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Uncover Latent PPP by Dynamic Factor Error Correction Model (DF-ECM) Approach: Evidence from Five OECD Countries

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Abstract:

This study explores a new modelling approach that bridges the gap between multilateral country-level data and the bilateral-model based, goods-market specific purchasing power parity (PPP) hypothesis. Under this approach, PPP is embedded in latent common factors, extractable from a large set of bilateral price disparities, and tested via an error-correction model where the factors act as error-correction leading indicators for exchange rate and inflation. Significant modelling results for five OECD countries using monthly data suggest that the extant finding of insignificant PPP using similar data should be due to errors-in-variables attenuation and that its correction lies in effective construction of latent variables.

JEL: F31, C22, C33

Keywords: Law of one price; errors-in-variables; latent dynamic factor; error correction

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1 Introduction

It is widely observed that real exchange rates exhibit slow mean reversion and weak equilibrating power over the dynamic adjustments of nominal rates. The phenomenon forms the basis of the PPP (purchasing power parity) puzzle, i.e. empirical verification of the 'Law of One Price' (LOP) underlying PPP is much weaker than expected, cf. Obstfeld and Rogoff (2000). The puzzle has been attributed to the considerable gap between what the PPP theory assumes and the conditions of available data, especially macro data (e.g. see Taylor and Taylor, 2004; Sarno, 2005). Two issues have come to the fore – aggregation and dynamics. Concerns over aggregation are focused on the fact that heterogeneity among types of traded goods, rates of trading costs as well as heterogeneity between traded and non-traded goods across different countries is simply too pronounced to assume away empirically. A direct solution is to test the theory at a micro level, e.g. the studies carried out by Barrett (2001), Barrett and Li (2002), and Parsley and Wei (2004); a more elaborate method is to try to filter out the heterogeneous features considered to be highly significant from disaggregate panel data before inferences on PPP at a certain aggregate level are made (see e.g. Crucini et al. 2005; Imbs et al. 2005). As for dynamics, time-series studies show that different dynamic features exist not only between exchange rate and price but also among prices of different countries. Nonlinear models are used by Taylor et al. (2001) and Sarno et al. (2004) to characterise the complicated price dynamics; various VAR models and dynamic panel methods are exploited to study the exchange rate pass-through to different prices and in different countries. The literature is still growing (see e.g. Bussière 2007 for a recent survey).

The present study attempts to tackle the two issues together via a novel route. The key contention here is that it is inadequate to attack dynamics alone without considering the attenuation issue due to aggregation when country-level data are used. In fact, the source of the problem is wider than aggregation. The theoretical base of PPP is a bilateral, goods-market model, in which a domestic economy trades with a 'foreign' entity. In reality, all single countries face multilateral purchasing power disparities and interest rate disparities with numerous foreign economies each with different resource endowments, goods and capital market traditions, as well as different policies that interfere frequently with its market conditions. The gap between theory and country-level data is simply too substantive to ignore. We propose to treat the gap as an 'errors-in-variables' issue and to deal with it by taking the bilateral-model based 'foreign' variables in PPP as latent.[1] Specifically, we assume that PPP is embodied in the common factors of a dynamic factor model (DFM) comprising bilateral purchasing

[1] Conventionally, the gap is filled by construction of a real and/or a nominal effective exchange rate for the home country. However, there is no unique way of constructing such measures. Different measures contain different problems, e.g. see Ellis (2001), Chinn (2006). Moreover, different measures may lead to different inferences with respect to the verification of PPP, e.g. see Pipatchaipoom and Norrbin (2006).

power disparities of a home country with a large number of foreign economies. This amounts to filtering the heterogeneous part of the data into the idiosyncratic errors of the DFM and discarding them as measurement errors. Once the common factors are extracted, they are postulated as proxies of the disparities driving the price and exchange rate adjustment of the home country. The PPP postulate is then tested via the error-correction model (ECM): a convenient form of dynamic models as it not only facilitates the commonly adopted presentation of PPP as a long-run equilibrium condition but also verifies the condition in a much more stringent manner than what mean-reversion tests or simple cointegration analysis can achieve (see e.g. Johansen 2006).

The above procedure is referred to as the dynamic-factor error-correction model (DF-ECM) approach. The DF-ECM approach is initially explored by Qin et al. (2007) for the purpose of measuring regional market integration, and its trial application to the developing Asian region has yielded encouraging results. The present study develops the approach by applying it to the verification of PPP for five OECD countries. Thirty foreign economies are chosen to represent the world market and their price disparities vis-à-vis each of the five countries form the basis of dynamic factor analysis (DFA). Monthly data for the period of 1975-2005 are used.

The rest of the paper is organised as follows: The next section presents the DF-ECM approach; Section 3 describes practical issues pertinent to the implementation of the approach; Section 4 discusses the main findings from the five cases; The last section concludes with a short summary.

2 Method of Investigation: The DF-ECM Approach

2.1 The DF-ECM procedure

Let us start with the real exchange rate, Q, defined by PPP:

$$Q = \frac{EP}{P_f} \tag{1}$$

where P denotes the aggregate price level of the home economy of interest, P_f denotes the price level of the corresponding foreign economy and E is the exchange rate measured in the foreign currency units per unit of the domestic currency. An increasingly common way of testing PPP is to extend (1) into a dynamic model of the following log-linear form and study the mean reversion parameter, β (see, e.g. Koedijk et al. 2004):

$$q_t = \alpha + \gamma(L)\Delta q_{t-1} + \beta q_{t-1} + u_t \tag{2}$$

where $q_t = \ln(Q_t)$, $\Delta q_t = q_t - q_{t-1}$, $\gamma(L)$ is a finite-order lag polynomial and u_t is the residual term. A shortcoming of (2) is its restriction of the dynamic characteristics being

identical of, *EP* and P_f, the two price variables denominated in the same currency.[2] We relax this restriction and re-parameterise the model into an ECM:

$$\Delta(e+p)_t = \alpha(L)\Delta(e+p)_{t-1} + \gamma(L)\Delta p_{f,t} + \phi q_{t-1} + u_t \tag{3}$$

Similar to (2), the variables in small capital letter are logarithms of the corresponding variables in (1). An attractive feature of (3) is that its explanatory variables are presented by two types of structurally interpretable and empirically almost uncorrelated shocks – short-run shocks (the first two terms on the right-hand side) and a long-run disequilibrium shock (the third term), see Qin and Gilbert (2001). Notice that the long-run shock actually plays the role of a leading indicator with error-dampening capacity. Empirical verification of PPP by model (3) entails a significantly negative feedback coefficient, $\phi < 0$, signifying that the exchange rate adjusts to maintain PPP in the long run. Unfortunately, the coefficient estimates are found to be insignificant in numerous studies where country-level data are used, especially for quarterly or monthly data. When found significant, as with some cases using annual data covering very long periods, the estimates tend to be extremely small. This constitutes the so-called PPP puzzle as described at the beginning of the paper.

Here, we attribute 'errors-in-variables' attenuation as a substantive cause of the problem. As mentioned in the previous section, a considerable gap exists between the extremely abstract PPP hypothesis and the available country-level data. PPP holds only on the basis of a number of conditions – there are only two countries trading; each tradable goods follows 'law of one price'; the factor prices and production functions of the non-tradable parts of the two economies should be identical; their aggregate price indices are perfectly comparable; and, of course, the two goods markets are completely open, without capital market friction or policy interference (see e.g. Isard 1995). Judging by these conditions, errors are inevitably part of the variables Δp_f and *q*, of (3) when these are represented either by data series from one country selected as the 'numéraire' foreign counterpart or by certain weighted aggregates of a group of countries. It is well-known that attenuation becomes non-negligible when the error/noise part of the data is persistent and substantive, as it can bias the OLS estimator in a regression towards zero.

From the standpoint of an applied modeller, an effective way to correct attenuation caused by diverse measurement errors is to construct latent variables via common factor models. Here, we propose to view the foreign variables in (3) as latent and corresponding to certain common shocks of the world. These shocks are extractable by means of DFMs. Let the set of all countries be $N = \{1,2,\cdots,n\}$, the set of foreign countries vis-à-vis country *d*, the home country of interest, be $N_{-d} = \{1,\cdots,d-1,d+1,\cdots n\}$. Two DFMs are needed for measuring respectively the latent long-run shock, *q*, and the latent short-run shock, Δp_f, in (3). The first is set to extract common factors from all the observable, bilateral real rates of economy *d* vis-a-vis each of the foreign economies, i.e. $q_f = e + p - p_f$ with $f \in N_{-d}$. Defining $\boldsymbol{q_f}' = (q_1 \cdots q_f \cdots q_n)$, we assert for country *d*:

[2] Notice that the restriction is not imposed in empirical studies of the exchange rate pass-through (e.g. see Campa and Goldberg 2005).

$$\begin{aligned} \boldsymbol{q}_{f,t} &= \Psi^{*}\xi_{t}^{*} + \varepsilon_{t}^{*} \\ \xi_{t}^{*} &= \Lambda^{*}(L)\xi_{t-1}^{*} + \nu_{t}^{*} \end{aligned} \tag{4}$$

In (4), ξ^{*} is an *m*-vector of latent common factors with $m << N_{-d}$, which are thereafter referred to as the long-run factors, Ψ^{*} is a parameter matrix and $\Lambda^{*}(L)$ is a vector of lag polynomial, ε^{*} and ν^{*} are error terms with the former being idiosyncratic shocks of the foreign economies vis-a-vis country *d*. In factor analysis, $\boldsymbol{q}_{f}$ is commonly referred to as the 'indicator set' or the set of 'manifest variables'.

Similar to (4), the second DFM for extracting the latent short-run shocks writes as:

$$\begin{aligned} \Delta\boldsymbol{p}_{f,t} &= \Psi\xi_{t} + \varepsilon_{t} \\ \xi_{t} &= \Lambda(L)\xi_{t-1} + \nu_{t} \end{aligned} \tag{5}$$

where the indicator set $\Delta\boldsymbol{p}_{f}^{'} = (\Delta p_{1} \cdots \Delta p_{f} \cdots \Delta p_{n})$ is a vector of the short-run foreign inflation shocks, and ξ is an *l*-vector of latent common factors with $l << N_{-d}$, thereafter referred to as the short-run factors.

