# Beyond Burgernomics and MacParity: Exchange Rate Forecasts Based on the Law of One Price

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**Abstract**:

Cumby (1996) suggests that (i) deviations from MacParity are stationary, (ii) relative Big Mac prices converge rapidly and (iii) provide significant information on likely future exchange rate movements. This paper examines to what extent these results can be generalised to other micro-level price data such as those provided in the UBS surveys *Prices and Earnings round the Globe*. In many respects, the MacParity findings do not constitute a special case, although they are sensitive to sample period and numeraire currency. Surprisingly, estimates based on aggregate CPI data for the same sample periods provide qualitatively similar results.

**JEL classification: F30**

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# 1. Introduction

> Some readers find our Big Mac index hard to swallow. This year, however, has been one to relish for Burgernomics. When the euro was launched at the start of the year, most forecasters expected it to rise. The Big Mac index, however, suggested the euro was overvalued against the dollar – and it has indeed fallen. (The Economist, April 1st 1999, p. 74)

> Most forecasters now predict a rebound in the euro next year. Burgernomics disagrees ... [the euro] is still a tad overvalued. (The Economist, December 9th 1999, p. 166).

As *The Economist* quite rightly claims in the first quote above, the Big Max index (BMI) was just about the only forecast[1] that predicted the euro to depreciate against the dollar after its launch. And as the second quote indicates, Burgernomics subsequently provided yet again a superior forecast of the continued fall of the euro in 2000, when compared to most financial analysts.

While Burgernomics may seem nothing but a light-hearted attempt to provide a slightly offbeat measure of exchange rate misalignment, more systematic evidence on the characteristics of the BMI was provided in a 1996 *NBER Working Paper* by Robert Cumby.[2] His analysis of BMI based real exchange rates versus the dollar came up with a number of surprising results that stood in stark contrast to some of the stylised facts emanating from the large literature on purchasing power parity (PPP) and the law of one price (LOOP). First, he discovered that nonstationarity could be easily rejected in his short panel of thirteen BMI real exchange rates. In contrast, one of the stylised facts in the PPP literature is that it is extremely difficult to get stationarity results unless one uses very long time series data or large panels (Rogoff 1996). Second, the convergence speed for the BMI parities turned out to be around twelve months – again much lower than the three to five years usually found in the PPP literature (Rogoff 1996).

The third and maybe most surprising result in Cumby's study is that deviations from relative MacParity provide significant information on likely future exchange rate movements. His analysis thus corroborates in a systematic way the more casual evi-

[1] According to its January 7th 1999 issue, "the euro [was] 13% percent overvalued against the dollar."
[2] Pakko and Pollard (1996) and Ong (1996) are two other papers that have systematically analysed the BMI data. Ong's results are supportive of the law of one price but, in contrast to the Cumby paper, focus mainly on the cross-sectional dimension of the data.

dence provided in the opening quotes on the euro/dollar exchange rate. The fact that such fundamentals as the LOOP can be of use in forecasting raises the question whether the random walk model should really be considered "the standard benchmark for empirical exchange rate performance" (Frankel and Rose 1995, p. 1691).

This paper examines to what extent Robert Cumby's (1996) striking results can be generalised. The analysis is based on individual price series published since 1970 by Union Bank of Switzerland (UBS), focusing on the same set of 13 countries so make direct comparisons possible. Methodologically the paper proceeds along similar lines to the Cumby paper, although there are some differences to the way it deals with the cross-section correlation often found in panel data. It also extends Cumby's analysis by examining whether the results are sensitive to the choice of numeraire currency, and to what extent they differ from those obtained with aggregate CPI data.

The results are ambivalent: In some ways the BMI results constitute a special case, in others not. Many of the findings in Cumby extend to the UBS price series, such as the stationarity of goods-level real exchange rates, the fairly rapid convergence to parity and the ability to forecast exchange rate changes in the medium to longer run. However, the results are sensitive to changes in the sample period and to the choice of numeraire currency. The differences between CPI and micro-level price data turn out to be a lot smaller than what one might have concluded on the basis of the stylised facts reported above.

The next section describes the dataset in more detail. Section three tests for unit roots in the UBS price data across the panel of countries considered by Cumby (1996). Section four provides estimates of the speed of convergence to the LOOP and section five analyses the potential of the UBS data in providing exchange rate forecasts at a number of different forecast horizons. Some extensions are provided in section 6, before a set of conclusions in section seven completes the paper.

## 2. Data description

UBS have so far released eleven surveys on *Prices and Earnings Round the Globe*. They have appeared every three years, the first covering prices in 1970 and the last listing prices in 2000. The price data relate to various types of goods at different levels of aggregation sampled in major cities around the world. The surveys include between 31 and 58 cities, some located in the same country. Overall, 66 cities from 55 countries have appeared in at least one issue. This paper uses data from cities located in the 14 countries originally analysed in Cumby (1996) to ensure comparability with his results. These are Australia (Sydney), Belgium (Brussels), Canada (Montreal, Toronto), Denmark (Copenhagen), France (Paris), Germany (Berlin, Düsseldorf, Frankfurt), Hong Kong, Italy (Milan, Rome), Japan (Tokyo), Netherlands (Amsterdam), Spain (Barcelona, Madrid), Sweden (Stockholm), United Kingdom (London) and the United States (Chicago, Houston., Los Angeles, New York, San Francisco). In cases where there is more than one city in a survey[3], the arithmetic mean across the available cities represents the given country.

To ensure a similar number of observations as in the Cumby study, only those price series are included that have appeared in at least nine consecutive surveys. This leaves twelve prices series for each country: (i) price level including services, (ii) clothing, (iii) household appliances, (iv) automobile, (v) food, (vi) restaurant meal, (vii) hotel stay, (viii) public transport, (ix) taxi ride, (x) rent, (xi) car service and (xii) services. For a given survey the prices refer to strictly comparable single items or baskets of goods and services, though some of the categories change over time. For instance, the cost of the basket of household goods contains a black-and-white television in the early surveys, but a colour television in the later ones. The appendix provides more detail, including examples of the typical kinds of goods and services contained in each price series (exact definitions can be found in the various UBS volumes).

