointegration relation ψ normalized on the US treasury bill rate became:

$$\psi' = [-0.04, -0.03, -1.01, 1.00, 1.00, -0.85, 0.09]$$

suggesting that this cointegration relation primarily contains information about the libor spread and the bond spread. This led to the following joint hypotheses on the full cointegration structure:

$$\mathcal{H}_1 : \quad \beta = \{H_1\varphi_1, H_2\varphi_2, H_3\varphi_3\}, \tag{18}$$

where the design matrices are defined as:

$$H_1 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}, H_2 = \begin{bmatrix} -1 & 0 & 0 & 0 \\ 1 & 1 & 0 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}, H_3 = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ -1 & 0 \\ 1 & 0 \\ 1 & 0 \\ -1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}.$$

Table 7: A structural representation of the cointegrating space.

Eigenvectors β				Weights α			
Var.	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	Eq.	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\alpha}_3$
Δp_t	**1.0**	$-$**0.50** (7.5)	0	$\Delta^2 p_t$	$-$**1.29** (10.5)	0.09 (0.6)	-0.08 (0.6)
Δp_t^*	$-$**0.37** (7.2)	**1.0** (0)	**0**	$\Delta^2 p_t^*$	$-$**0.43** (6.0)	**-0.65** (7.7)	$-$**0.46** (5.3)
i_t^l	**0**	$-$**0.50**	$-$**1.0**	Δi_t^l	-0.01 (0.6)	0.01 (0.7)	-0.00 (0.3)
i_t^{l*}	**0.37** (7.2)	0	**1.0**	Δi_t^{l*}	-0.01 (0.9)	**0.03** (2.8)	0.00 (0.3)
i_t^s	$-$**0.37** (7.2)	0	**1.0**	Δi_t^s	-0.02 (1.8)	-0.01 (1.2)	$-$**0.05** (4.1)
i_t^{s*}	0	0	**-1.0**	Δi_t^{s*}	-0.01 (0.7)	**0.04** (2.1)	**0.05** (2.6)
$ppp_t^{1)}$	**0.18** (2.5)	**-0.24** (3.0)	0	Δppp_t	-1.0 (0.8)	-0.3 (0.3)	$-$**3.8** (2.5)
cnst.	-0.000	-0.002	0.000				

1) *ppp* has been divided by 100 to avoid reporting small coefficients

The likelihood ratio statistic for testing the eight overidentifying restrictions, asymptotically $\chi^2(8)$, was 1.03 and the structure is clearly acceptable with a p-value of 1.00. In Table 7 the estimated β_{ij} coefficients and their asymptotic standard errors are reported. All of them are strongly significant, thereby implying that the suggested structure is both formally and empirically identified. Furthermore, as the recursive graphs in Appendix D, figures D1-D4, demonstrate, these empirical effects have been remarkably constant over the last 10 years, a total of 120 observations!

The first two vectors are very similar to the two cointegrating relations in the small model, illustrating the invariance of the cointegration property. The main difference occurs in the first vector where the Japanese short rate now enters instead of the long rate. The third vector is written here as:

$$i_t^s - i_t^{s*} = (i_t^l - i_t^{l*}) + v_t \tag{19}$$

We interpret it as an international term structure relationship in which the short interest differential is proportional to the long interest differential.

Figure 6 shows the graphs of the three equilibrium error correction mechanisms. It appears that the third cointegration relation (19) is only weakly mean-reverting, consistent with the rather high root left in the model when $r = 3$. In particular, the period from 1979 to 1984, containing the years of monetary targeting, seems to describe a transition

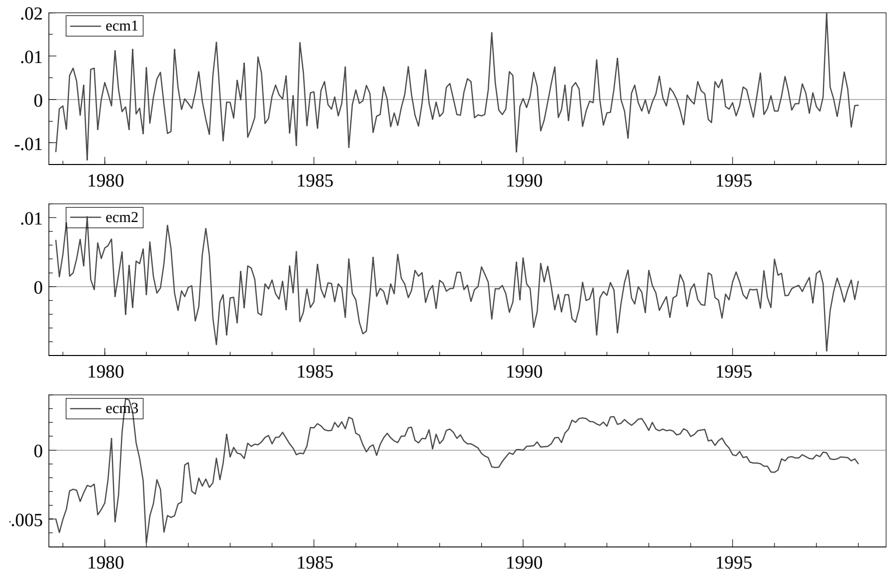

Figure 6: The graphs of the three ecm-mechanisms defined in Table 7.

towards a new steady-state between long- and short-term interest rates. A similar transition behavior can be noticed in the second cointegration relation, the relation between US inflation, Japanese inflation and bond rate, and the *ppp* term. Cf. the graphical inspection in Section 2. The first cointegration relation, the relation between Japanese inflation and libor rate, the US real bond rate and the *ppp*, seems to have defined a remarkably strong cointegrated relationship: the deviations from this relation looks almost as white noise for the whole period.

The adjustment coefficients[6] (t-ratios in brackets) of the first two relations are similar to those of the small system, with the exception that the US inflation rate now significantly adjusts to all three equilibrium errors and Japanese inflation no longer adjusts to the second. It is particularly noteworthy that the two short rates, but not the bond rates, are significantly adjusting to the third steady-state relation, against the expectation's hypothesis which predicts that short-term interest rates drive long-term rates. The small but significant adjustment of US short rate to the second and third cointegrating vector reflects its role as a money market determined interest rate, but the lack of adjustment in the long-term bond rates seems to suggest that the transmission of the

[6]Because the residual covariance matrix is almost diagonal, the interpretation is not likely to change much under a linear transformation of the VAR. Nevertheless, a more complete interpretation will be given in Section 9.

money market effects onto the long-rates is not there, or only weakly so.