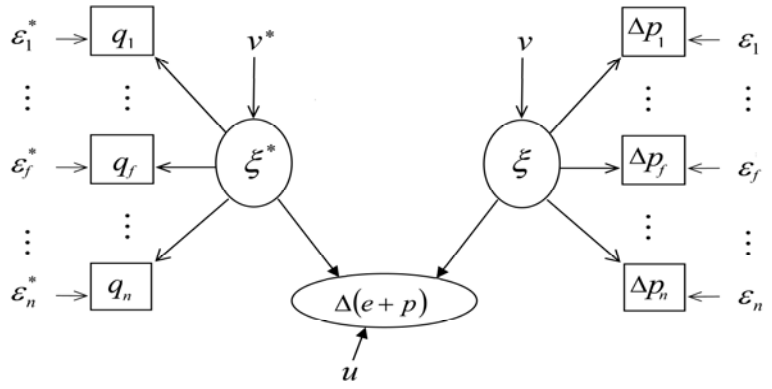
Introducing the common factors from (4) and (5) into (3) leads to a DF-ECM model:

$$\Delta(e+p)_{t} = \alpha(L)\Delta(e+p)_{t-1} + \Gamma(L)^{'}\xi_{t} + \Phi'\xi_{t-1}^{*} + u_{t} \tag{6}$$

where $\Gamma(L) = (\gamma_{1}(L) \quad \cdots \quad \gamma_{l}(L))$ is a *l*-vector of lag polynomial and $\Phi' = (\phi_{1} \quad \cdots \quad \phi_{m})$ is a *m*-vector of negative-feedback coefficients.

Notably, the present DF-ECM approach differs from most of the recent econometric studies involving DFMs, such as the ALI (automated leading indicator) approach linking DFM with VAR (vector auto-regression) by Camba-Mendez et al. (2001), and the extended structural VAR models by common factors explored by Forni et al. (2003), Bernanke et al. (2005), Favero et al. (2005) and Stock and Watson (2005). The common factors in those studies are extracted from indicators of different entities, whereas the indicators are of the same entity in the present case. The DFMs are used here primarily for filtering out measurement errors. In that sense, our approach bears close resemblance to the method of structural equation models with latent variables (SEMWLV) widely used outside econometrics, e.g. see Bedeian et al. (1997), Wansbeek and Meijer (2000), where models like (4) and (5) are referred to as the measurement equations and models such as (6) are labelled as the structural equations. However, (6) is a simpler structural equation in the sense that the modelled endogenous variable is not latent, unlike what is normally assumed in SEMWLV literature. On the other hand, both our measurement equations and our structural equation are dynamic, whereas most models in SEMWLV literature are static. Figure 1 illustrates the static version of our approach via a path diagram.

Figure 1. Static Path Diagram for Equations (4), (5) and (6)



Notice that (6) can be extended into two variants through relaxing the term $\Delta(e+p)$, which effectively allows for different dynamic pass-through of Δe_t and Δp_t. This is useful when two types of exchange rate regimes are considered. When exchange rate is fixed or under tight control, PPP works primarily via domestic price changes. Hence we have:

$$\Delta p_t = \alpha_a(L)\Delta p_{t-1} + \delta_a(L)\Delta e_t + \Gamma_a(L)'\xi_t + \Phi'_a\xi^*_{t-1} + u_{a,t} \tag{6a}$$

Whereas under the regime of a free-floating currency, the nominal exchange rate is expected to shoulder most of the adjustment with respect to PPP:

$$\Delta e_t = \alpha_b(L)\Delta p_t + \delta_b(L)\Delta e_{t-1} + \Gamma_b(L)'\xi_t + \Phi'_b\xi^*_{t-1} + u_{b,t} \tag{6b}$$

As the number of parameters in (6a) or (6b) rapidly increases when *m* and *l* are larger than two or three, the computer-automated model reduction software, PcGets, is employed for primary model simplification search, or, using the software's terminology, 'testimation'. The key advantage of PcGets is that it carries out testimation by the general → specific approach in a consistent and efficient manner. This means that the specific model thus produced is guaranteed to be data-coherent and parsimoniously encompass the general model at the starting point, see Hendry (1995), Hendry and Krolzig (2001), Owen (2003), Phillips (2005). In other words, the specific model has survived all the commonly used diagnostic tests.

2.2 Useful Statistic Indicators

A number of statistics and parameter estimates are particularly useful for informing us about the power of PPP. Some are from the ECM procedure, and others from the DFMs.

The first and foremost is the vector of the feedback coefficients, Φ, in (6). Note that the signs of these coefficients depend upon the signs of the relevant coefficients in Ψ^* of (4), e.g. ϕ_1 for the first element of ξ^* is expected to be negative if: $\sum_{i=1}^{n-1}\psi^*_{i1} > 0$,

$\Psi^* = \{\psi_{ij}^*\}_{m,n-1}$. Since there is more than one long-run factor in most cases, a simple linear combination of the surviving factors from PcGets testimation is carried to yield one EC (error correction) term (see the next section).

The next sets of statistics are summaries of the model fit from the PcGets testimation. These include, respectively, the adjusted R^2, Schwarz information criterion, the numbers of parameters of the general model at the start of testimation and of the specific model at the end. Since PcGets conducts testimation based on an array of parsimonious encompassing tests, there is no need for us to check and report these diagnostic tests here.

A popular means of verifying PPP empirically is unit-root analyses of real exchange rates. However, it has been shown that different testing methods can generate conflicting results, e.g. by Pipatchaipoom and Norrbin (2006), and that the unit-root approach may be too restrictive with respect to economic reasoning, e.g. by Coakley et al. (2005). We believe that the present ECM approach is more stringent than unit-root tests. Nevertheless, several unit-root tests are performed on the EC terms of the DF-ECMs at the final stage.

Two useful statistics are derived from the DFMs. The first is the correlation coefficient of each indicator variable, q_f, with its fitted value by the DFMs. This statistic is referred to as 'communality' in factor analysis when all the indicator variables are standardised.[3] The second statistics is the temporal correlation coefficient of all the indicator variables with their fitted values in a DFM at each sample observation, e.g. $\tau_t^2 = corr^2\left[\boldsymbol{q}_t, \left(\hat{\Psi}^* \hat{\xi}^*\right)_t\right]$ if based on (4). This statistics exploits the fact that all indicators are of the same entity. We refer to this statistics as the covariation coefficient. A time series of these coefficients is expected to show how the panel of bilateral PPPs for one economy co-moves with the set of the common factors over time. It also serves as an indication of the size of the measurement errors in the form of idiosyncratic shocks.

3 Implementation of the DF-ECM Approach

The DF-ECM approach is applied to five OECD countries: Canada, France, Germany, Japan and UK. Monthly data are collected for the period of 1975-2005.[4] These include consumer price indices (CPI) and US dollar denominated exchange rates. The Appendix Table A1 gives the details of all the series and their sources.

[3] Tucker and MacCallum (1997) give detailed discussion about the statistics. As the number of long-run common factors may vary across different countries, adjusted R^2 is used here instead of the simple R^2.

[4] In the earlier drafts of the paper, quarterly models were also presented as it was uncertain before any experiments whether monthly models would generate any significant results. The quarterly model results are now omitted to make the paper shorter.

3.1 Implementation of DFMs

Choice of the indicator set: In addition to the above five countries, twenty six economies are selected by the criterion that the selected country set covers over 70%~80% of the total trade for each of the five countries.[5] This makes N_{-d} contain 30 economies and N=31 for each of the five. All the indicator series are adjusted to zero-mean series. The long-run indicator sets are also standardised, but the short-run indicator sets are not as the short-run indicators are already US$ comparable foreign inflation rates.

Determination of the number of factors: Two recently developed procedures of consistent estimators are utilized. One is developed by Bai and Ng (2007) and the other by Onatski (2005). The larger of the two estimates is adopted when they differ. Table 1 reports the estimated results of the two procedures.

Factor extraction: DFMs (4) and (5) are estimated using the technique developed by Camba-Mendez et al. (2001). Kalman filter algorithm is used with the initial parameter estimates obtained via principal component analysis. One advantage of this is that the algorithm can handle an unbalanced data panel like ours, where the CPI data series start later than 1975M01 for countries like China and Czech Republic, and quarterly CPI like that of Australia (see the Appendix Table A1). As for the short-run indicator set, there are only 29 indicators when Australia drops out.

Determination of the number of lags: The experiment starts from L=1 and moves on to L=2 and L=3. A lag number is then chosen with reference to information criteria, such as Akaike and Schwarz criteria. It is found through numerous DFM experiments that one lag is adequate for the extraction of short-run factors by (5) whereas two or three lags are necessary for long-run factors by (4). The results are given in Table 1.

Table 1. Specification of the DFMs (4) and (5)

	Number of factors (Onatski procedure / Bai-Ng procedure)			Lag length for DFM (4)
	Long run	Short run (quarterly)	Short run (monthly)	
Canada	5 / 3	5 / 1	5 / 1	2
France	6 / 6	5 / 1	5 / 1	3
Germany	6 / 3	5 / 1	5 / 1	2
Japan	6 / 3	5 / 1	5 / 1	3
UK	6 / 5	5 / 1	5 / 1	2

Note: The larger number is adopted for the number of factors when the estimates of the two procedures differ. The lag length for DFM (5) remains one.

3.2 Implementation of DF-ECMs

Models (6a) and (6b) are the focal point of experiments, though (6) is tried first for each country (to keep this paper brief, the results are not reported here). OLS is used for model estimation. Notice, however, that the estimation method is comparable to a 2SLS

[5] The trade data are checked from the Trade Profile Statistics by the World Trade Organisation.

(two-stage least squares) procedure, where common factors extracted from (4) and (5) are used effectively as IVs (instrumental variables) to circumvent the errors-in-variables problem. As mentioned before, model simplification is a primary task here. We start by trying various lag lengths and found that six lags are generally adequate. The default setting of liberal model selection by Hendry and Krolzig (2001) is used for model testimation. Since coefficient constancy is a major concern, model testimation is performed for different sample periods, starting from the full sample, then for sub-samples of 1980-2005 and 1985-2005 respectively. The resulting specific models are further simplified, mainly through reparameterisation and linear combination of the long-run factors, using PcGive (for details on reparameterisation, see e.g. Hendry 1995). Recursive estimation is used here to monitor coefficient constancy. Hansen parameter instability test (1992) is also calculated.

4 Application Results

Data series of the both modelled variables, Δe_t and Δp_t, for each of the five countries are plotted in Appendix Figures A1-A5. In order to compare the DF-ECM results with conventional results, standard ECMs for the five countries are run using real effective exchange rates (REERs) as the EC terms (see Appendix Table A1 for data information of these REERs). REERs are chosen here for the main reason that it is more comparable to the multilateral setting of the DFM-based real rate measures than bilateral real rate measures. Besides, REERs have been used in empirical tests of PPP by numerous researchers (see e.g. Corbae and Ouliaris 1991, Bahmani-Oskooee 1995, Ellis 2001, Chinn 2006).

4.1 General results

The most noticeable result from Appendix Tables A2-A6 is that the DFM-based real rates, i.e. the long-run EC terms, are all significant and that their feedback coefficients display a high degree of constancy, as shown by the Hansen test statistics given under the coefficient estimates. The constancy can also be seen from the recursive estimation graphs plotted in the bottom panels of Appendix Figures A1-A5. In contrast, the long-run EC terms in the form of $\ln(REER)$ of those standard ECMs are all insignificant except for model (6a) in Japan and (6b) in UK, where, however, the Hansen statistics reveal significant coefficient instability.[6] The insignificance of $\ln(REER)$ is consistent with the extant finding in the literature. The cause is often attributed to the nonstationary feature of *REER*. This is reconfirmed by the unit-root tests on the $\ln(REER)$ series shown in Table 2. In the table, unit-root tests of the DF-based EC terms are also presented. It is easy to see that the nonstationary feature is more pronounced in $\ln(REER)$ than in the DF-based EC terms, though the test results on these latter terms are quite mixed, reinforcing the findings by Pipatchaipoom and Norrbin (2006).