Much of the recent literature on the LOOP[4] has stressed the fact that every good normally considered as a traded good – i.e. one that qualifies for LOOP considerations – normally also contains a non-traded component due to, for instance, local distribution and marketing costs. This also applies to most of the goods covered here, although there are likely to be differences as to their relative importance. For instance, most people would consider food prices to reflect traded goods to a greater extent than rents. To examine whether the degree of tradability matters, I will look for systematic differences in

[3] In those cases not all cities are available for the entire period.
[4] e.g. Engel and Rogers (1996), Haskel and Wolf (1999), Crucini et al. (2000) and Lutz (2000).

the results between the goods prices that can be put into these categories. Here the series (ii) – (v) are considered tradable and (vi) to (xi) nontradable.

The sections to follow also provide estimates for the BMI to allow a direct comparison with those for the UBS price series. The BMI data have been published annually by *The Economist* since 1986, usually in April of each year. Following Robert Cumby the first two BMI surveys in 1986 and 1987 are merged, since each only covers a limited number of (partly non-overlapping) countries and they are only a few months apart (September 1996 and January 1987). Thus the starting point for the BMI data is the 1987 survey; for countries not observed in 1987 the 1986 observation is used. The exchange rate data employed for the forecasting exercise comes from the IMF's *International Financial Statistics* database on CD-ROM (October 2000 release), updated with the most recent monthly paper issue. Again, more details are provided in the appendix.

Figure 1 provides a visual impression of the data. Define the goods-specific real exchange rate (RER),

$$r_{i,t} = \frac{P_{i,t}}{E_{i,t} P_t^{US}} , \tag{1}$$

as the common currency price in period $t$ of a given good in country $i$ relative to the United States. The exchange rate variable $E_{i,t}$ is defined as the domestic currency price of one US dollar. The thirteen graphs show $r_{i,t}$ for each UBS goods category and for the BMI. The main impression one gets with respect to the time-series dimension of the data is that, despite some fluctuations, there is not much evidence of any trends in the levels. Comparing initial and end points of the data, there only seems to be a slight increase in prices vis-à-vis the US for some of the less tradable goods, such as 'public transport', 'restaurant meal', 'hotel stay', 'rent' and the overall 'price of services'. It is also interesting to look across the different goods categories. For some goods the deviations around parity tend to be a lot larger than for others. Again, this applies to those goods one would tend to categorise as non-tradables, such as 'public transport', 'taxi ride', 'car services', 'restaurant meal', 'hotel stay' and 'rent'. However, some of these fluctuations at the individual goods and service price level appear to even out when one moves to the multiple goods and service price series, such as the categories 'price level including rent' and 'price of services'. Compared to the various UBS price series, deviations from BMI parity have been smaller and less volatile.

## 3. Testing for stationarity

If PPP and LOOP hold in the long run, deviations from parity should be temporary and the RER a stationary process. With a sufficient number of time-series observations, one can employ the standard augmented Dickey-Fuller (ADF) test for stationarity, where $H_0$: $\rho = 1$ against $H_1$: $\rho < 1$ in the testing regression equation

$$\Delta \ln r_{i,t} = \phi_i + (\rho_i - 1)\ln r_{i,t-1} + \sum_{j=1}^{k} \gamma_{ij} \Delta \ln r_{i,t-j} + \upsilon_{i,t} \qquad i = 1...N,\ \ t = 1...T\,. \tag{2}$$

Here $N$ denotes the number of countries and $T$ the number of time series observations for each country. In our case, however, $T$ is too small for the BMI and each of the UBS price series to consider using such ADF tests due to their low power. Instead, this paper follows Cumby (1996) by testing for joint stationarity across all $N$ cross-sections via the $t$-bar test due to Im, Pesaran and Shin (1997) (IPS), a panel unit root test. This is based on the average $t$-ratio

$$\bar{t}_N = \frac{1}{N}\sum\nolimits_{i=1}^{N} \frac{(\hat{\rho}_i - 1)}{s.e.(\hat{\rho}_i)} \tag{3}$$

derived by estimating (2) for the $N$ cross-sectional units. As IPS show, $\sqrt{N}\left(\bar{t}_N - \mu\right)/\sigma \rightarrow N(0,1)$, as $T \rightarrow \infty$, $N \rightarrow \infty$, where $E(\bar{t}_i) = \mu$ and $\mathrm{var}(\bar{t}_i) = \sigma^2$. IPS also provide detailed tables with critical values for different combinations of $N$, $T$ and $k$.[5]

There are some general caveats regarding the use of panel unit root tests. For instance, panel unit root tests will be biased if there are cointegrating relationships across the various country-specific series (Banerjee et al. 2001). Another bias might result from cross-sectional correlations, since panel unit root tests such as the IPS test are only valid with *iid* disturbances. Ignoring cross-sectional correlations can have serious effects on the test outcomes as O'Connell (1998) shows. For this reason IPS suggest subtracting time-specific means from the data prior to estimation, but this can only deal

[5] There are other panel unit root tests. Compared to Levin and Lin's (1993) test, the IPS test has the major advantage that it does not restrict the autoregressive coefficients to be the same across countries. IPS show that their test performs better than the Levin-Lin test. An alternative test that allows for heterogeneity is the Fisher-test due to Maddala and Wu (1999). However, while it has the merit of even allowing for differences in the specification of the test regressions, one needs to obtain the small-sample test-distribution for each individual cross-section first. So, for the purpose of this paper it was decided to employ the IPS test.

with simple cross-country correlations.[6] Yet another difficulty, as noted by Maddala and Wu (1998), is that the tests are set up for an extreme contest. Under the null hypothesis all series have a unit root whereas under the alternative none have, but of course, there are many possibilities in between these extremes.