How do these results compare with the US-German findings reported in JM? In fact the results are quite similar. For example, JM established a relation towards which German inflation was adjusting, a relation towards which US inflation was adjusting, and an international term structure relationship. Hence, both studies find price adjustment, rather than exchange rate adjustment towards long-run steady-states. Both studies find evidence of the crucial importance of interest rates, both long and short, to facilitate the very slow price adjustment towards fundamental steady-state. Altogether, the results in this section and those contained in JM, can be interpreted as clear evidence of the dependence of the US economy in the post- Bretton Woods period on Japanese (in particular) and European capital markets to finance its large trade deficits.

6.2 The role of short-term interest rates in the Japanese model

To gain a further perspective on the role of the short- relative to the long-term interest rates we report, in Table 8, a comparative analysis of the combined effects as measured by $\hat{\alpha}_r\hat{\beta}_r' = \hat{\Pi}_r$, where the subscript r stands for the restricted estimates as reported in Tables 4 and 7.

Comparing the Japanese inflation rate equations with and without the short-term interest rates demonstrated why the first cointegration relation, $H_1\varphi_1$, needed to be modified when moving from the small to large system: when including the libor rates into the analysis it replaces the Japanese bond rate as a significant determinant of Japanese inflation. Interestingly, the equation for US inflation with the short yields included changes quite dramatically: it is now significantly adjusting to the Japan-US bond spread (positively) and to the libor spread (negatively) and much more significantly to the real exchange rate.

This may suggest that the yield gap plays a different role in this system from the one predicted by the pure expectations model of the term structure. According to the latter the long bond yield should contain information on the current short rate and some average of expected future short rates. In this view, the addition of the short rates should not significantly affect the existing results. The fact that the results are significantly affected suggests that short rates contain important information, over and above that contained in the long rates, and this presumably relates to the operation of monetary policy.

The bond rate equations reported in Table 8 are essentially unchanged with the inclusion of the short rates: the Japanese bond rate is still weakly exogenous and the US bond rate is affected by the inflation rate differential relative to the *ppp* term both in the small and the extended system.

The equation for the *ppp* term in the extended system shows significant additional effects from the short interest rate spread (the central bank policy effect) and the bond rate spread becomes more significant. Most empirical work has failed to find a good description of the variations of nominal exchange rates over this period. Although the significant effects in the *ppp* equation are not large in magnitude it might nevertheless be of interest to reformulate them into an explanation of nominal exchange rate:

$$\Delta ppp_t = -0.04(i^l_{t-1} - i^{l*}_{t-1}) + 0.04(i^s_{t-1} - i^{s*}_{t-1})$$
$$\Delta s_t = \Delta p_t - \Delta p^*_t + 0.04(i^l_{t-1} - i^{l*}_{t-1}) - 0.04(i^s_{t-1} - i^{s*}_{t-1})$$

Thus, it seems as if the Yen/dollar rate has appreciated with US-JP bond rate spread and depreciated with the spread in the short rates. These are interesting results in particular when compared to the German US case, where we found the opposite to hold, i.e. the Dmk/dollar rate appreciated with increasing US-German treasury bill spread, but depreciated with increasing US-German bond spread. One hypothetical explanation is that the Japanese central bank have bid up the libor rate to counteract the consequences on trade of an appreciations of the Yen/dollar rate, whereas in the German-US case speculative markets might have excessively bid up the dollar rate by the demand for US bonds.

From Table 8 it also appears that there is no significant steady-state adjustment towards the real Yen/Dollar rate in the *ppp* equation. But, instead, the two inflation rates and the US bond rate and libor rate have reacted significantly on deviations from long-run *ppp*. This was not the case for Germany-US where *ppp* was significantly adjusting in the *ppp* equation as well as in US inflation rate equation.

Both the US and Japanese short interest rates show significant reaction towards each other, but also effects from the bond yield spread. The US libor rate is also affected by the inflation rate spread and the real exchange rate, but this does not seem to be the case for Japanese libor rate, again (possibly) suggesting that it has been more influenced by monetary policy interventions than by market forces.

Finally, it is worth contrasting the lack of significant inflationary effects in the interest rate equations with the significant interest rate effects in the inflation rate equations. Similar results were found in JM. It strongly suggests that the short-run adjustment effects are from bond rates to the short rates, which are then influencing inflation rates.

The effects described by the adjustment coefficients α and by the Π matrix are of short-run character. The topic of the next section is to

Table 8: The combined long-run effects in the extended Japanese model

The combined effects $\Pi = \alpha\beta'$							
Eq.	Δp_t	Δp_t^*	i_t^l	i_t^{l*}	ppp_t		
$\Delta^2 p_t$	**-1.30**	**0.54**	**0.97**	**-1.21**	**-0.57**		
	(-14.2)	(5.3)	(7.6)	(-13.2)	(13.9)		
$\Delta^2 p_t^*$	0.02	$-$**0.27**	**0.26**	-0.05	0.04		
	(0.5)	(-5.2)	(3.9)	(1.1)	(1.8)		
Δi_t^l	-0.00	0.00	-0.01	0.00	-0.00		
	(0.2)	(1.7)	(-1.2)	(0.3)	(-0.6)		
Δi_t^{l*}	**-0.02**	**0.02**	-0.01	-0.01	**-0.01**		
	(-2.6)	(3.4)	(0.6)	(-1.7)	(3.2)		
Δppp_t	-0.01	-0.01	**0.02**	-0.01	-0.00		
	(1.3)	(1.2)	(2.2)	(-1.7)	(0.7)		
	Δp_t	Δp_t^*	i_t^l	i_t^{l*}	ppp_t	i_t^s	i_t^{s*}
$\Delta^2 p_t$	**-1.34**	**0.57**	0.02	**-0.54**	**-0.24**	**0.38**	0.08
	(-13.7)	(4.6)	(0.1)	(2.9)	(9.0)	(2.7)	(0.5)
$\Delta^2 p_t^*$	-0.09	**-0.50**	**0.83**	$-$**0.66**	**0.07**	$-$**0.32**	**0.47**
	(1.7)	(-6.9)	(6.6)	(6.1)	(5.0)	(3.9)	(5.4)
Δi_t^l	-0.01	0.01	-0.00	-0.00	0.00	0.00	0.00
	(1.3)	(1.1)	(0.2)	(0.2)	(1.3)	(0.1)	(0.1)
Δi_t^{l*}	**-0.02**	**0.03**	-0.02	**-0.00**	**0.01**	0.00	-0.00
	(3.2)	(3.6)	(1.0)	(0.9)	(4.0)	(0.4)	(-0.1)
Δppp_t	-0.01	0.00	$-$**0.04**	**0.04**	0.00	**0.04**	$-$**0.04**
	(0.8)	(0.0)	(1.8)	(2.0)	(0.6)	(2.9)	(2.5)
Δi_t^s	-0.01	-0.01	**0.06**	$-$**0.05**	0.00	$-$**0.04**	**0.05**
	(1.4)	(0.8)	(3.5)	(-4.1)	(0.5)	(-3.9)	(4.1)
Δi_t^{s*}	$-$**0.03**	**0.04**	$-$**0.07**	**0.05**	**0.01**	**0.05**	$-$**0.05**
	(-2.5)	(2.7)	(2.6)	(2.1)	(3.3)	(3.0)	(2.6)

study the long-run impact of a shock to the variables on the system.