[6] In fact, the ECMs using $\ln(REER)$ often suffer from unsatisfactory diagnostic tests, though these are not reported here to keep the paper short.

Here, we attribute the insignificance to another cause – measurement error attenuation. As seen from the graphs of $\{\tau_t^2\}$ based on DFM (4) in Appendix Figures A1-A5, these covariation coefficient series remain small (mostly well below 30%) and erratic, suggesting that idiosyncratic errors form a major part of the data at each observation point. In other words, substantive measurement errors are present in $\boldsymbol{q}_f$ if the set is used directly to construct the theoretical entity of real exchange rates, such as *REER*. Notably, the measurement error problem may not be unrelated to the nonstationary problem, since one source of nonstationarity is accumulation of independent errors. Inspecting the rescaled plots of the $\ln(REER)$ series together with those $\hat{\xi}^*$ series in Appendix Figures A1-A5, we can see that the $\ln(REER)$ series tend to exhibit longer periods of random drifts than the $\hat{\xi}^*$ series in general. The unscaled plots show that the $\ln(REER)$ series are far less volatile than the $\hat{\xi}^*$ series. That explains why the coefficients of the DFM-based EC terms are substantially smaller in magnitude than those of $\ln(REER)$. On the other hand, we see from these graphs that the $\hat{\xi}^*$ series are different between models (6a) and (6b) of the same country. This

Table 2. Unit-root Test Statistics On Selected EC Terms

Country	Tests	$\hat{\xi}_t^*$ for (6a)	$\hat{\xi}_t^*$ for (6b)	$\ln(REER)$
Canada	ADF	-1.3151 (2)	-3.1624*** (2)	-1.4901 (0)
	Phillip-Perron	-1.3392 [6]	-2.5309** [4]	-1.5705 [1]
	DF-GLS	-0.1323 (2)	-0.883 (2)	-1.4478 (0)
	Ng-Perron (MZ_t)	-0.1446 (2)	-0.8229 (2)	-1.4405 (0)
France	ADF	-2.5951*** (1)	-3.3505*** (4)	-2.5925* (1)
	Phillip-Perron	-2.5452** [7]	-3.4732*** [18]	-2.4999 [1]
	DF-GLS	0.4408 (1)	0.0816 (4)	-0.7230 (1)
	Ng-Perron (MZ_t)	0.468 (1)	0.0742 (4)	-0.7272 (1)
Germany	ADF	-1.8253* (1)	-1.4513 (1)	-2.0515 (0)
	Phillip-Perron	-2.0562** [8]	-1.6748* [9]	-2.394 [5]
	DF-GLS	-1.7822 (1)	-1.1219 (1)	-1.8122 (0)
	Ng-Perron (MZ_t)	-1.7844* (1)	-1.211 (1)	-1.7963* (0)
Japan	ADF	-2.2115** (0)	-2.5068** (0)	-2.3792 (1)
	Phillip-Perron	-2.2115** [2]	-2.86*** [8]	-1.9717 [1]
	DF-GLS	-0.2969 (0)	-1.0756 (0)	-1.1135 (1)
	Ng-Perron (MZ_t)	-0.2860 (0)	-1.0683 (0)	-1.0986 (1)
UK	ADF	-2.1424** (1)	-2.3913** (1)	-1.9726 (1)
	Phillip-Perron	-1.9461** [9]	-2.3059** [3]	-1.7762 [3]
	DF-GLS	1.1856* (0)	-2.3262** (1)	-1.9292* (1)
	Ng-Perron (MZ_t)	1.2203 (0)	-2.3006** (1)	-1.9279* (1)

Note: The sample periods used correspond to those used in the model estimation and reduction (see Appendix Tables A2-A6). ADF denotes augmented Dickey-Fuller test; DF-GLS is Elliott-Rothenberg-Stock test (1996); Only MZ_t out of the four tests in (Ng-Perron, 2001) is reported to save space. *, ** and *** indicate rejection of the unit-root null hypothesis at 10%, 5% and 1% respectively. The numbers in parentheses are the number of lags used in the tests and these numbers are chosen on the basis of information criteria. The number in the square brackets of Phillip-Perron test (1988) is bandwidth determined by means of Bartlett kernel.

finding further supports the view that the PPP hypothesis is deeply latent in aggregate data which are full of noises interwoven in complicated dynamics.

As for the expected signs of the coefficients of the significant long-run factors, these can be checked against Table 3, where $\sum_{i=1}^{n}\psi_{ij}^{*}$ (j=m) and the associate standard errors from DFM (4) are reported. Since all the standard errors are fairly large, the implied 95% confidence intervals are generally too wide to restrict any of the feedback coefficients in (6a) or (6b) within the strictly negative range. In terms of the adjustment speed, it is interesting to note that the feedback coefficient estimates of the exchange rate models (6b) are larger in absolute value than those of the inflation models (6a). This evidence is in support of the common view that goods prices are far less responsive than nominal exchange rates to external shocks under the freely floating regime.

Notice that the short-run common factors play an important role in the DF-ECMs as well. This is particularly striking when the R^2 statistics between the DF-ECMs and the corresponding *REER*-based ECMs are compared (see Appendix Tables A2-A6). On the whole, exchange rates are more responsive than inflation to the short-run factors and react to them in a more instantaneous manner. This feature renders support to the relative version of PPP.

As five short-run factors and five to six long-run factors are found necessary for each country, automated model testimation by PcGets becomes essential, as shown in Table 4. In fact, a great deal more of testimation experiments have been carried out than what is reported here. One particular feature easily revealed during PcGets testimation is that the DF-ECMs do not fit well with subsamples including the prior-1980 data for some countries, e.g. Japan. On the whole, the DF-ECMs fit better with post 1980 sub-samples than the full sample. If the adjusted R^2 statistics in Table 4 are compared with those of the DF-ECMs in Appendix Tables A2-A6, one can easily see that further model reduction through reparameterisation helps to improve model fit moderately.

Finally, let us look at the correlation coefficients between the indicator sets and their explained parts by DFMs (4) and (5) respectively. The coefficients are given in Table 5 and Table 6 and ranked by size. Two features are worth commenting on. First, the

Table 3. Coefficient Estimates Of The Long-run Factors Based On DFM (4)

Country	Long-run factors	ξ_1^*	ξ_2^*	ξ_3^*	ξ_4^*	ξ_5^*	ξ_6^*
Canada	$\sum_{i=1}^{30}\psi_{ij}^{*}$	0.6081	0.9977	0.4802	0.6178	0.6992	N/A
	Standard error	(1.1320)	(1.7913)	(1.6057)	(1.6186)	(1.3635)	
France	$\sum_{i=1}^{30}\psi_{ij}^{*}$	1.7250	-0.6758	0.2060	1.4969	0.1618	0.0617
	Standard error	(3.1159)	(2.9753)	(2.4701)	(2.7346)	(5.2093)	(3.5536)
Germany	$\sum_{i=1}^{30}\psi_{ij}^{*}$	0.6645	-0.8278	0.4149	0.3089	0.6127	0.1842
	Standard error	(8.8665)	(3.2152)	(4.9475)	(6.5889)	(5.4767)	(1.9901)
Japan	$\sum_{i=1}^{30}\psi_{ij}^{*}$	2.8930	-0.5342	0.1620	1.9058	-0.2133	-0.2055
	Standard error	(3.3003)	(5.7676)	(6.2744)	(5.9941)	(6.0424)	(4.9895)
UK	$\sum_{i=1}^{30}\psi_{ij}^{*}$	1.5919	-0.0701	1.1703	-0.0926	0.9258	0.0793
	Standard error	(4.7444)	(14.6691)	(4.3594)	(3.5113)	(2.6611)	(4.3556)

correlation coefficients in Table 5 are substantially larger than those in Table 6, indicating that slow mean reversion must prevail among the bilateral real rate series of the indicator set of DFM (4). Secondly, the correlation rankings across countries are far more similar in Table 6 than in Table 5. This is because the short-run indicator sets differ from each other only by one indicator, namely that of the home country under study. Notice also that France, Germany and Japan rank fairly high in the coefficient sequences of Table 6. This helps to explain why short-run common factors play such a significant role in the DF-ECMs of these three countries.

Table 4. Summary Statistics Of Model-Fit Via PcGets Testimation of (6a) And (6b)

		Sample	General model		Specific model		Number of parameters
	Equation	starting	Adjusted	Schwarz	Adjusted	Schwarz	from general →
Country		point	R^2	criterion	R^2	criterion	specific
	Δe_t	1975M08	0.0692	-7.7213	0.076	-8.3877	54 → 4
Canada		1980M01	0.0701	-7.6551	0.1104	-8.3971	54 → 7
	Δp_t	1975M08	0.3019	-10.668	0.3027	-11.315	54 → 5
		1980M01	0.2663	-10.591	0.2703	-11.325	54 → 5
	Δe_t	1975M10	0.9413	-9.0411	0.9414	-9.5694	55 → 15
France		1980M01	0.9665	-9.4696	0.9648	-9.9344	55 → 20
	Δp_t	1975M10	0.6818	-11.493	0.6765	-12.071	55 → 10
		1980M01	0.7147	-11.673	0.7011	-12.188	55 → 17
	Δe_t	1975M08	0.978	-9.9772	0.9789	-10.478	55 → 20
Germany		1980M01	0.9831	-10.127	0.9827	-10.663	55 → 17
	Δp_t	1975M08	0.2045	-11.024	0.2109	-11.597	55 → 12
		1980M01	0.2522	-10.974	0.2429	-11.598	55 → 12
	Δe_t	1975M08	0.3116	-6.4049	0.3134	-7.1198	55 → 1
Japan		1985M01	0.3704	-6.2411	0.3633	-7.0829	55 → 6
	Δp_t	1975M08	0.3198	-10.307	0.3109	-10.901	55 → 14
		1980M01	0.3437	-10.502	0.3228	-11.216	55 → 12
	Δe_t	1975M08	0.5651	-7.0798	0.5755	-7.7225	55 → 8
UK		1980M01	0.5666	-6.9828	0.5636	-7.7184	55 → 5
	Δp_t	1975M08	0.3824	-10.013	0.3547	-10.547	55 → 11
		1980M01	0.3895	-10.21	0.3629	-10.729	55 → 17

Note: six lags are used in the general models. All samples end at 2005M12.