The main aim of this paper is to examine the relative performance of the UBS price data compared to the BMI results in Cumby (1996). Thus it makes sense to employ the same test. To deal with at least some of the problems described in the preceding paragraph, the possibility of cross-sectional correlations will be dealt with explicitly. The procedure is to first implement the standard test, then redo the analysis (i) with cross-sectional means removed and (ii) using the estimated cross-sectional correlations found in the data to generate test-statistics via parametric bootstrap.

The results of these tests are presented in Table 1. First, there are the results for the twelve UBS price series. Underneath, one finds means across all UBS series, which are included as a basic descriptive device. Next there are the means for the series considered tradable and then for those considered nontradable (as described in the previous section) to examine whether tradability makes a difference. The last two rows give the results for the BMI data to allow a comparison with the UBS data. Two sets of results for the BMI data are presented, the first for the Cumby sample (up to 1996) and the second for the longest possible updated sample (up to 1998) for the same group of countries[7]. All *t*-bar statistics that are significant at the ten percent level or lower are shaded.

Column one shows the results for the IPS test with a constant in each Dickey-Fuller regression, column 2 those for the version with constant and time trend. There are no results with lagged differences as the results for the UBS series based on (2) with additional ADF terms revealed the latter to be unnecessary[8]. The ten percent critical values

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[6] This procedure may lead to other problems, as Mark (2001, Ch. 2) notes. If there is exactly one stochastic trend in the data, it will be removed by taking the time-specific means out of the data, thus biasing the test towards finding stationarity.

[7] There was no data for Belgium in 1999 and 2000 and for the Netherlands in 2000.

[8] Employing a ten percent significance level, only 1.8% of lagged first differences across the 12 UBS series were significant using an ADF(1) test with constant, and nine percent using the ADF(1) test with constant and trend. For almost all series in these two tests there was at most one country per series with a significant ADF term. Altogether, given that there were fewer significant terms than one would expect due to chance at this significane level, it was decided to stick with the simple test. This has the added advantage that more degrees of freedom can be preserved, since each ADF terms mean losing one additional initial condition. It is also worth reporting that the one exception was the 'public transport' series in the

in IPS closest to our sample specification are those for $N$ = $T$ = 10. They are –1.93, –2.06 and –2.32 at the ten, five and one percent significance levels for the test with constant, and –2.59, –2.74 and –3.03 for the test with constant and trend (see Table 4 in IPS). The results of the tests for the UBS dataset are overwhelmingly in favour of stationarity. For the simple test with a constant only, the unit root null hypothesis is rejected at the one percent significance level in all cases. For the test with constant and trend, six series lead to rejections at the one-percent, three at the five-percent and two at the ten-percent significance levels. The only non-rejection here is for 'price level incl. rent', but only marginally with a test-statistic of –2.56 compared to the 10-percent critical value –2.59. Cumby's results for the BMI are confirmed in the first row, though the evidence against a unit root is less strong for the longer BMI series where it cannot be rejected when a time trend is included.

Columns three and four repeat the exercise after the cross-section means have been subtracted from the data. The results are mostly unchanged. For the test without a time trend the null hypothesis is still decisively rejected, albeit four times at the five-percent rather than the one-percent significance level. The rejection rate deteriorates somewhat for the test with constant and trend as there are now three non-rejections for the UBS series instead of one in column before. The test results for the BMI remain basically unchanged, but remember that the use of deviations from time may only remove very simple forms of cross-sectional correlations.

The third version of the test therefore uses a parametric bootstrap to derive the empirical distributions based on the actual cross-sectional correlations found in the data. This involved the following procedure for each price series: (i) the null hypothesis was estimated for each country, i.e. $\ln r_{i,t} = \ln r_{i,t-1} + e_{it}$ and $\ln r_{i,t} = \phi_i + \ln r_{i,t-1} + e_{it}$ for the versions without and with time trend; (ii) the correlation matrix was estimated as $\hat{\mathbf{\Sigma}} = T^{-1}\sum_{t=1}^{T} \hat{\mathbf{e}}_{\mathbf{t}}\hat{\mathbf{e}}'_{\mathbf{t}}$ where $\hat{\mathbf{e}}_{\mathbf{t}} = \left(\hat{e}_{1t},...;\hat{e}_{Nt}\right)$ was the vector of residuals; (iii) next, 50+$T$ random standard normal residuals were generated $N$ times with the joint distribution given by $N(\mathbf{0}, \hat{\mathbf{\Sigma}})$, with $T$ corresponding to the time dimension of the actual series; (iv) $N$ artificial series were generated using these residuals and the process assumed under the

test with constant and trend. Here five ADF terms were significant, but the result of this test was the same as without ADF terms: a strong rejection of the unit root null hypothesis.

null (in the case with constant the drifts estimated from the sample, $\hat{\phi}_i$, were used for each series); (v) the IPS test was performed on the last *T* observations. Steps (iii) – (v) were repeated 5,000 times to generate the distribution of test-statistic from which the *p*-values for the actual tests in Table 1 were determined.

The resulting *p*-values are shown in columns 5 and 6. Controlling for all the cross-correlations in the data makes little difference for the test with constant only, as the null can still be rejected for all series: for two at the ten-percent, nine at the five-percent and one at the one-percent significance level. However, it has a more marked effect on the test with constant and trend. Here the unit root null can only be rejected for half the series (although two of them – 'public transport' and 'taxi ride' – come very close). The BMI results based on the Cumby sample are still highly significant, but adding two more time-series observations leads to non-rejection of the null for the longer BMI series, as the last row shows.

Which version of the tests is more relevant, the one with a constant only or that with both constant and trend? A tentative answer can be based on the significance of the time trends. At a ten-percent significance level 14.1% of the estimated trend coefficients are significant, a little but not a lot more than what one would expect by chance. There is also a certain degree of concentration across the goods categories. Most are found in the case of 'automobile' and 'car service' with four significant country-specific time trends, but in these two cases the test statistics with time trend are also significant. There are four more cases where two country-specific time trends are significant. Here the choice of specification would make a difference for three of them ('food', 'public transport' and 'taxi ride'). For all other categories there is at most one significant time-trend at the ten percent level – just what one would expect by chance even if there were all insignificant.