7 Weak exogeneity and the long-run impact of shocks

Based on the VAR model, the change in a variable Δx_{it} can be decomposed into its predictable part, the conditional expectation $E_{t-1}\{\Delta x_{it} \mid \Delta x_{t-1}, \beta' x_{t-1}\}$, and the unpredictable part, ε_{it}, given the information available at time t-1. The empirical investigation in this section focuses on the long-run impact of these " unanticipated shocks" on the system. Formally, the econometric analysis is based on the inverted VAR model in the so called moving average representation of the vector process:

$$x_t = C\sum_1^t \varepsilon_i + C\Phi_1 \sum_1^t D_i + C\Phi_2 \sum_1^t S_i + C^*(L)(\varepsilon_t + \mu + \Phi_1 S_t + \Phi_2 D_t) + B \tag{20}$$

where $C = \beta_\perp(\alpha'_\perp \Gamma \beta_\perp)^{-1}\alpha'_\perp$, $C^*(L)$ is an infinite polynomial in the lag operator L, and B is a function of the initial values. $\alpha_\perp$ and $\beta_\perp$ are $p \times (p-r)$ matrices orthogonal to α and β. The total impact matrix C

has reduced rank $(p-r)$ and can be decomposed (similarly as $\Pi = \alpha\beta'$) in two $p \times (p-r)$ matrices:

$$C = \widetilde{\beta}_{\perp}\alpha'_{\perp},$$

where $\widetilde{\beta}_{\perp} = \beta_{\perp}(\alpha'_{\perp}\Gamma\beta_{\perp})^{-1}$. The interpretation is that $\alpha'_{\perp}\Sigma\varepsilon_i$ is a measure of the $p-r$ common stochastic trends which influence the variables x_t with the weights $\widetilde{\beta}_{\perp}$.

It appears that the first common trend, $\alpha'_{\perp.1}\Sigma\hat{\varepsilon}_i$, is equal to the cumulated shocks to the Japanese long bond yield, consistent with the weak exogeneity result of Table 6. The second common trend is dominated by shocks to the US bond yield, whereas the remaining third and fourth trends seem primarily to be related to shocks to the short-term interest rates. $\alpha_{\perp.3}$ is essentially the sum of permanent shocks to the libor rates, describing the common long-run movements in the short-term interest rates. The last common trend captures the impact of shocks to the real exchange rate and shocks to the libor spread (shocks to the uncovered interest rate parity).

We have no standard errors for the coefficients $\widetilde{\beta}_{\perp.ij}$ and the coefficients in bold face are only indicative. The results are in accordance with our previous findings suggesting that the developments in 'world' financial markets are driven by the dominant rate yields - the US and Japanese short and long rates. The variables that adjust most strongly are prices, the *ppp* term and the short term interest rates. This latter finding reinforces the point made earlier that the Fisher conditions do not seem work in the predicted manner.

Based on (20) it is straightforward to calculate the full impulse response functions for a unitary change of $\hat{\varepsilon}_{it}$. Because we are primarily interested in the final impact of a permanent shock to the variable Δx_i, only the estimate of the final impact matrix C is reported in the lower part of Table 9. Standard errors of estimates are calculated using results in Paruolo (1997). Significant coefficients with a p-value of 0.05 or less are indicated with bold face. The entries of a column in the C matrix can be interpreted as the total (cumulated) impact of a shock, ε_{it}, on the other variables of the system, whereas a row can be interpreted as the weights with which permanent shocks to the variables of the system have influenced the long-run movements in the variable x_{it}. Hence, the significance of each entry c_{ij} gives an indication if the shock ε_i to the variable x_i has exhibited a permanent effect on the variable x_j.

We note that shocks to

(1) the two inflation rates have very small and generally insignificant effects on the other variables of the system,

Table 9: The common trends $\alpha_\perp$ and the associated weights $\beta_\perp$ and the C-matrix

	$\widetilde{\beta}_{\perp.1}$	$\widetilde{\beta}_{\perp.2}$	$\widetilde{\beta}_{\perp.3}$	$\widetilde{\beta}_{\perp.4}$	$\alpha_{\perp.1}$	$\alpha_{\perp.2}$	$\alpha_{\perp.3}$	$\alpha_{\perp.4}$
Δp_t	**0.4**	**-0.8**	0.2	-0.1	0.0	-0.0	-0.0	0.0
Δp_t^*	**0.7**	**-0.4**	0.1	0.1	0.0	0.0	0.0	-0.1
i_t^l	**1.2**	**-0.2**	0.1	-0.1	**1.0**	-0.1	-0.1	0.1
i_t^{l*}	**0.6**	**0.9**	0.0	-0.1	0.0	**1.0**	0.1	0.0
i_t^s	**0.9**	**-0.6**	**0.5**	-0.0	0.0	-0.2	**0.9**	**0.6**
i_t^{s*}	**0.3**	**0.6**	**0.5**	-0.1	0.0	0.1	**1.0**	**-0.7**
ppp_t	**-0.51**	**0.61**	**-0.04**	**0.99**	0.0	0.2	0.2	**1.0**

The estimates of the long-run impact matrix C

	$\Sigma\varepsilon_{\Delta p}$	$\Sigma\varepsilon_{\Delta p*}$	$\Sigma\varepsilon_{i^b}$	$\Sigma\varepsilon_{i^{b*}}$	$\Sigma\varepsilon_{i^s}$	$\Sigma\varepsilon_{i^{s*}}$	$\Sigma\varepsilon_{ppp}$
Δp_t	0.0 (0.6)	−0.0 (1.1)	**0.5** (4.9)	**−0.7** (5.4)	**0.3** (3.5)	**0.3** (2.6)	**−0.2** (4.2)
Δp_t^*	-0.0 (1.1)	−0.0 (0.5)	**0.8** (8.1)	**-0.3** (2.6)	**0.2** (3.1)	0.1 (0.9)	**0.1** (2.8)
i_t^l	-0.01 (1.5)	0.0 (0.9)	**1.3** (7.5)	−0.1 (0.6)	−0.0 (0.4)	0.2 (1.2)	−0.1 (0.8)
i_t^{l*}	**-0.03** (3.0)	**0.1** (2.3)	**0.5** (2.3)	**1.0** (3.6)	−0.3 (1.8)	0.2 (1.0)	0.1 (1.0)
i_t^s	**-0.02** (2.0)	0.0 (0.5)	**1.0** (5.2)	**-0.5** (2.2)	**0.5** (3.7)	**0.6** (3.1)	-0.0 (0.3)
i_t^{s*}	**-0.03** (3.3)	**0.1** (2.0)	0.1 (0.6)	**0.7** (2.3)	0.2 (1.3)	**0.7** (2.6)	0.1 (1.1)
ppp_t	0.0 (0.2)	−0.0 (0.2)	−0.5 (0.2)	0.6 (0.7)	0.5 (0.1)	**-0.8** (2.6)	**1.1** (10.0)

(2) the two long-term bond yields have significant cumulative impacts on inflation rates and also on short-term interest rate yields,

(3) the short-term interest rates have some positive long-run effects on the inflation rates, but primarily on each other and the real exchange rate, and finally

(4) the *ppp* term has significant long-run effects on itself and the two inflation rates.