Table 5. Ranked Correlation Coefficients Between the Indicators in q_t and the Fitted $\left(\hat{\Psi}^* \hat{\xi}_t^*\right)$ of DFM (4)

	Canada	France	Germany	Japan	UK
1	0.973 USA	0.967 Austria	0.971 Malaysia	0.977 Malaysia	0.970 Belgium
2	0.963 Malaysia	0.965 Malaysia	0.970 Austria	0.976 India	0.962 Germany
3	0.958 Denmark	0.958 Saudi Arabia	0.969 Czech Rep.	0.975 Belgium	0.962 Malaysia
4	0.952 Austria	0.957 Czech Rep.	0.968 Saudi Arabia	0.973 Netherlands	0.961 Netherlands
5	0.948 Belgium	0.955 USA	0.954 USA	0.970 Germany	0.955 Austria
6	0.943 Netherlands	0.948 India	0.953 India	0.969 France	0.953 India
7	0.942 France	0.941 Singapore	0.946 Hong Kong	0.969 Czech Rep.	0.950 Denmark
8	0.941 Germany	0.933 China	0.941 Singapore	0.965 USA	0.950 France
9	0.939 Thailand	0.923 Denmark	0.935 China	0.965 Thailand	0.941 China
10	0.937 Singapore	0.921 Taiwan	0.931 Thailand	0.961 Denmark	0.941 Thailand
11	0.932 Poland	0.921 Ireland	0.926 Ireland	0.954 Austria	0.940 Saudi Arabia
12	0.928 Switzerland	0.920 Thailand	0.925 Netherlands	0.952 Sweden	0.938 USA
13	0.912 India	0.916 Belgium	0.920 Italy	0.951 Taiwan	0.936 Singapore
14	0.912 Taiwan	0.910 Netherlands	0.919 Taiwan	0.951 Norway	0.935 Taiwan
15	0.910 Italy	0.910 Poland	0.882 Poland	0.948 Italy	0.921 Sweden
16	0.907 Spain	0.908 Italy	0.880 Sweden	0.948 Ireland	0.907 Norway
17	0.884 China	0.886 Germany	0.867 Spain	0.938 Canada	0.903 Canada
18	0.884 Ireland	0.883 Hong Kong	0.847 Denmark	0.938 Spain	0.893 Italy
19	0.875 Norway	0.870 Canada	0.844 Canada	0.930 China	0.887 Czech Rep.
20	0.875 Japan	0.861 Switzerland	0.833 UK	0.917 Australia	0.885 Hong Kong
21	0.873 Saudi Arabia	0.858 Spain	0.823 Japan	0.914 Saudi Arabia	0.879 Spain
22	0.863 Czech Rep.	0.839 Japan	0.822 Norway	0.913 Hong Kong	0.877 Ireland
23	0.858 Hong Kong	0.830 UK	0.821 Belgium	0.897 Switzerland	0.877 Switzerland
24	0.832 Turkey	0.820 Sweden	0.811 Switzerland	0.895 Singapore	0.830 Poland
25	0.822 Sweden	0.801 Turkey	0.797 Turkey	0.892 Turkey	0.829 Mexico
26	0.771 UK	0.759 Norway	0.764 France	0.873 Poland	0.821 Australia
27	0.747 South Korea	0.673 Australia	0.736 Australia	0.872 South Korea	0.820 South Korea
28	0.697 Mexico	0.672 South Korea	0.735 South Korea	0.842 Mexico	0.805 Turkey
29	0.497 Australia	0.601 Mexico	0.735 Mexico	0.834 UK	0.785 Japan
30	0.063 Brazil	0.074 Brazil	0.069 Brazil	0.048 Brazil	0.066 Brazil

Note: The adjusted R^2 is used, instead of the simple R^2 in order to make comparable the cases with different numbers of factors.

Table 6. Ranked Correlation Coefficients Between the Indicators in Δp_f And the Fitted $\left(\hat{\psi}\hat{\xi}_t\right)$ of DFM (5) Using Three-Month Rates

	Canada	France	Germany	Japan	UK
1	0.538 Malaysia	0.552 Malaysia	0.537 Malaysia	0.535 Malaysia	0.519 South Korea
2	0.469 France	0.469 Denmark	0.469 Denmark	0.469 Denmark	0.467 Denmark
3	0.457 Norway	0.457 Norway	0.455 Belgium	0.457 Norway	0.453 Netherlands
4	0.456 Belgium	0.454 Belgium	0.454 Norway	0.455 Belgium	0.453 Belgium
5	0.452 Hong Kong	0.450 Hong Kong	0.444 Austria	0.452 Germany	0.450 Germany
6	0.449 Austria	0.446 Austria	0.431 France	0.448 Austria	0.445 Austria
7	0.441 Japan	0.410 Germany	0.429 Hong Kong	0.431 France	0.428 France
8	0.436 Germany	0.403 Japan	0.412 Japan	0.412 Italy	0.404 USA
9	0.390 Italy	0.388 Italy	0.395 Italy	0.408 South Korea	0.399 Italy
10	0.379 Singapore	0.382 Singapore	0.384 Singapore	0.401 Ireland	0.381 Hong Kong
11	0.374 Netherlands	0.371 India	0.373 Taiwan	0.383 Singapore	0.377 Norway
12	0.373 Taiwan	0.369 Taiwan	0.367 Switzerland	0.371 Taiwan	0.368 Ireland
13	0.360 Switzerland	0.367 Switzerland	0.366 India	0.366 Switzerland	0.365 Switzerland
14	0.359 Sweden	0.340 Sweden	0.346 Sweden	0.362 Hong Kong	0.357 Sweden
15	0.356 Poland	0.323 USA	0.327 Turkey	0.348 Turkey	0.342 Spain
16	0.356 Denmark	0.323 Turkey	0.320 USA	0.344 Sweden	0.328 Thailand
17	0.343 Mexico	0.318 Mexico	0.316 Poland	0.328 Mexico	0.326 Malaysia
18	0.324 Ireland	0.315 Poland	0.313 Mexico	0.326 USA	0.314 Poland
19	0.324 Turkey	0.302 Ireland	0.296 Thailand	0.319 India	0.292 Singapore
20	0.313 USA	0.290 Spain	0.286 Spain	0.309 Poland	0.292 Taiwan
21	0.310 India	0.288 Thailand	0.284 Ireland	0.287 Spain	0.285 Canada
22	0.291 Thailand	0.228 Canada	0.230 Canada	0.286 Thailand	0.249 Turkey
23	0.235 Spain	0.188 South Korea	0.193 South Korea	0.238 Canada	0.194 Japan
24	0.196 South Korea	0.175 Netherlands	0.172 Netherlands	0.170 Netherlands	0.167 Mexico
25	0.141 Czech Rep.	0.145 Brazil	0.141 Brazil	0.153 Brazil	0.144 India
26	0.129 Saudi Arabia	0.107 Czech Rep.	0.102 Czech Rep.	0.112 Czech Rep.	0.143 Brazil
27	0.089 Brazil	0.085 Saudi Arabia	0.083 Saudi Arabia	0.082 Saudi Arabia	0.112 Australia
28	0.078 Australia	0.070 Australia	0.079 UK	0.064 Australia	0.087 Czech Rep.
29	0.062 UK	0.068 UK	0.078 Australia	0.057 UK	0.064 Saudi Arabia
30	0.038 China	0.035 China	0.044 China	0.021 China	0.034 China

Note: Three-month rates are used here to extract the short-run common factors by DFM (5) because the Australia CPI series is in quarterly only. However, the short-run common factors used in the DF-ECM models are obtained from monthly rates. The Adjusted R^2 is used, instead of the simple R^2 in order to make comparable the cases with different numbers of factors.

4.2 Individual countries

Canada: The DF-ECMs show reasonable fit (see Appendix Table A2) with fairly constant long-run coefficients (see Appendix Figure A1). The long-run coefficients in (6a) are clearly consistent with the positive coefficient estimates of ξ_2^*, ξ_4^* and ξ_5^* from DFM (4) in Table 3. As for ξ_1^*, the large standard error of 1.132 (Table 3) makes its 95% confidence interval cover as low as -1.65, well allowing for the positive feedback coefficient of +0.0002 in (6a) of Appendix Table A2. The feedback coefficient estimate of (6b) is about three times of that of (6a), indicating a much stronger PPP response in the exchange rate dynamics than the inflation dynamics.
France: Model (6b) fits remarkably well in sharp contrast to the poor fit of the *REER*-based ECMs (see Appendix Table A3). The two $\hat{\xi}^*$ series for (6a) and (6b) are almost identical. The signs of the feedback coefficients are consistent with those of the factor loading coefficients from (4) implied in Table 3.
Germany: Only ξ_3^* survives the testimation in (6b), though the model fits remarkably well, even better than (6a), mainly due to the explanatory power of the short-run common factors (see Appendix Table A4). The relatively weak EC term here is also reflected in the unit-root test results in Table 2.
Japan: PcGets testimation reveals that sensible DF-ECMs only become possible for the post-1980 periods. In fact, only in the current-period does the first short-run factor survive in the full-sample testimation of model (6b) (see Table 4). This is also discernible from the recursive estimation graphs in Appendix Figure A4, where convergence to constancy of the feedback coefficients occurs around the end of the 1980s. The $\ln(REER)$ term is significant in model (6a) but its coefficient fails the constancy test (see Appendix Table A5).
UK: Noticeably from Appendix Figure A5, the dynamic pattern of $\ln(REER)$ resembles that of the $\hat{\xi}^*$ series of model (6b), except for the post-2000 period. This may help to explain why the $\ln(REER)$-based EC term is significant in the comparable model. But the coefficients suffer from non-constancy (see Appendix Table A6).

5 Concluding Comments

This study explores a new modelling approach to empirically verify PPP. Under the new approach, PPP is embodied in latent common factors, extractable from a large set of bilateral price disparities, and tested via an error-correction model where the factors act as error-correction leading indicators for exchange rate and inflation. The indicators are found significant in monthly inflation and exchange rate models for five OECD countries. The finding reverses the commonly held belief, based on numerous previous results, that PPP is at best a very long-run relationship at the macro level, verifiable only with low-frequency data over very long sample periods.

A key reason for the present PPP evidence is that the new approach provides us with an effective means of correcting attenuation caused by the errors-in-variables problem. The source of the problem is the immense gap between multilateral country-level data and the bilateral-model based, goods-market specific purchasing power parity (PPP) hypothesis. So far, the problem has only been tackled via the use of micro market data. In the present study, country-level data are used and the errors are identified mainly as

the idiosyncratic shocks in DFMs and filtered out before the dynamic model containing PPP in the form of ECM is estimated. The PPP-based price disparities are treated as latent theoretical constructs.

Another advantage of the new approach is the combination of dynamic factors and the ECM approach. Conceptually, the long-run common factors match with the leading indicator interpretation of the EC term in an ECM, and the ECM lends its structural interpretation conveniently to both the long-run and the short-run factors. Empirically, the ECM and the associate general-to-specific modelling strategy renders more robust results than those by various means of nonstationarity tests.