To conclude this section then, the main finding is that Cumby's results for the persistence of deviations can be generalised to other micro-level data[9]. Surprisingly, there is no difference with respect to the tradability of goods. The means of the relevant

[9] One major difference between the UBS and BMI datasets is the time span covered. While the BMI covers a twelve-year period only, the UBS data spans three decades. As discussed in Maddala and Kim (1998, Ch. 4) the time span matters when testing for stationarity. Given that both datasets have a similar number of observations, but the UBS sample a much longer time span, one should expect the tests based on the UBS series to yield more powerful tests.

test statistics across the two categories hardly differ and there is no differences with respect to the possible non-rejections when the test with constant and trend is considered.

## 4. The speed of convergence

Apart from the stationarity question, the other emphasis in the recent PPP literature has been the speed of convergence to parity. One of the surprising results in Robert Cumby's study is that the convergence speed for the BMI turns out to be much lower than those usually estimated in the PPP literature.[10] This section examines whether the UBS price series are also characterised by rapid convergence to parity. As in Cumby (1996), the speed of convergence is estimated for the joint panel of thirteen real exchange rates vis-à-vis the US via the standard autoregression[11]

$$\ln r_{i,t} = \phi_i + \rho \ln r_{i,t-1} + e_i \tag{4}$$

for each price series.

Table 2 presents the results for two different specifications of (4). The first set of results is based on the simple pooled estimator where $\phi_i = \phi$. The second version allows for country-specific dummies (i.e. fixed effects), $\phi_i$.[12] In that case the estimated adjustment pattern relates to relative rather than absolute LOOP, since the country-specific effects allow for a permanent gap in relative prices. These could be the consequence of trade barriers, transaction costs, non-traded inputs or different levels of taxation. To deal with cross-sectional correlations in the residuals and the resulting bias in the estimated standard errors, Table 2 reports panel-corrected standard errors (PCSE).[13] Furthermore, as Robert Cumby points out, in the regression model with country-specific fixed effects the estimates of the convergence coefficient are biased downwards in dynamic panels

[10] Subsequent studies using micro-level price data (Haskel and Wolf 1999 and Lutz 2000) have also found the speed of convergence to be be more rapid than the stylised facts reported for aggregate data (e.g. Frankel and Rose 1996 and Lothian and Taylor 1996).
[11] As reported in the previous section, the ADF regressions on which the IPS panel unit root tests were based indicated that there was no evidence of serial correlation. Thus higher order autoregressive terms are ignored here, which has the advantage of leaving more degrees of freedom.
[12] We do not report results where time-specific dummies are included. They fundamentally alter the interpretation of the results, since those specifications net out all unilateral movements of US prices versus the other currency prices. Thus all convergence tendencies due to changes in dollar prices or the external value of the dollar versus all other currencies would be removed. This was confirmed by preliminary estimates (available from the authors on request) where the inclusion of time dummies raised the persistence of real exchange rates substantially, compared to the estimates reported here.
[13] Given the results in Beck and Katz (1995) we refrain from using GLS.

unless $T$ is large. Table 2 therefore also provides biased-adjusted convergence estimates, $\rho$ adj., based on Nickell (1981)[14]. Because the estimates of $\rho$ for the UBS price data refer to three-year intervals and those for the BMI to annual data, they cannot be compared numerically. Instead, one should look at the estimated half-lives (in years) of shocks to the various price series. As in the previous table, there are also means for the UBS series as a whole and for those series deemed tradable and non-tradable.

The pooled data estimates suggest that half-lives for the UBS data are not out of line with the stylised facts for aggregate data, with a mean across all series of 3.3 years. In contrast, the BMI half-lives are considerable higher at around seven years. However, adding country specific dummies changes the picture considerably: half-lives are now on average 1.6 years for the UBS data (looking at the adjusted $\rho$'s). For the BMI series the reduction is even more impressive, with the Cumby sample estimate at just over half a year and the longer BMI series estimate at just over a year. Thus the extremely low estimate for the Cumby sample does not quite extend to the longer BMI sample. The last column in the table provides $p$-values of an $F$-test for the significance of the country dummies. They are highly significant (i.e. below one percent) in all cases except for 'household appliances' and 'car service', but even then they are significant at the 10% level. This implies that absolute LOOP is quite clearly rejected by the data. Note, however, that the autoregressive coefficients are less precisely estimated when country dummies are included, with hardly any of the UBS coefficients significantly different from zero. This means that in principle one cannot reject the hypothesis that all adjustment to relative parity is completed within the three year period at which the UBS data is observed.

As regards the difference between tradables and nontradables, there is some evidence here that the speed of adjustment is greater for the goods categorised as tradables, but the difference remains fairly small. The mean for tradables is 1.5 years, whereas for non-tradables it is close to two years. We can thus conclude this section by restating the main findings. Only a model of relative LOOP leads to the low half-lives reported in

[14] Nickell shows that the bias has the following form for large $N$ (eq. (18) in Nickell 1981):

$$\operatorname*{plim}_{N\to\infty}(\hat{\rho}-\rho)=\frac{-(1+\rho)}{T-1}\left\{1-\frac{1}{T}\frac{1-\rho^{T}}{(1-\rho)}\right\}\left\{1-\frac{2\rho}{(1-\rho)(T-1)}\left[1-\frac{1}{T}\frac{1-\rho^{T}}{(1-\rho)}\right]\right\}^{-1}.$$

Cumby (1996). The half-lives estimated for the UBS data are not quite as low as those for the BMI. Overall, therefore, the BMI results can again be generalised to the UBS data, even if the results are not quite as dramatically different to the PPP stylised facts as they are for the BMI.

## 5. Forecasting exchange rate changes

Having confirmed in the previous section that there is a more general tendency – beyond the BMI – for micro-level real exchange rates to adjust more quickly to shocks when what is reported in the literature for aggregate data, this section examines the nature of the adjustment more closely. Cumby (1996) also addressed this issue and found that the adjustment of nominal exchange rates mattered more than that of currency-specific prices. Unfortunately the UBS dataset does not allow us to address the relative importance since local currency prices are only observed at three-year intervals and there are no alternative data sources. However, since nominal exchange rate data can be obtained easily from standard sources, we can examine to what extent exchange rates act as an equilibrating force in the process. The key question in this section therefore is whether there is a systematic relationship between deviations from PPP/LOOP and subsequent exchange rate changes.