A row-wise inspection shows that inflation rates are primarily influenced by shocks to the bond rates, the short-term interest rates and real exchange rates. The Japanese bond rate, being weakly exogenous, is not influenced by any of the other variables, whereas the US bond rate is to some extent influenced by permanent shocks to the Japanese variables. The libor rates are primarily influenced by permanent shocks to the bond rates and to each other, whereas the *ppp* term is significantly affected by shocks to the US libor rate.

These results reinforce the findings from our analysis of the common trends and long-run relations which were in conflict with the basic premise of the expectations hypothesis of the term structure.

8 A parsimonious short-run adjustment model

Using the identified cointegration relations reported in Table 7 we first estimated a multivariate dynamic equilibrium error correction model for the full system. Because the Japanese bond rate was found to be strongly exogenous we reestimated the system conditional on the marginal model for the Japanese bond rate. By first removing insignificant lagged variables from the system based on an F-test and then removing insignificant coefficients from the equations based on a Likelihood Ratio test we arrived at the following parsimonious model:

$$
\begin{bmatrix} \Delta^2 p_t \\ \Delta^2 p_t^* \\ \Delta i_t^{l*} \\ \Delta i_t^s \\ \Delta i_t^{s*} \\ \Delta ppp_t \end{bmatrix} = \begin{bmatrix} \underset{(2.2)}{1.88} & 0 & 0 & 0 & 0 \\ \underset{(5.5)}{2.67} & 0 & 0 & \underset{(2.0)}{0.54} & \underset{(3.8)}{0.63} \\ \underset{(4.6)}{0.31} & 0 & \underset{(4.1)}{0.21} & \underset{(3.3)}{-0.13} & 0 \\ \underset{(3.0)}{0.21} & \underset{(6.7)}{0.44} & 0 & \underset{(3.4)}{-0.14} & 0 \\ 0 & 0 & \underset{(14.0)}{1.15} & 0 & \underset{(4.7)}{0.17} \\ \underset{(5.2)}{0.44} & \underset{(2.2)}{0.14} & \underset{(2.2)}{-0.14} & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta i_t^l \\ \Delta i_{t-1}^l \\ \Delta i_{t-1}^{l*} \\ \Delta i_{t-1}^s \\ \Delta i_{t-1}^{s*} \end{bmatrix} +
$$

$$
+ \begin{bmatrix} -\underset{(19.4)}{1.20} & 0 & 0 \\ \underset{(6.7)}{-0.31} & \underset{(9.7)}{-0.63} & \underset{(4.5)}{-0.33} \\ 0 & \underset{(3.7)}{0.02} & 0 \\ 0 & 0 & \underset{(3.6)}{-0.03} \\ 0 & 0 & \underset{(3.3)}{0.04} \\ 0 & 0 & \underset{(3.8)}{0.04} \end{bmatrix} \begin{bmatrix} ecm1_{t-1} \\ ecm2_{t-1} \\ ecm3_{t-1} \end{bmatrix} + \Phi D_t + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \\ \varepsilon_{4t} \\ \varepsilon_{5t} \end{bmatrix}
$$

$$
\hat{\Sigma}(r_{ij}) = \begin{bmatrix} (0.00322) & & & & & \\ 0.08 & (0.00178) & & & & \\ -0.06 & 0.19 & (0.00023) & & & \\ -0.08 & 0.02 & -0.11 & (0.00024) & & \\ -0.12 & 0.18 & -0.13 & 0.01 & (0.00038) & \\ 0.04 & -0.03 & -0.05 & 0.02 & 0.01 & (0.00031) \end{bmatrix}
$$

where

$$
\begin{aligned}
ecm1 &= \Delta p - 0.37(\Delta p^* - i^{l*}) - 0.37 i^s + 0.0016 ppp \\
ecm2 &= \Delta p^* - 0.5\Delta p - 0.5 i^l - 0.0023 ppp \\
ecm3 &= (i^s - i^{*s}) - (i^{*l} - i^l)
\end{aligned}
$$

and the off-diagonal terms of the $\hat{\Sigma}(r_{ij})$ matrix are given as residual correlations, whereas the diagonal terms correspond to the residual standard errors. The residual cross correlations are generally very small.

The estimated coefficients of the included dummy variables are presented in Appendix A. The LR test of the overidentifying restrictions, distributed as $\chi^2(127)$, was 137.8 and the restrictions were accepted with a p-value of 0.24. Furthermore, the coefficients on the lagged bond rate in the *ppp* equation were restricted to be equal with opposite signs. It was accepted with a p-value of 0.83. Of the 127 exclusion restrictions only 24 are related to the system variables. They were tested separately and accepted with a p-value of 0.63. The remaining exclusion restrictions are mostly associated with the many intervention dummies needed to account for the large movements in US libor rate during the period of monetary targeting in the beginning of the eighties, but also with some rather big fluctuations in the Japanese libor rate when it was introduced at the end of the seventies. Furthermore, the monthly seasonal dummy variables are only included in the US and Japanese inflation rate equations.

In terms of the contemporaneous effects, we note that the strongly exogenous Japanese bond rate appears in all equations except for the US libor rate. No additional significant contemporaneous effects were discovered.

Among the lagged variables effects we note that US bond rate has adjusted positively to the lagged change in the US-Japanese libor spread and positively to $ecm2$, i.e. US bond rate has increased when US inflation is high relative to the Japanese inflation and the real exchange rate. The latter effect is likely to capture the effect of financing the trade deficit by issuing US bonds. Both libor rates react exclusively to the domestic lagged variables, and the *ppp* term reacts only to the lagged change in the bond yield spread.

In terms of equilibrium error correction, we note that Japanese inflation only reacts to the first *ecm* term, whereas US inflation adjusts to all three. The real exchange rate and the short-term interest rates are all exclusively adjusting to the third *ecm*, i.e. the relation between the long-term and the short-term spread.