Appendix Table 1: Variable and Data Sources

Economy	Variable and source	Particulars
Australia	CPI and US$ exchange rate from Datastream; CPI is from Australian Bureau of Statistics	CPI is quarterly
Austria	CPI = OEI64 of IFS; US$ exchange rate from Datastream	
Belgium	CPI = BGI64 of IFS; US$ exchange rate from Datastream	
Brazil	CPI = BRI64 of IFS; US$ exchange rate from Datastream	CPI sample starts from: 1980M02
Canada	CPI = CNI64 of IFS; US$ exchange rate from Datastream REER from Datastream (OECD source)	
China	CPI = CHI64 of IFS; US$ exchange rate from Datastream; For data prior to 1993 are from State Bureau of Statistics of China	CPI sample starts from: 1982M01
Czech Republic	CPI = CZI64 of IFS; US$ exchange rate from Datastream	CPI sample starts from: 1991M01; exchange rate starts from: 1993M01
Denmark	CPI = DKI64 of IFS; US$ exchange rate from Datastream	
France	CPI = FRI64 of IFS; US$ exchange rate from Datastream REER from Datastream (OECD source)	REER sample starts from: 1980M01
Germany	CPI = BDI64 of IFS; US$ exchange rate from Datastream REER from Datastream (OECD source)	
Hong Kong	CPI = HKI64 of IFS; US$ exchange rate from Datastream	
India	CPI = INI64 of IFS; US$ exchange rate from Datastream	
Ireland	CPI = IRI64 of IFS; US$ exchange rate from Datastream	
Italy	CPI = ITI64 of IFS; US$ exchange rate from Datastream	
Japan	CPI = JPI64 of IFS; US$ exchange rate from Datastream REER from Datastream (OECD source)	
Korea, South	CPI = KOI64 of IFS; US$ exchange rate from Datastream	
Malaysia	CPI = MYI64 of IFS; US$ exchange rate from Datastream	
Mexico	CPI = MXI64 of IFS; US$ exchange rate from Datastream	
Netherlands	CPI = NLI64 of IFS; US$ exchange rate from Datastream	
Norway	CPI = NWI64 of IFS; US$ exchange rate from Datastream	

Poland	CPI = POI64 of IFS; US$ exchange rate from Datastream	Sample for both series: 1988M1 — 2005M12
Saudi Arabia	CPI = SII64 of IFS; US$ exchange rate from Datastream	CPI sample: 1980M2 — 2005M12
Singapore	CPI = SPI64 of IFS; US$ exchange rate from Datastream	
Spain	CPI = ESI64 of IFS; US$ exchange rate from Datastream	
Sweden	CPI = SDI64 of IFS; US$ exchange rate from Datastream	
Switzerland	CPI = SWI64 of IFS; US$ exchange rate from Datastream	
Taiwan	CPI and US$ exchange rate from Datastream; CPI is from Directorate General of Budgets, Accounting and Statistics, Executive Yuan of Taiwan	
Thailand	CPI = THI64 of IFS; US$ exchange rate from Datastream	
Turkey	CPI = TKI64 of IFS; US$ exchange rate from Datastream	
UK	CPI = UKI64 of IFS; US$ exchange rate from Datastream REER from Datastream (OECD source)	
USA	CPI = USI64 of IFS	

Note: All the series are monthly for the period of 1975M1 — 2005M12 except for those noted in the particulars. IFS denotes *International Financial Statistics* by IMF.

Appendix Table A2. Specific Models of (6a) and (6b) Versus ECMs of *REER*: Canada

(6a)

$$\Delta\hat{p}_t = \underset{(0.1398)}{\underset{(0.00025)}{0.003}} + \underset{(0.3827)}{\underset{(0.0506)}{0.1695}}\,\Delta p_{t-4} - \underset{(0.3045)}{\underset{(0.00006)}{0.0002}}\,\xi_{1,t-2} + \underset{(0.0963)}{\underset{(0.00002)}{0.0002}}\,\hat{\xi}^{*}_{t-1}$$

$$\hat{\xi}^{*}_{t-1} = \left(\xi^{*}_{1} - 4\xi^{*}_{2} - 0.5\xi^{*}_{4} - 2\xi^{*}_{5}\right)_{t-1}; \qquad R^2 = 0.3305 \quad \overline{R}^2 = 0.3249$$

Using *REER*:

$$\Delta\hat{p}_t = \underset{(1.0824)^{**}}{\underset{(0.0078)}{0.0035}} + \underset{(1.3596)}{\underset{(0.0233)}{0.1555}}\,\Delta_3 p_{t-1} + \underset{(1.5807)^{**}}{\underset{(0.0527)}{0.2623}}\,\Delta p_{t-4} - \underset{(0.1596)}{\underset{(0.0087)}{0.0267}}\,\Delta\Delta e_{t-1}$$

$$- \underset{(1.0707)^{**}}{\underset{(0.0016)}{0.0005}}\,\ln\left(REER\right)_{t-1} \qquad R^2 = 0.28$$

(6b)

$$\Delta\hat{e}_t = \underset{(0.111)}{\underset{(0.0007)}{0.0006}} - \underset{(0.2195)}{\underset{(0.051)}{0.0986}}\,\Delta e_{t-2} + \underset{(0.4243)}{\underset{(0.0003)}{0.0012}}\,\xi_{1,t} + \underset{(0.0404)}{\underset{(0.0006)}{0.0016}}\,\Delta\xi_{2,t} + \underset{(0.8646)^{**}}{\underset{(0.0013)}{0.0026}}\,\xi_{3,t} - \underset{(0.2023)}{\underset{(0.001)}{0.0019}}\,\xi_{4,t} - \underset{(0.18)}{\underset{(0.0001)}{0.0006}}\,\hat{\xi}^{*}_{t-1}$$

$$\hat{\xi}^{*}_{t-1} = \left(\xi^{*}_{1} - \xi^{*}_{2} - \xi^{*}_{3} - \xi^{*}_{4} + 2\xi^{*}_{5}\right)_{t-1}; \qquad R^2 = 0.1274; \quad \overline{R}^2 = 0.112$$

Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.2789)}{\underset{(0.033)}{0.0332}} - \underset{(0.2104)}{\underset{(0.0525)}{0.139}}\,\Delta e_{t-9} - \underset{(0.2724)}{\underset{(0.007)}{0.0071}}\,\ln\left(REER\right)_{t-1} \qquad R^2 = 0.0222$$

Note: Samples used for DF-ECMs: 1976M01-2005M12; Samples for *REER* equations: 1977M01-2005M12. $\overline{R}^2$ denotes the adjusted R^2. The intercept term is kept in all models irrespective of its statistical significance in order to obtain the R^2 statistics. The statistics in the upper brackets under the coefficient estimates are the standard errors; those in the lower brackets are Hansen parameter instability test statistics. Its 5% critical value is 0.47. Statistical significance at the 5% and 1% levels are marked by * and ** respectively.

Appendix Table A3. Specific Models of (6a) and (6b) Versus ECMs of *REER*: France

(6a)

$$\Delta\hat{p}_t = \underset{(0.0003)\ (0.0192)}{0.0024} + \underset{(0.0472)\ (0.1157)}{0.1633}\ \Delta p_{t-1} + \underset{(0.0447)\ (0.0657)}{0.2965}\ \Delta p_{t-6} - \underset{(0.0148)\ (0.085)}{0.0817}\ \Delta e_t - \underset{(0.0002)\ (0.0986)}{0.0009}\ \xi_{1,t}$$
$$+ \underset{(0.00002)\ (0.4901)^*}{0.0001}\left[\xi_{1,t-5} + \xi_{1,t-6}\right] + \underset{(0.00016)\ (0.1205)}{0.0005}\ \xi_{3,t-3} + \underset{(0.0001)\ (0.2132)}{0.0004}\ \xi_{5,t-4} - \underset{(0.00009)\ (0.0209)}{0.0006}\ \hat{\xi}^*_{t-1}$$
$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 + \xi^*_3 + 0.24\,\xi^*_4 + 0.2\,\xi^*_5 + 0.2\,\xi^*_6\right)_{t-1}; \qquad R^2 = 0.7204 \qquad \overline{R}^2 = 0.7133$$

Using *REER*:

$$\Delta\hat{p}_t = \underset{(0.0132)\ (0.174)}{-0.0087} + \underset{(0.0443)\ (0.6086)^*}{0.3128}\ \Delta_3 p_{t-1} + \underset{(0.0457)\ (0.4105)}{0.1603}\ \Delta p_{t-3} + \underset{(0.0462)\ (0.1593)}{0.388}\ \Delta p_{t-6}$$
$$- \underset{(0.0039)\ (0.2664)}{0.0092}\ \Delta e_{t-1} + \underset{(0.0028)\ (0.1773)}{0.0019}\ \ln\left(REER\right)_{t-1} \qquad R^2 = 0.6257$$

(6b)

$$\Delta\hat{e}_t = \underset{(0.0007)\ (0.0573)}{0.0056} - \underset{(0.0447)\ (0.4127)}{0.155}\Delta e_{t-4} - \underset{(0.069)\ (0.0599)}{0.6915}\left[\Delta p_t + \Delta p_{t-4}\right] + \underset{(0.0001)\ (0.3084)}{0.0103}\xi_{1,t} + \underset{(0.0005)\ (0.4472)}{0.0018}\xi_{1,t-4} - \underset{(0.0003)\ (0.2187)}{0.0025}\xi_{2,t}$$
$$- \underset{(0.0003)\ (0.3293)}{0.0007}\xi_{2,t-4} - \underset{(0.0004)\ (0.0547)}{0.0014}\Delta\xi_{3,t} + \underset{(0.0003)\ (0.0913)}{0.0017}\Delta_3\xi_{3,t} - \underset{(0.0004)\ (0.1786)}{0.0015}\Delta\xi_{3,t-2} + \underset{(0.0005)\ (0.0522)}{0.0012}\xi_{3,t-4} - \underset{(0.0001)\ (0.0407)}{0.0007}\hat{\xi}^*_{t-1}$$
$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 + 0.2\xi^*_2 + 1.5\xi^*_3 + 0.6\xi^*_4 + \xi^*_5\right)_{t-1}; \qquad R^2 = 0.9629; \quad \overline{R}^2 = 0.9616$$

Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.1895)\ (0.1593)}{0.3189} - \underset{(0.5249)\ (0.1152)}{1.373}\ \Delta p_t - \underset{(0.0405)\ (0.1586)}{0.0674}\ \ln\left(REER\right)_{t-1} \qquad R^2 = 0.0356$$

Note: Samples used for DF-ECMs: 1979M01-2005M12; Samples for *REER* equations: 1980M01-2005M12. See also the note in Appendix Table A2.