To examine this issue, the relationship between exchange rate changes and deviations from parity is estimated as follows,

$$\ln E_{i,t+h} - \ln E_{i,t} = \alpha + \beta \ln r_{i,t} + u_{i,t}, \tag{5}$$

where $E_{i,t+h}$ is the exchange rate $h$ months ahead of $t$. If deviations from Big Mac and UBS parities help predict future exchange rate changes, we would expect $\beta > 0$ and the estimated coefficient to be significant. In other words, an overvaluation as measured by relative goods prices would lead to a depreciation of the domestic currency relative to the dollar, which corresponds to a rise in the exchange rate. Note that this forecasting equation nests the random walk model of exchange rates often considered as the

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Cumby (1996) finds on the basis of a Monte-Carlo experiment that even for $N$ = 13 this large sample formula does an extremely good job of capturing the size of the bias.

benchmark against which to evaluate forecasting performance. If the random walk model is the best forecasting tool, the RER should not be related to future exchange rates changes and thus the corresponding null hypothesis is $H_0$: $\beta = 0$. In this case (5) becomes

$$\ln E_{i,t+h} = \alpha + \ln E_{i,t} + u_{i,t} \tag{6}$$

which has the nominal exchange rate follow a random walk with drift process.

The estimation proceeds as in the previous section. Two different versions of equation (5) are estimated for the BMI and each of the UBS price series. The estimates for the UBS series refer to the 1973 to 1997 period only. The first time-series observation (1970) is ignored since it falls into the Bretton Woods period and the last because it is too recent for there to be a sufficient number of subsequent exchange rate observations. Three different forecast horizons are considered: one for the short-term (3 months), one for the medium term (12 months) and one for the longer term (24 months).

The results are reported in Table 3. For each regression the estimate for $\beta$ and its corresponding *t*-ratio based on panel-corrected standard errors (PCSE) are shown in the tables. Consider first the UBS data. Here nearly all the estimated correlations have the expected positive sign. In terms of specification, the results are worse for the pooled data than for the estimates with country effects. The correlations are weaker numerically and less frequently statistically significant.[15] The mean $\beta$ for the pooled sample ranges from 0.03 to 0.14, depending on the forecast horizon, compared to 0.05-0.30 for the relative LOOP specification. For all 36 estimates (twelve series and three forecasting horizons) only six correlations are significant for the pooled sample at the five percent level, but 16 for the specification with country effects (looking at the shaded *t*-ratios in the table).

Considering the different forecasting horizons, the relationship between deviations from the LOOP and subsequent exchange rate changes gets stronger the longer one looks ahead. This applies to both the size and significance of the estimated $\beta$'s. This result conforms with the general notion of PPP/LOOP as pertaining more to the long than the short run. But there is even some evidence of significant forecasting power at short time horizons for the UBS series (‘clothing’, ‘restaurant meal’, ‘rent’ and ‘price of services’).

Looking across the series, there are only three for which there is not at least one significant forecasting relationship: ‘household appliances’, ‘public transport’ and ‘hotel stay’. Interestingly, there is now a noticable difference between the results for the goods categorised as tradable and non-tradable, compared to the results from the previous two sections. The mean correlations are in most cased twice as large for tradables than for non-tradables and their *t*-ratios are also larger on average. Overall, the results for the UBS data are possibly even stronger than those based on the BMI, as shown in the last two rows, since for the BMI one obtains significant relationships only for the 24 month horizon.[16]

## 6. Sensitivity analysis and interpretation

So far we have learned that Robert Cumby’s surprising results for the BMI can be generalised to other micro-level price series: there is evidence of significant mean-reversion, the estimated speed of convergence is considerably faster than what has been found for aggregate data, and deviations from the LOOP are in many cases significantly correlated with future exchange rate movements. All these findings go against the stylised facts on LOOP/PPP reported in the literature. This section addresses three further issues: do the results stand up to reference currencies other than the US dollar? Are the results really that different from those for aggregate data? Do the euro forecasts based on the BMI carry over to the other price series considered here?

*a) Changing the reference currency*

This part reports the results when the econometric analysis so far is repeated for real exchange rates relative to the German deutschmark and UK £ sterling. The detailed results can be found in Tables A1 – A3 in Appendix 2, but a summary of the main findings is provided in Table 4. The unit root null hypothesis is again rejected for most

[15] The tests are one-sided, due to the specific alternative hypothesis considered.

[16] The results for the BMI differ from Cumby in two respects. First, he reports the significant correlation for a one-year horizon in a specification with country *and* time effects. Second, he does not use panel-corrected standard erros (PCSE). We earlier gave reasons why we prefer using PCSEs to a specification with common time dummies.

of the UBS series, but somewhat less frequently than for the real exchange rates based on the dollar. The least successful results relate to the deutschmark as numeraire currency, with the £ based real exchange rates somewhere in between. The same pattern can be noted for the BMI where, as in the case of the dollar, it is again more difficult to reject the unit root for the longer BMI series than for the Cumby sample. Across all six tests (three specifications with and without trend each) and all three reference currencies the percentage of rejections is greatest for the UBS dataset (79.6%), followed by the Cumby BMI sample (66.7%), and the lowest resulting from the use of the longer BMI series (33.3%).

Next, we consider the speed of convergence. Table 4 reports the estimated half-lives across the various specifications, series and reference currencies. Again, the details relating to the real exchange rates relative to the deutschmark and sterling can be found in Appendix 2. With respect to the reference currency, the general pattern observed for the stationarity tests also applies here. The dollar based estimates exhibit the lowest half-lives, followed by sterling as numeraire, while referencing to the deutschmark gives the slowest convergence speeds. Across the series it is now the UBS dataset that has the highest half-lives. For the BMI, one again finds that the Cumby sample constitutes a somewhat extreme case which cannot be generalised to the longer BMI sample. Note, however, that all the estimates for the speed of convergence reported in Table 4 are low compared to the stylised facts reported for aggregate data.