The results reported in this section confirm the crucial role of the long- and short-term interest rates for the development of the real exchange rate in this period. It is interesting to note that an increase in the spread between the US and Japanese bond rates is associated with a depreciation of the dollar consistent with the predictions of the UIP, whereas the opposite is the case with an increase in the libor spread. As mentioned in the previous section this latter result could reflect the impact on the libor rate of interventions by the Japanese central bank to in trying to prevent the unpleasant consequences of Yen/$ exchange rate appreciations on Japanese trade.

9 Summary and conclusions

In this paper we have investigated the existence of interest rate and inflation rate linkages between Japan and the USA for the post Bretton Woods period of floating exchange rates. This was done by empirically examining the joint determination of a number of key parity conditions such as uncovered interest rate parity, purchasing power parity, the Fisher condition and a term structure relationship. Our starting point was a similar study based on US-German data which strongly suggested that these parity conditions have not been valid as stationary relationships on their own. Only by allowing for interactions between the standard parities stationarity was recovered. The analysis of the new Japanese-US data has further corroborated this result.

Since many of the hypotheses tested in this paper were motivated by JM we will first give a brief summary of similarities and dissimilarities and then discuss how they might have corroborated the hypothetical explanations of JM.

1. The empirical finding that the German and US bond rates are weakly exogenous for the long-run parameters was very strong in the Germany-US data. In the JP-US system this was the case for the Japanese bond rate, but not for the US bond rate. The latter was reacting to the inflation rate differential and the real Dollar/Yen rate. In both systems the results strongly suggested that it is the short-term interest rates that are adjusting to the cointegrating relations, rather than the long-term bond rates. In particular, the long-term bond rates do not seem to be affected by the short-term rates, but the latter are clearly influenced by the former. Altogether, the results provide evidence against the expectations hypothesis of the term structure of interest rates.

2. The nonstationarity of the real dollar rates was related to nonstationary movements in the long-term bond rates and to some extent in the inflation rates. The nonstationarity in the bond spread was reflected correspondingly in nonstationary movements in the spread between the short-term interest rates and again to some extent in the inflation rates. These results were remarkably similar in the two systems.

3. A very strong result both in the German-US and the Japanese-

US system was that the main adjustment towards sustainable real exchange rates took place in the inflation rates. This is an interesting finding considering that both the Dmk/$ and the Yen/$ exchange rates have been floating in the post Bretton Woods period. Hence, the substantial variability in spot exchange rates (compared to the much lower price variability) in this period cannot be interpreted as movements toward sustainable real exchange rates, but rather as the outcome of speculative behavior in the exchange market. This conform strongly with the interpretation in Krugman (1993). Significant equilibrium correction in the *ppp* was found in the German-US case, whereas no such effects were found in the Japanese-US case.

4. Another finding closely related to the above result is that nominal interest rates drive inflation rates and not the other way around. This was a very strong result that was confirmed both by the estimated short-run effects and the long-run impact effects in both systems. Inflation rates were found to have essentially no effects at all for nominal interest rates, whereas the effects from nominal interest rates to inflation rates were positive (the so called price puzzle effect). This seems to provide strong evidence against the Fisher parity condition.

5. In both cases the deviations from fundamental *ppp* was compensated by corresponding movements in both the long-term and the short-term interest rate spread, but interestingly with different signs: the Yen/dollar rate has appreciated when US-JP long-term bond spread has gone up, but depreciated with the spread in the short rates, whereas the Dmk/dollar rate has appreciated with increasing US-German treasury bill spread, but depreciated with increasing US-German bond spread. One hypothetical explanation in the Japan-US case is that it reflects the impact of the Japanese central bank interventions to prevent the unpleasant consequences of Yen/$ exchange rate appreciation on Japanese trade, and in the German-US case that speculative markets have bid up the dollar rate when buying US bonds, thereby counteracting an adjustment towards steady-state in the German-Us case.

The above findings were shown to be remarkable robust (empirically as well as econometrically) over a period of many fundamental changes. The results of the present paper emphasize the crucial role the real dollar rate has played for international monetary finance in the post Bretton Woods period. The following hypothetical explanations in JM obtained further support in the Japanese-US system:

1. The role of the dollar as a reserve currency (the 'safe haven' effect) have facilitated relatively cheap financing of the large US current account deficits in this period and counteracted an adjustment of the real dollar rate toward its stationary value (fundamental value). This might explain one of the 'market failure' puzzles: why an adequate adjustment toward purchasing power parity between the USA and the other two major currency blocks has not taken place or only very slowly so.

2. The large differences between national savings rates, and in particular the large US trade deficits, seemed to be an important reason why the long-term bond rates were found to be so crucial in this system. The globalization of the capital markets has probably strengthened this effect.

3. Though the role of central bank policy for stabilizing the short-term capital market has evidently been crucial as the turbulent years of monetary targeting in the eighties demonstrated, its role for controlling inflation seems much more modest than usually believed. This relates directly to the concern expressed by Alan Blinder (1997) and cited in the introduction about the empirical failure of the term structure and its monetary policy implications.

4. Finally, the estimated equation systems were able to explain the variation in the data with a remarkable precision. In that sense it seem quite unlikely that one would be able to estimate competing models that could beat the present results in terms of precision and stability. In that sense the results could be used as a benchmark against which theoretically more sophisticated models can be judged.

In sum, the Japanese-US linkages between interest rates, inflation rates and exchange rates reported in this paper do not conform to a standard textbook interpretation. We believe these results are challenging and merit further analysis.

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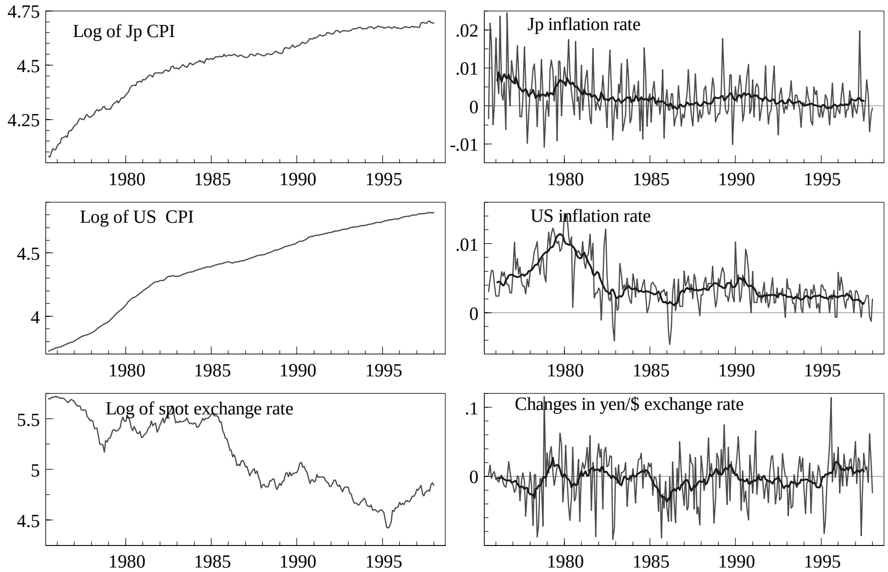
11 Appendix A: The intervention dummies