Appendix Table A4. Specific Models of (6a) and (6b) Versus ECMs of *REER*: Germany

(6a)

$$\Delta p_t = \underset{(0.0002)\ (0.0794)}{0.0025} - \underset{(0.0507)\ (0.216)}{0.1201}\ \Delta p_{t-4} - \underset{(0.0234)\ (0.2537)}{0.0743}\ \Delta e_t + \underset{(0.00026)\ (0.1796)}{0.0007}\ \xi_{1,t} - \underset{(0.0001)\ (0.3281)}{0.0005}\ \left[\xi_{3,t-4} + \xi_{3,t-5}\right]$$
$$- \underset{(0.0001)\ (0.311)}{0.0005}\left[\xi_{4,t-1} + \xi_{4,t-6}\right] + \underset{(0.0001)\ (0.0453)}{0.0005}\ \xi_{5,t-6} - \underset{(0.00003)\ (0.059)}{0.0003}\ \hat{\xi}^*_{t-1}$$
$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 - 0.6\xi^*_2 + \xi^*_3 - 2.7\xi^*_5\right)_{t-1}; \qquad R^2 = 0.2458 \qquad \overline{R}^2 = 0.2309$$

Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.1292)\ (0.2253)}{0.2043} - \underset{(0.0516)\ (0.0777)}{0.102}\ \Delta e_{t-7} - \underset{(0.4237)\ (0.0553)}{1.1374}\ \Delta\Delta p_{t-1} - \underset{(0.0274)\ (0.223)}{0.043}\ \ln\left(REER\right)_{t-1} \qquad R^2 = 0.0348$$

(6b)

$$\Delta\hat{e}_t = \underset{(0.0003)\ (0.1906)}{0.0019} - \underset{(0.0076)\ (0.3011)}{0.0183}\ \Delta e_{t-1} + \underset{(0.0388)\ (0.3682)}{0.105}\ \Delta e_{t-6} - \underset{(0.0528)\ (0.3201)}{0.2083}\ \left[\Delta p_t + \Delta p_{t-3}\right] + \underset{(0.00008)\ (0.0641)}{0.0107}\ \xi_{1,t} - \underset{(0.0004)\ (0.3766)}{0.0012}\ \xi_{1,t-6}$$
$$- \underset{(0.0002)\ (0.5609)^*}{0.0032}\ \xi_{2,t} - \underset{(0.0004)\ (0.0396)}{0.0024}\ \xi_{3,t} - \underset{(0.0003)\ (0.1107)}{0.001}\left[\xi_{3,t-1} + \xi_{3,t-3}\right] - \underset{(0.0003)\ (0.1204)}{0.0012}\ \xi_{4,t-3} + \underset{(0.0003)\ (0.0484)}{0.0021}\ \xi_{5,t}$$
$$+ \underset{(0.0002)\ (0.0877)}{0.001}\left[\xi_{5,t-1} + \xi_{5,t-3}\right] - \underset{(0.0001)\ (0.0523)}{0.0004}\ \hat{\xi}^*_{t-1}$$
$$\hat{\xi}^*_{t-1} = \xi^*_{3,t-1}; \qquad R^2 = 0.9827\ ; \quad \overline{R}^2 = 0.9821$$

Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.1292)\ (0.2253)}{0.2043} - \underset{(0.0516)\ (0.0777)}{0.102}\ \Delta e_{t-7} - \underset{(0.4237)\ (0.0553)}{1.1374}\ \Delta\Delta p_{t-1} - \underset{(0.0274)\ (0.223)}{0.043}\ \ln\left(REER\right)_{t-1} \qquad R^2 = 0.0348$$

Note: Samples used for DF-ECMs of (6b): 1977M08-2005M12; Samples for all the other models: 1975M01-2005M12. See also the note in Appendix Table A2.

Appendix Table A5. Specific Models of (6a) and (6b) Versus ECMs of *REER*: Japan

(6a)

$$\Delta\hat{p}_t = \underset{(0.0469)}{\underset{(0.0002)}{0.0024}} - \underset{(0.0416)}{\underset{(0.0484)}{0.3647}}\,\Delta p_{t-2} - \underset{(0.0708)}{\underset{(0.0478)}{0.2297}}\,\Delta p_{t-3} + \underset{(0.1439)}{\underset{(0.0063)}{0.0213}}\,\Delta e_{t-1} + \underset{(0.0304)}{\underset{(0.00005)}{0.0001}}\,\Delta_4\xi_{1,t} + \underset{(0.1467)}{\underset{(0.0001)}{0.0005}}\left[\xi_{2,t} + \xi_{2,t-2}\right]$$

$$+ \underset{(0.0475)}{\underset{(0.0003)}{0.0008}}\,\xi_{3,t-5} + \underset{(0.0806)}{\underset{(0.0001)}{0.0005}}\left[\xi_{4,t} + \xi_{4,t-1} + \xi_{4,t-2} + \xi_{4,t-3}\right] - \underset{(0.0162)}{\underset{(0.0002)}{0.0006}}\left[\xi_{5,t-1} + \xi_{5,t-3}\right] - \underset{(0.0667)}{\underset{(0.00006)}{0.0006}}\,\hat{\xi}^*_{t-1}$$

$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 + 0.7\xi^*_3 - \xi^*_4 + 0.3\xi^*_5 - 0.8\xi^*_6\right)_{t-1}; \qquad R^2 = 0.3763 \quad \overline{R}^2 = 0.3437$$

Using *REER*:

$$\Delta\hat{p}_t = \underset{(0.9539)^{**}}{\underset{(0.0079)}{0.0345}} + \underset{(0.1017)}{\underset{(0.0545)}{0.1038}}\,\Delta p_{t-1} - \underset{(0.1282)}{\underset{(0.0521)}{0.2728}}\,\Delta p_{t-2} - \underset{(0.0409)}{\underset{(0.0344)}{0.1374}}\,\Delta_3\Delta p_{t-3}$$

$$+ \underset{(1.494)^{**}}{\underset{(0.005)}{0.0142}}\,\Delta\Delta e_{t-1} - \underset{(0.9779)^{**}}{\underset{(0.0016)}{0.007}}\,\ln\left(REER\right)_{t-1} \qquad R^2 = 0.2062$$

(6b)

$$\Delta\hat{e}_t = \underset{(0.0656)}{\underset{(0.0017)}{0.0037}} + \underset{(0.0476)}{\underset{(0.0493)}{0.1207}}\,\Delta e_{t-2} + \underset{(0.3366)}{\underset{(0.0005)}{0.0061}}\,\xi_{1,t} + \underset{(0.0799)}{\underset{(0.0004)}{0.0072}}\,\xi_{3,t} - \underset{(0.0427)}{\underset{(0.00039)}{0.00089}}\,\hat{\xi}^*_{t-1}$$

$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 - \xi^*_3 + 2\xi^*_4 + \xi^*_5\right)_{t-1}; \qquad R^2 = 0.3552; \quad \overline{R}^2 = 0.3455$$

Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.1346)}{\underset{(0.0625)}{0.0662}} - \underset{(0.0401)}{\underset{(0.0565)}{0.1156}}\,\Delta e_{t-6} - \underset{(0.1367)}{\underset{(0.0131)}{0.0133}}\,\ln\left(REER\right)_{t-1} \qquad R^2 = 0.0187$$

Note: Samples used for all the models: 1980M01-2005M12. See also the note in Appendix Table A2.

Appendix Table A6. Specific Models of (6a) and (6b) Versus ECMs OF *REER*: UK

(6a)

$$\Delta\hat{p}_t = \underset{(0.0459)}{\underset{(0.0004)}{0.0037}} + \underset{(0.1399)}{\underset{(0.0364)}{0.1793}}\,\Delta_2 p_{t-1} + \underset{(0.0390)}{\underset{(0.0469)}{0.211}}\,\Delta p_{t-6} + \underset{(0.1943)}{\underset{(0.00008)}{0.0003}}\,\xi_{1,t} - \underset{(0.0875)}{\underset{(0.00008)}{0.00018}}\,\xi_{1,t-1} - \underset{(0.1097)}{\underset{(0.0004)}{0.0014}}\,\xi_{3,t}$$

$$+ \underset{(0.0278)}{\underset{(0.0003)}{0.0009}}\,\xi_{3,t-2} - \underset{(0.535)}{\underset{(0.0003)}{0.0013}}\,\xi_{3,t-3} + \underset{(0.0627)}{\underset{(0.0003)}{0.0012}}\,\Delta_6\xi_{4,t} - \underset{(0.3105)}{\underset{(0.0002)}{0.0007}}\,\Delta_2\xi_{5,t} - \underset{(0.0627)}{\underset{(0.00006)}{0.0004}}\,\hat{\xi}^*_{t-1}$$

$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 - 1.8\xi^*_2 - 1.2\xi^*_3 - 0.7\xi^*_4 + 0.7\xi^*_6\right)_{t-1}; \qquad R^2 = 0.4063 \quad \overline{R}^2 = 0.3866$$

Using *REER*:

$$\Delta\hat{p}_t = \underset{(0.3919)}{\underset{(0.0124)}{0.0126}} + \underset{(0.4165)}{\underset{(0.052)}{0.2321}}\,\Delta p_{t-1} + \underset{(0.1799)}{\underset{(0.0531)}{0.1025}}\,\Delta p_{t-5} + \underset{(0.0994)}{\underset{(0.0499)}{0.2794}}\,\Delta_3\Delta p_{t-6} - \underset{(0.3929)}{\underset{(0.0028)}{0.0025}}\,\ln\left(REER\right)_{t-1}$$

$$R^2 = 0.2167$$

(6b)

$$\Delta\hat{e}_t = \underset{(0.0706)}{\underset{(0.0011)}{0.0018}} - \underset{(0.3327)}{\underset{(0.0363)}{0.07}}\,\Delta e_{t-6} - \underset{(0.1142)}{\underset{(0.1903)}{0.4614}}\,\Delta\Delta p_t + \underset{(0.461)}{\underset{(0.0004)}{0.0078}}\,\xi_{1,t} - \underset{(0.3076)}{\underset{(0.0018)}{0.0071}}\,\xi_{3,t} + \underset{(0.0257)}{\underset{(0.0017)}{0.0065}}\,\xi_{4,t}$$

$$+ \underset{(0.1531)}{\underset{(0.0016)}{0.0032}}\,\xi_{4,t-2} - \underset{(0.2534)}{\underset{(0.0014)}{0.0065}}\,\xi_{5,t} - \underset{(0.0405)}{\underset{(0.0013)}{0.003}}\,\xi_{3,t-6} - \underset{(0.0674)}{\underset{(0.0002)}{0.0009}}\,\hat{\xi}^*_{t-1}$$

$$\hat{\xi}^*_{t-1} = \left(\xi^*_1 + \xi^*_2 - 1.3\xi^*_4 - \xi^*_6\right)_{t-1}; \qquad R^2 = 0.6064; \quad \overline{R}^2 = 0.5947$$

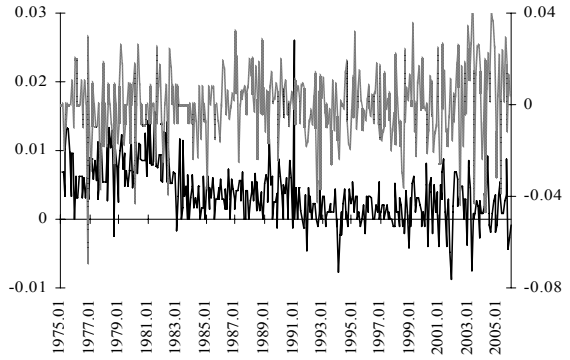
Using *REER*:

$$\Delta\hat{e}_t = \underset{(0.4680)}{\underset{(0.0821)}{0.1856}} - \underset{(0.4916)^{*}}{\underset{(0.0183)}{0.0415}}\,\ln\left(REER\right)_{t-1} \qquad R^2 = 0.0164$$

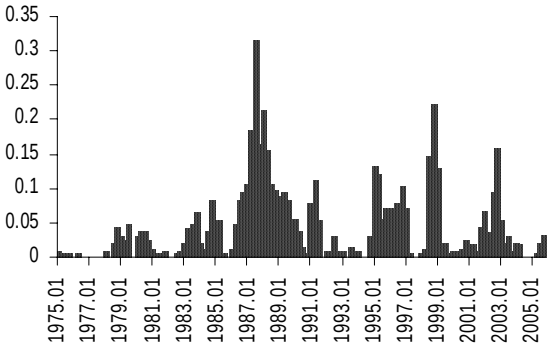
Note: Samples used for DF-ECMs: 1980M01-2005M12; Samples for *REER* equations: 1979M10-2005M12. See also the note in Appendix Table A2.