Moving now to the correlations between deviations from LOOP and subsequent exchange rate changes, it is again the dollar based estimates that are most successful for the UBS dataset. The sterling based estimates in particular are very poor, with hardly any significant positive correlations. Table A3 in Appendix 2 also reveals that for sterling many estimated correlations are of the wrong sign, something that is not the case for the dollar and the deutschmark. In contrast to the UBS results, for the BMI series the real exchange rates relative to the deutschmark and sterling are more successful at forecasting exchange rate changes than the dollar based ones. Overall, it appears to be somewhat easier to predict exchange rates with the BMI than the various series in the UBS dataset.

*b) Comparison with CPI estimates*

While equation (5) may be useful for forecasting purposes, it does not contain an economically meaningful relationship in itself. It simply does not make sense to expect exchange rates to respond to deviations from MacParity or the parities for any of the UBS series on their own. The only way the relationship in (5) can be rationalised is to regard micro-level price deviations to be a stand-in (i.e. instrument) for the overall or macro-level degree of price divergence. Since measures of overall price levels such as the consumer price index (CPI) come in index form and are usually based on different consumption baskets, one can never rule out measurement problems as one of the potential causes for PPP failures and/or large half-lives.[17]

This leads to the question to what extent the results based on the micro-level price series considered in this paper really improve on CPI based estimates. To suggest an answer, the empirical analysis is repeated using the standard CPI series from the IMF's *International Financial Statistics*. The periods of observation are exactly matched to those for the UBS and BMI datasets (see Appendix 1 for more details). Results are reported in Table 5.

Consider first the panel unit root test based on IPS. Just as with the long BMI series, it is quite difficult to reject the unit root null hypothesis using CPI data for the same periods. For the US dollar and sterling as reference currency there is one rejection across the six specifications, for the DM none. In contrast, when CPI data is matched to the UBS periods, non-stationarity can be rejected in almost all cases with the dollar and deutschmark as reference currency and in half the cases for sterling. Overall, therefore, the rejection of non-stationarity is much easier using the longer span data, irrespective whether it is the micro-price series from the UBS dataset or CPI data. The strong evidence of stationarity for the BMI data using the Cumby sample period does neither generalise to the longer BMI series nor to the CPI data for the BMI periods.

Moving to the panel convergence and half-life estimates, the CPI based estimates for the dollar as numeraire currency are similar to those reported in the literature when the pooled model is used. When country dummies are added, the half-life esti-

[17] A unilateral change in the weighting scheme in one country, for instance, would be equivalent to a structural change. It is well-known (see e.g. Maddala and Wu 1998) that such changes, if not properly ac-

mates are reduced substantially for CPI data matched to the BMI periods. There is less of a reduction when CPI data is matched to the UBS periods. Again, just as with the micro-level price data, one obtains the lowest estimates when the dollar is the reference currency. But again, just as with the results on stationarity, it is possible to obtain results for the CPI that are fairly similar to those with micro-level price data, once the sample periods are matched exactly.

Lastly, consider the evidence on CPI based exchange rate forecasts. Again, the comparison reveals that the CPI based forecasts are of similar quality to those with micro-level price data. They may not work in the short-run (i.e. three months ahead), but for the twelve and 24-month ahead forecasts they are almost all significantly correlated with deviations from relative PPP (i.e. when country effects are included), irrespective of the reference currency. The one exception is the 12-month ahead forecast for the UBS periods with sterling as reference currency. There are even several significant correlations when the pooled sample is used, especially for the UBS sample periods.

*c) Alternative euro forecasts*

The case for using Big Mac prices to predict exchange rate changes presented at the beginning of this paper rested partly on the fact that the BMI was about the only exchange rate forecast that suggested the euro might fall after its introduction in 1999. Keeping with the general theme of this paper, it remains to be seen whether this result would have also been obtained using the UBS price series. Figure 2 provides us with some evidence on this issue. It shows the average dollar real exchange rates across the six euro currencies (Belgium, France, Germany, Italy, Netherlands and Spain) in the UBS dataset for both the BMI and six UBS price series. The six UBS series are those deemed tradable in the previous analysis plus the two aggregated price series (‘price level including rent’ and ‘price of services’).

The line depicting the BMI clearly shows what *The Economist* discovered as the euro was introduced: a clear overvaluation of the euro on the basis of this parity. However, looking at the six UBS series in 1997 – the closest observation to the introduction of the euro – one gets a mixed picture. According to two of the series (‘household appli-

counted for, will bias unit root tests towards finding non-stationarity. Thus some of the difficulties in rejecting the unit root null hypothesis reported in the literature may be exactly for this reason.

ances', 'automobile') the euro currencies were overvalued in 1997, but the four other series suggested an undervaluation. In fact, if one considers the entire period covered in Figure 2, one is struck by the contrast between the BMI parities that suggested a large overvaluation until 2000, and the UBS series that suggested undervaluation of the European currencies for most of the period. The only two periods were the two datasets were roughly in accordance were 1988 – all except the UBS 'price level including rent' series implied an overvaluation versus the dollar – and 2000 when all pointed to an undervaluation. But even in 2000 the degree of the suggested undervaluation differed enormously, the UBS series being much more extreme.

Perhaps the most striking thing – if one believes the LOOP should hold in the medium run – about the BMI shown in the Figure is the fact that it could remain overvalued for so long. This stands in stark contrast to its usefulness in forecasting the euro decline after 1999. One correct prediction does obviously not make a successful forecasting tool – throughout the preceding decade the BMI was persistently wrong about future movements in the value of the six European currencies surveyed here versus the dollar. Even if *The Economist* humorously claims,

> Burgernomics is far from perfect, but our mouths are where our money is. (The Economist, April 29th 2000, p. 103),

this may not be the most palatable investment advice.