Table 10: The estimated intervention effects in the short-run adjustment model

	Δp_t	Δp_t^*	Δi_t^l	Δi_t^s	Δi_t^{s*}	Δppp_t
Di78.10	0	0	0	0	0	0.0010 (4.6)
D79.11.12	0	0	0	0.0007 (6.4)	0	0
D80.03	0	0	0	0	0	0
Di80.03	0	0	0	0	0.0017 (11.3)	0
D80.05	0	0	0	0	-0.0031 (14.9)	0
D80.07	0	-0.008 (4.8)	0	0	0.0012 (5.8)	0
D80.11	0	0	0	0	0.0012 (5.7)	0
D81.01	0	0	0	0.0008 (4.9)	-0.0012 (5.9)	-0.0008 (2.5)
D81.03	0	0	0	0.0005 (3.4)	-0.0012 (5.9)	0
Di81.05	0	0	0.0002 (2.3)	0	-0.0020 (13.4)	0
D81.10	0	-0.007 (4.0)	-0.0004 (3.0)	-0.0008 (4.3)	-0.0007 (3.4)	0
D81.11	0	0	0	0.0003 (2.2)	-0.0012 (5.7)	0
Di82.01	0	0	0	0	0	0
D82.08	0	0	0	0	-0.0021 (10.4)	0
Di84.12	0	0	-0.0002 (2.1)	0	-0.0004 (2.7)	0
D88.08	0	0	0	0.0009 (6.0)	0	0
D89.02	0	0	0	0.0007 (4.5)	0	0
D91	0.01 (10.9)	0	0	0	0	0

12 Appendix B: Graphs of the data



The graphs of Jp and US prices and spot exchange rate in levels and differences.

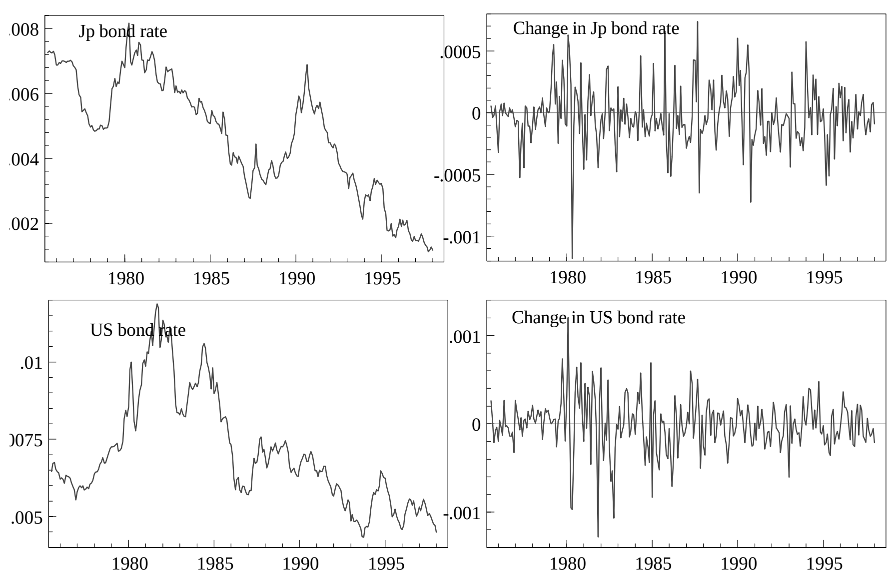


Figure B1: The graphs of Jp and US 10 year bond rates in levels and differences

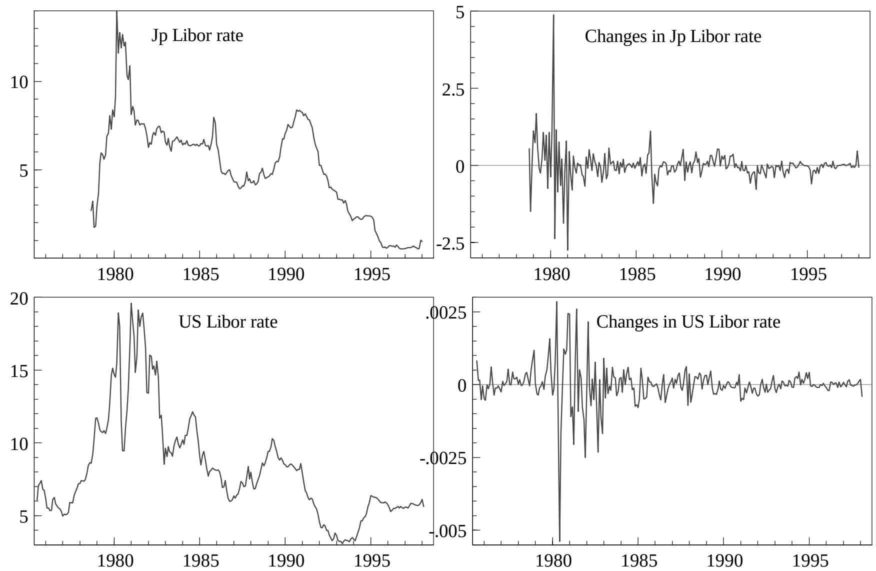


Figure B2: The graphs of the Jp and US libor rates in levels and differences.

13 Appendix C: Recursive graphs in the small model

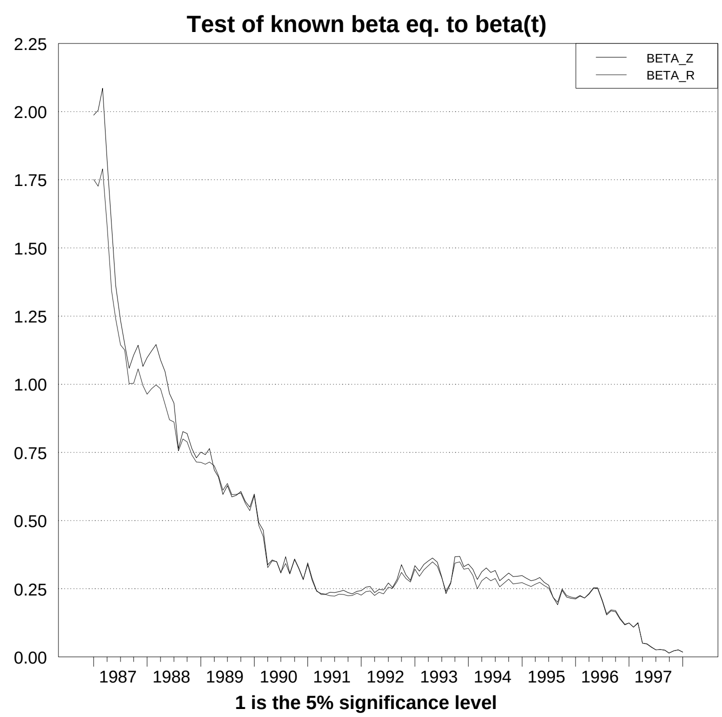


Figure C1. Recursively calculated test statistics (1.0 corresponds to the

5% significance level) for the constancy of the β vectors in the small model.