Appendix Figure A1. Canada

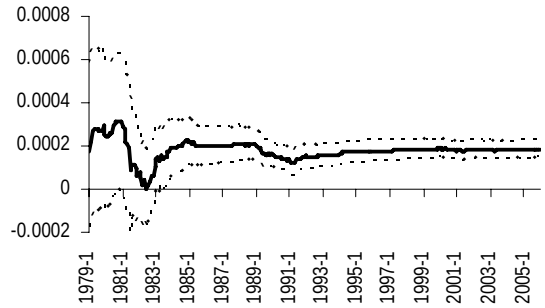
Data series: Δp_t — solid line scaled on the left axis; Δe_t — grey line scaled on the right axis



Covariation coefficient series $\{\tau_t^2\}$ of DFM (4)



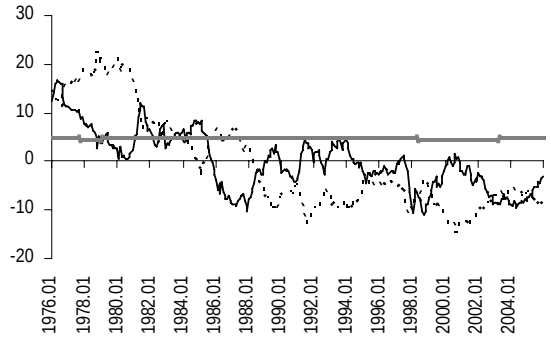
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6a); (dotted curves: the 95% confidence intervals)



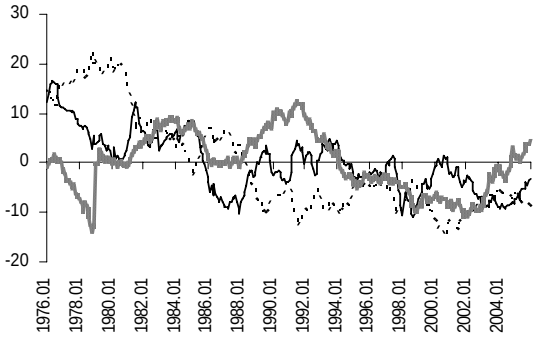
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6b); (dotted curves: the 95% confidence intervals)



Dotted line: $\hat{\xi}_t^*$ in (6a); solid line: $\hat{\xi}_t^*$ in (6b); grey line: ln(*REER*)

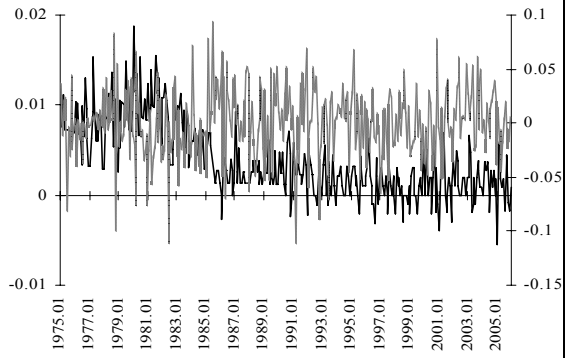


Rescaled plot of ln(*REER*) and $\hat{\xi}_t^*$ in (6a) and (6b) by the mean and standard deviation of $\hat{\xi}_t^*$ in (6b)

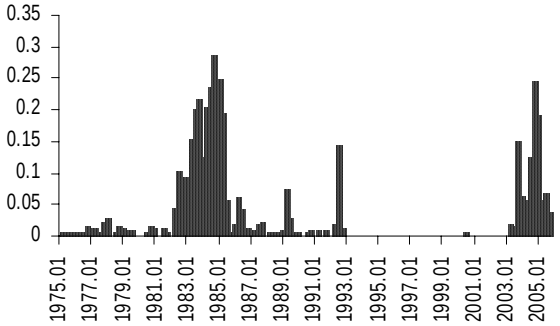


Appendix Figure A2. France

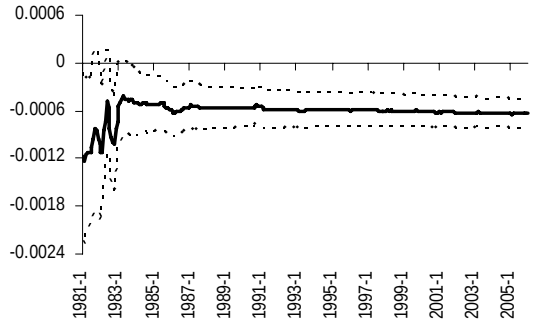
Data series: Δp_t — solid line scaled on the left axis; Δe_t — grey line scaled on the right axis



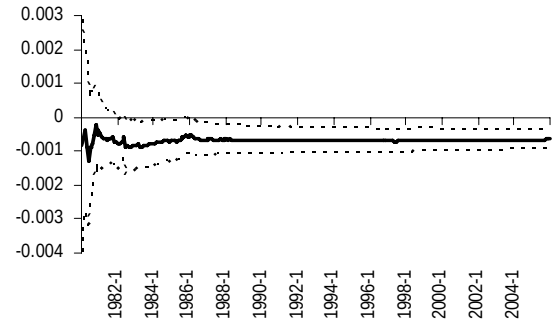
Covariation coefficient series $\{\tau_t^2\}$ of DFM (4)



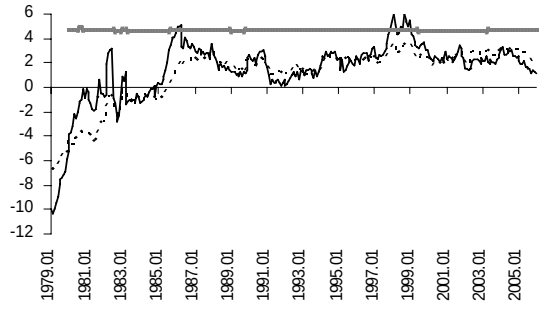
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6a); (dotted curves: the 95% confidence intervals)



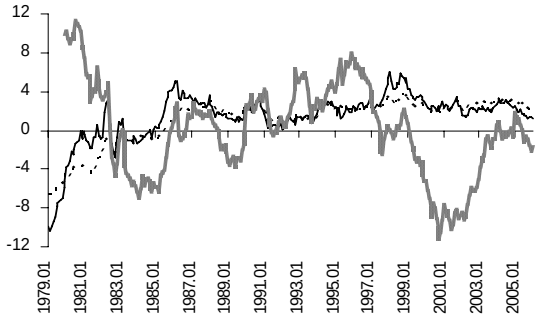
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6b); (dotted curves: the 95% confidence intervals)



Dotted line: $\hat{\xi}_t^*$ in (6a); solid line: $\hat{\xi}_t^*$ in (6b); grey line: ln(*REER*)

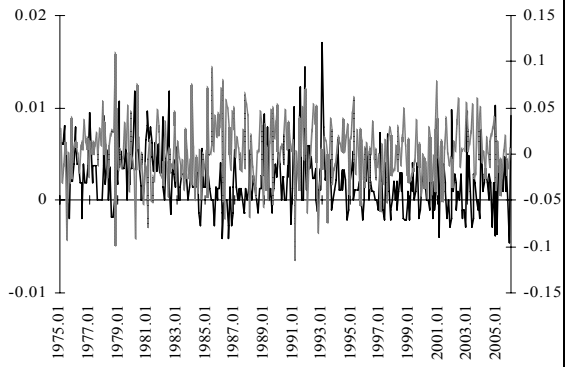


Rescaled plot of ln(*REER*) and $\hat{\xi}_t^*$ in (6a) and (6b) by the mean and standard deviation of $\hat{\xi}_t^*$ in (6b)

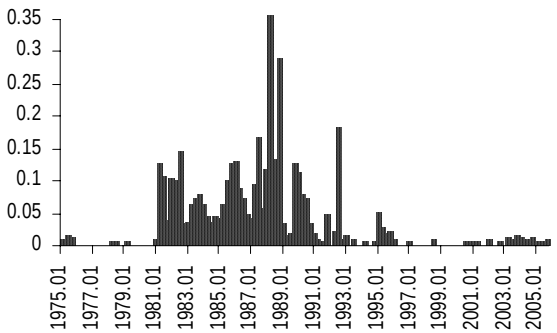


Appendix Figure A3. Germany

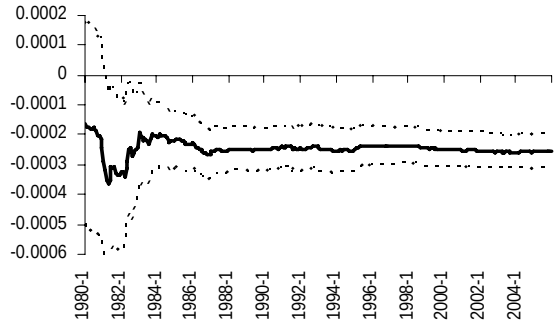
Data series: Δp_t — solid line scaled on the left axis; Δe_t — grey line scaled on the right axis



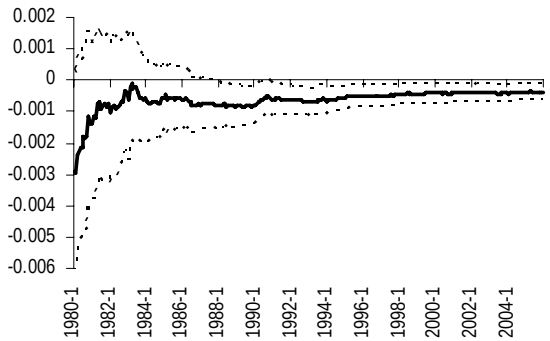
Covariation coefficient series $\{\tau_t^2\}$ of DFM (4)