## 7. Conclusions

This paper has examined to what extent Robert Cumby's (1996) striking results on the usefulness of relative Big Mac prices in forecasting exchange rate changes can be generalised. The analysis was based on individual price series published since 1970 by Union Bank of Switzerland (UBS), focusing on the same set of 13 countries to make direct comparisons possible. Throughout the analysis results for Big Mac prices were employed as a comparative benchmark. Methodologically the paper proceeded along similar lines to Cumby the Cumby paper. In some cases the analysis was extended, in particular concerning the cross-section correlation often found in panel data. Another extension involved analysing real exchange rates relative to two other base currencies besides the dollar. Lastly, the results for the micro-level price data were contrasted with those one would have obtained for exactly the same time periods by using aggregate CPI data.

Do the MacParity findings constitute a special case? The answer is ambivalent: they do in some ways but in others they do not. Many of the findings in Cumby extend to the UBS price series, such as the stationarity of goods-level real exchange rates, the fairly rapid convergence to parity and the ability to forecast exchange rate changes in the medium to longer run. However, it has also become clear that extending the BMI to the longest period for which there is data for the countries covered by Cumby leads to less extreme results. For instance, it becomes less easy to reject non-stationarity, and the estimated speed of convergence rises. The BMI results are also somewhat sensitive to the choice of numeraire currency. Again, the most striking results can be obtained following Cumby by referencing against the dollar. When one uses the deutschmark or sterling as the numeraire, the results still go in the same direction, but nonstationarity is less easily rejected and convergence speeds rise. However, while the ability to predict exchange rate changes deteriorates for the UBS series, it actually improves for the BMI.

Another additional issue that has been addressed in this paper is to what extent the use of micro-level price data actually makes a difference. To shed light on this issue the analysis was repeated with aggregate CPI data matched exactly to the sample periods of the UBS and BMI datasets. The overall impression is that the difference between CPI and micro-level price data results is less clearly pronounced than what might have been concluded by viewing the initial BMI evidence in Cumby (1996). For the time periods considered here, the CPI performs very similarly to the micro-level price series.

Finally, the paper returned to the opening remarks from the Economist that emphasised the successful euro forecast of the BMI when it was introduced in 1999. When contrasted with the UBS parities, the BMI based euro forecast wass rather an exception. Most UBS price series would have suggested in 1997 that European currencies were undervalued – not overvalued. Moreover, the biggest problem with the BMI based parities is that they predicted the European currencies to depreciate throughout the entire 1987-1999 period. Where the UBS and BMI series agree, however, is that by now the Euro is most definitely undervalued by LOOP standards.

# Appendix 1: Description of the Data

*1. Prices and Earnings round the Globe (UBS)*

So far there have been eleven editions with price data for July 1970, July/Aug 1973, May/June 1976, June/July 1979, March/April 1982 and 1985, April/May 1988 and 1991, 2. Quarter in 1994, 1997 and 2000. The goods are practically identical across the cities in any given survey but can vary across surveys. The following 12 series are used in this paper:

(1) *price level including rent*: cost of a weighted basket of goods and services;

(2) *clothing*: average (real exchange rate) based on six different categories of clothing;

(3) *household appliances*: cost of a set of household goods, e.g. TV, refrigerator etc.;

(4) *medium-priced automobile*: price of a typical mid-sized car;

(5) *food prices*: cost of a food basket;

(6) *restaurant meal*: price of a dinner for one or main dish;

(7) *hotel stay*: cost of a double room with bath and breakfast for two, incl. service, in a first class hotel;

(8) *public transport*: price of a one-way ride on public transport (bus, streetcar or subway) of about 10 km (6 miles) or at least 10 stops;

(9) *taxi ride*: price of a 5km ride (3miles) during daytime within city limits;

(10) *rent*: average (real exchange rate) based on six rent categories differing in terms of quality and number of rooms;

(11) *car service*: average labour costs (not including price of spare parts, if needed, and oil change) for a 15000 km (approx. 9000 miles) checkup;

(12) *services price*: cost of a weighted package of between 10 and 28 items.

*2. Big Mac Index (The Economist)*

This paper uses the 14 issues from 1986 to 1998, with data refering to: 1$^{st}$ Sep 1986, 17$^{th}$ Jan 1987, 28$^{th}$ Mar 1988, 11$^{th}$ April 1989, 30$^{th}$ April 1990, 9$^{th}$ April 1991, 10$^{th}$ April 1992, 13$^{th}$ April 1993, 5$^{th}$ April 1994, 7$^{th}$ April 1995, 22$^{th}$ April 1996, 7$^{th}$ April 1997, 6$^{th}$ April 1998. Data for the first two periods was amalgameted as described in the text. Newer survey were not used since data for the complete set of 14 countries in Cumby (1996) is only available up to 1998.

*3. Exchange rates and consumer price indices (International Monetary Fund)*

The data was extracted from the October 2000 CD-Rom verison of the *International Financial Statistics* and updated with later monthly paper issue. Data is for monthly averages.

The CPI data (code ...64...ZF..) is monthly (quarterly for Australia because no monthly data available).

The time periods of the exchange rate and CPI data were matched to the dates of the Big Mac and UBS price surveys. For the UBS sample period, the following values were employed: 1970 (July); 1973 (mean of July/August); 1976 (mean of May/June); 1979 (mean of June/July); 1982 and 1985 (mean of March/April); 1988 and 1991 (mean of April/May); 1994, 1997 and 2000 (mean of April/May/June). For the BMI sample period, the value for 1987 is for January and for the other periods it is based on the April CPI value.