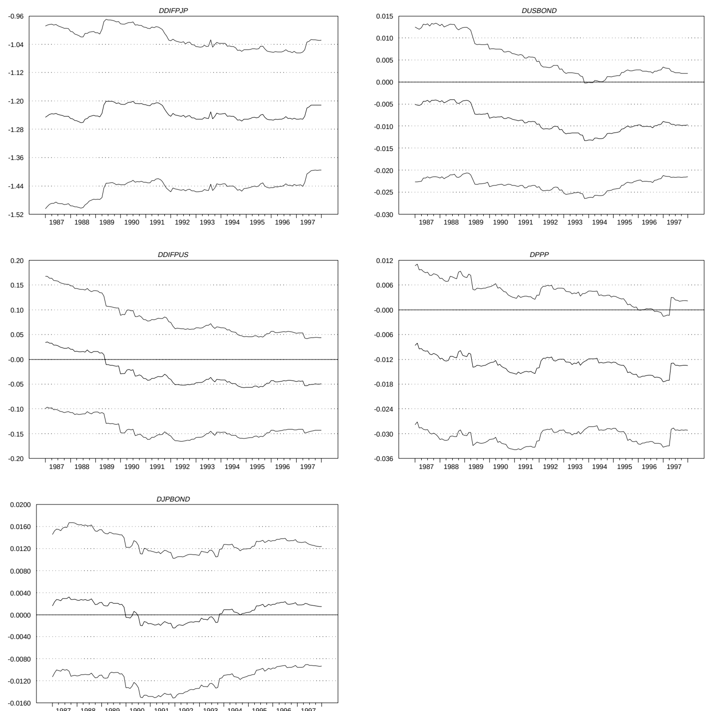


Figure C2. Recursively calculated α-coefficients to the first cointegration vector in the small model.

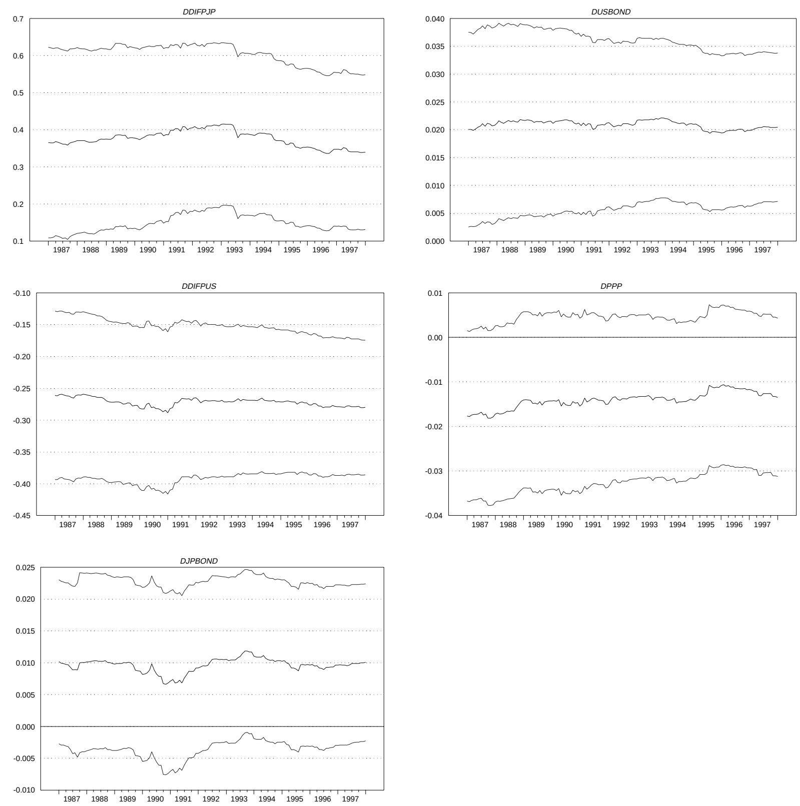


Figure C3. Recursively calculated α-coefficients to the second cointegration vector in the small model.

14 Appendix D: Recursive graphs in the big model

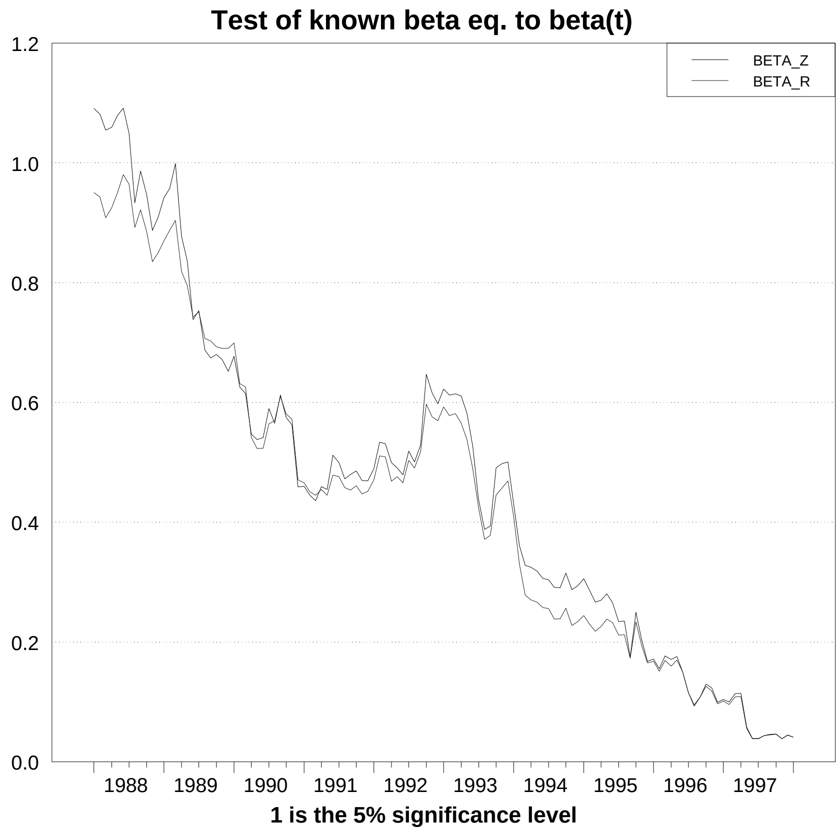


Figure D1. Recursively calculated test statistics (1.0 corresponds to the 5% significance level) for the constancy of the β vectors in the big model.