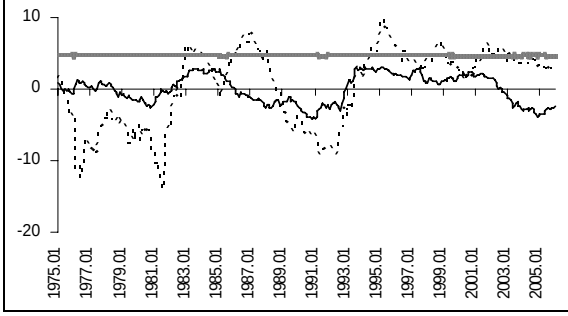
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6a); (dotted curves: the 95% confidence intervals)



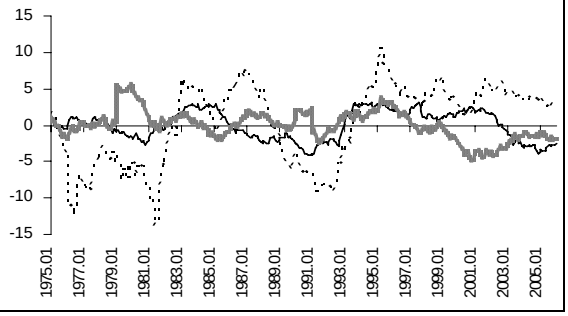
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6b); (dotted curves: the 95% confidence intervals)



Dotted line: $\hat{\xi}_t^*$ in (6a); solid line: $\hat{\xi}_t^*$ in (6b); grey line: ln(*REER*)

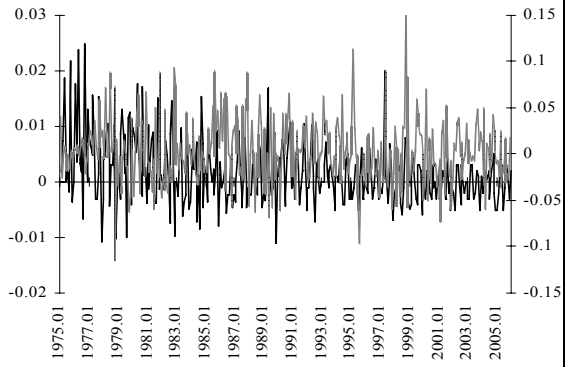


Rescaled plot of ln(*REER*) and $\hat{\xi}_t^*$ in (6a) and (6b) by the mean and standard deviation of $\hat{\xi}_t^*$ in (6b)

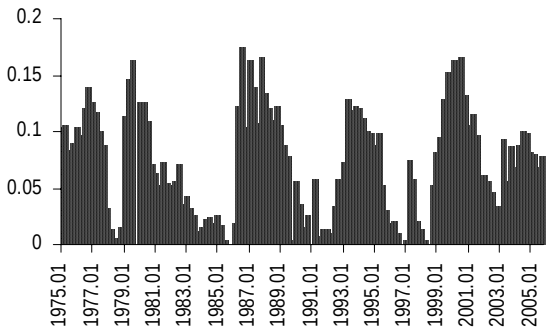


Appendix Figure A4. Japan

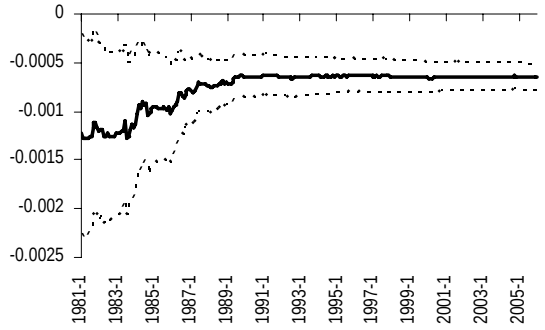
Data series: Δp_t — solid line scaled on the left axis; Δe_t — grey line scaled on the right axis



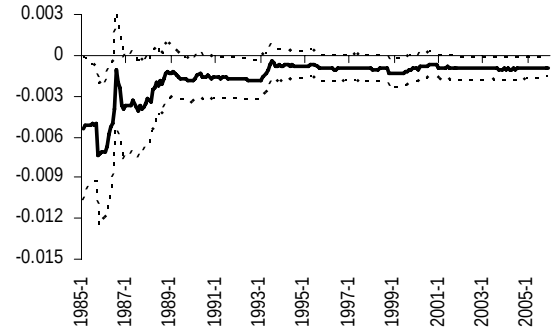
Covariation coefficient series $\{\tau_t^2\}$ of DFM (4)



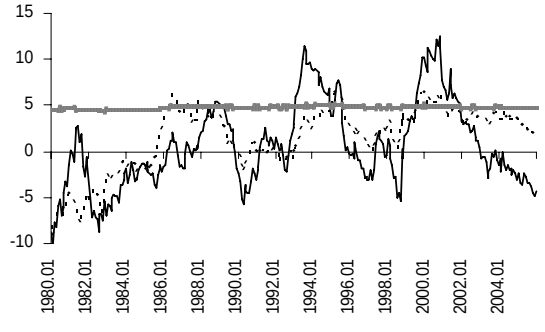
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6a); (dotted curves: the 95% confidence intervals)



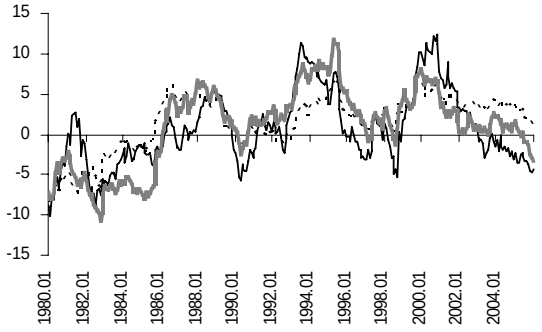
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6b); (dotted curves: the 95% confidence intervals)



Dotted line: $\hat{\xi}_t^*$ in (6a); solid line: $\hat{\xi}_t^*$ in (6b); grey line: ln(*REER*)

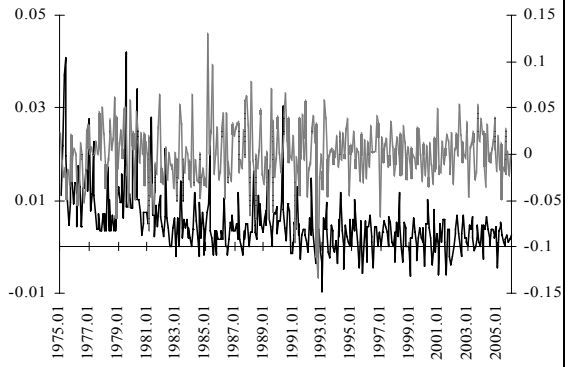


Rescaled plot of ln(*REER*) and $\hat{\xi}_t^*$ in (6a) and (6b) by the mean and standard deviation of $\hat{\xi}_t^*$ in (6b)

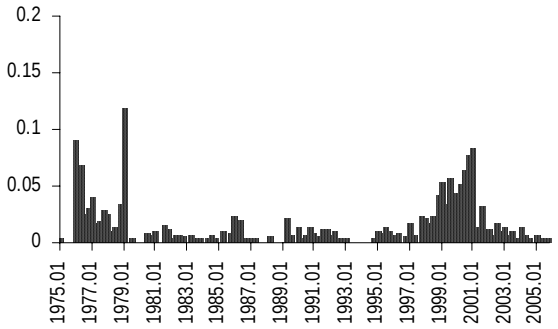


Appendix Figure A5. UK

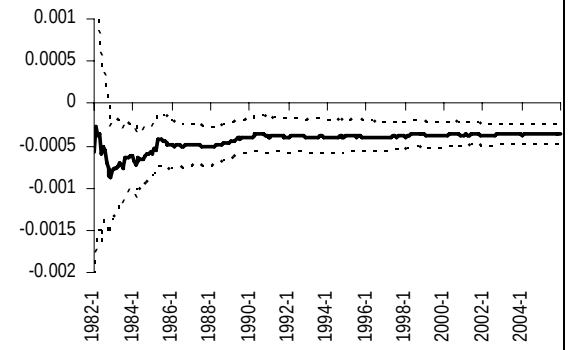
Data series: Δp_t — solid line scaled on the left axis; Δe_t — grey line scaled on the right axis



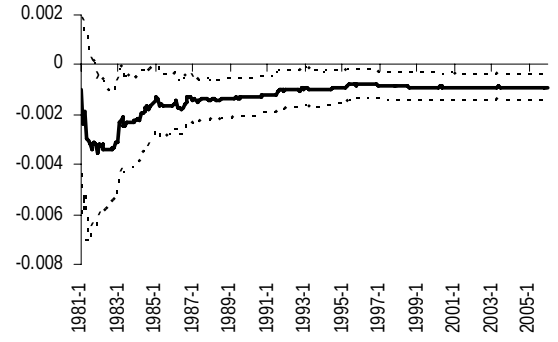
Covariation coefficient series $\{\tau_t^2\}$ of DFM (4)



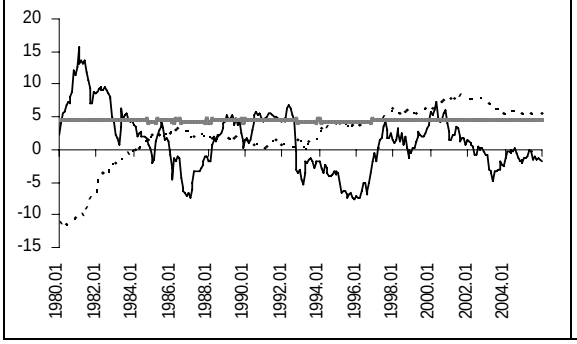
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6a); (dotted curves: the 95% confidence intervals)



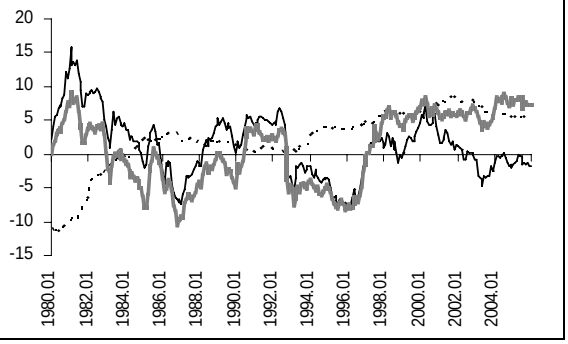
Recursive coefficient estimates of $\hat{\xi}_t^*$ in (6b); (dotted curves: the 95% confidence intervals)



Dotted line: $\hat{\xi}_t^*$ in (6a); solid line: $\hat{\xi}_t^*$ in (6b); grey line: ln(*REER*)



Rescaled plot of ln(*REER*) and $\hat{\xi}_t^*$ in (6a) and (6b) by the mean and standard deviation of $\hat{\xi}_t^*$ in (6b)



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