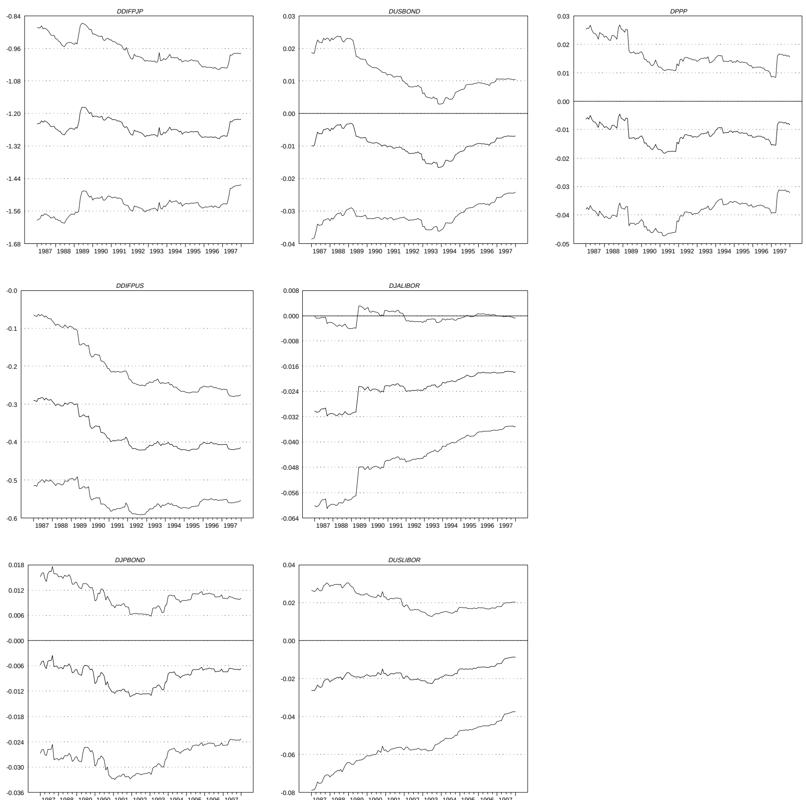


Figure D2. Recursively calculated α-coefficients to the first cointegration vector in the big model.

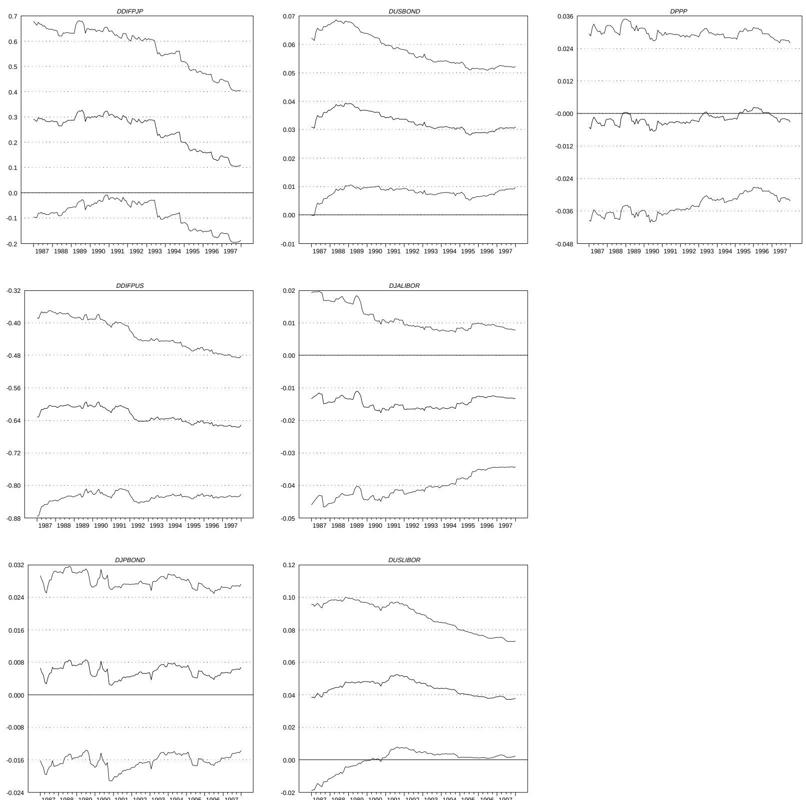


Figure D3. Recursively calculated α-coefficients to the second cointegration vector in the big model.

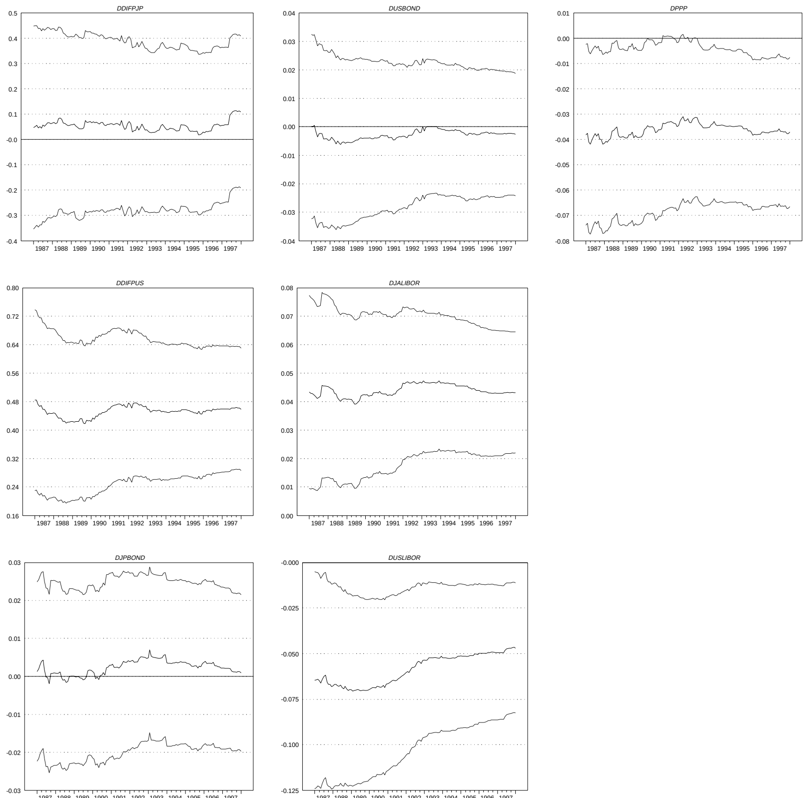


Figure D4. Recursively calculated α-coefficients to the third cointegration vector in the big model.