# Appendix 2

**Table A 1: IPS Panel Unit Root Tests for other Reference Countries**

| | | Germany | | | | | | UK | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | *none* | | *demeaned* | | *bootstrap* | | *none* | | *demeaned* | | *bootstrap* | |
| *No* | *Price series* | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* |
| 1 | Price level incl. rent | -1.95 | -2.38 | -2.38 | -2.24 | (0.182) | (0.325) | -2.47 | -2.76 | -2.31 | -2.15 | (0.049) | (0.169) |
| 2 | Clothing | -1.74 | -3.21 | -2.71 | -2.87 | (0.313) | (0.020) | -2.57 | -2.93 | -2.62 | -2.96 | (0.035) | (0.094) |
| 3 | Household appliances | -3.62 | -3.85 | -3.21 | -3.24 | (0.001) | (0.003) | -3.60 | -3.66 | -3.19 | -3.40 | (0.001) | (0.005) |
| 4 | Automobile | -2.09 | -2.40 | -2.47 | -3.12 | (0.124) | (0.320) | -2.14 | -3.29 | -2.56 | -3.03 | (0.136) | (0.042) |
| 5 | Food | -1.92 | -2.13 | -2.31 | -2.33 | (0.195) | (0.520) | -2.38 | -2.55 | -2.22 | -2.24 | (0.043) | (0.222) |
| 6 | Public transport | -2.21 | -2.12 | -2.67 | -3.11 | (0.065) | (0.527) | -2.68 | -3.05 | -2.64 | -3.02 | (0.023) | (0.073) |
| 7 | Taxi ride | -2.41 | -2.84 | -2.25 | -3.05 | (0.024) | (0.089) | -3.95 | -4.45 | -1.94 | -2.76 | (0.003) | (0.012) |
| 8 | Car service | -2.28 | -2.03 | -2.56 | -3.25 | (0.080) | (0.578) | -2.62 | -3.33 | -2.48 | -2.96 | (0.024) | (0.053) |
| 9 | Restaurant meal | -2.66 | -2.74 | -2.32 | -2.91 | (0.019) | (0.143) | -2.15 | -2.70 | -2.40 | -2.90 | (0.068) | (0.112) |
| 10 | Hotel stay | -2.11 | -3.07 | -2.89 | -3.32 | (0.097) | (0.036) | -2.47 | -3.12 | -2.91 | -3.33 | (0.017) | (0.022) |
| 11 | Rent | -2.23 | -2.35 | -2.33 | -2.56 | (0.047) | (0.329) | -1.91 | -4.11 | -2.43 | -2.32 | (0.249) | (0.007) |
| 12 | Price of services | -3.08 | -3.95 | -3.04 | -3.14 | (0.004) | (0.004) | -2.37 | -2.54 | -3.18 | -3.48 | (0.052) | (0.230) |
| *Mean (all)* | | *-2.36* | *-2.76* | *-2.60* | *-2.93* | *0.096* | *0.241* | *-2.61* | *-3.21* | *-2.57* | *-2.88* | *0.058* | *0.087* |
| *Mean (tradables, 2-5)* | | *-2.34* | *-2.90* | *-2.68* | *-2.89* | *0.158* | *0.216* | *-2.67* | *-3.11* | *-2.65* | *-2.91* | *0.054* | *0.091* |
| *Mean (nontradables, 6-11)* | | *-2.32* | *-2.52* | *-2.50* | *-3.03* | *0.055* | *0.284* | *-2.63* | *-3.46* | *-2.47* | *-2.88* | *0.064* | *0.047* |
| BMI (Cumby) | | -1.75 | -2.36 | -2.57 | -2.87 | (0.316) | (0.337) | -2.30 | -2.26 | -2.60 | -3.04 | (0.070) | (0.418) |
| BMI (longest joint sample) | | -1.69 | -2.28 | -2.08 | -2.42 | (0.362) | (0.394) | -1.70 | -2.33 | -2.11 | -2.45 | (0.356) | (0.360) |

Notes: see Table 1.

**Table A 2: Convergence Speeds for other Reference Countries**

| | | Germany | | | | | | | | UK | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Pooled data | | | Country effects | | | | | Pooled data | | | Country effects | | | | |
| *No* | *Price series* | $\rho$ | s.e. ($\rho$) | Half life (years) | $\rho$ | s.e. ($\rho$) | $\rho$ adj. | Half life (years) | p-value CE's | $\rho$ | s.e. ($\rho$) | Half life (years) | $\rho$ | s.e. ($\rho$) | $\rho$ adj. | Half life (years) | p-value CE's |
| 1 | Price level incl. rent | 0.685 | (0.13) | 5.5 | 0.213 | (0.22) | 0.369 | 2.1 | (0.000) | 0.619 | (0.13) | 4.3 | 0.196 | (0.17) | 0.349 | 2.0 | (0.000) |
| 2 | Clothing | 0.658 | (0.12) | 5.0 | 0.435 | (0.16) | 0.587 | 3.9 | (0.102) | 0.448 | (0.15) | 2.6 | 0.137 | (0.18) | 0.253 | 1.5 | (0.000) |
| 3 | Household appliances | 0.338 | (0.15) | 1.9 | -0.003 | (0.17) | 0.097 | 0.9 | (0.000) | 0.331 | (0.13) | 1.9 | 0.034 | (0.13) | 0.138 | 1.0 | (0.000) |
| 4 | Automobile | 0.554 | (0.13) | 3.5 | 0.316 | (0.18) | 0.453 | 2.6 | (0.065) | 0.547 | (0.14) | 3.4 | 0.359 | (0.20) | 0.501 | 3.0 | (0.006) |
| 5 | Food | 0.822 | (0.09) | 10.6 | 0.517 | (0.15) | 0.682 | 5.4 | (0.004) | 0.762 | (0.10) | 7.7 | 0.240 | (0.15) | 0.367 | 2.1 | (0.000) |
| 6 | Public transport | 0.683 | (0.08) | 5.4 | 0.409 | (0.15) | 0.575 | 3.8 | (0.031) | 0.587 | (0.10) | 3.9 | 0.152 | (0.17) | 0.282 | 1.6 | (0.000) |
| 7 | Taxi ride | 0.710 | (0.09) | 6.1 | 0.237 | (0.14) | 0.379 | 2.1 | (0.000) | 0.437 | (0.15) | 2.5 | -0.145 | (0.16) | -0.052 | 0.0 | (0.003) |
| 8 | Car service | 0.395 | (0.17) | 2.2 | 0.232 | (0.19) | 0.391 | 2.2 | (0.178) | 0.306 | (0.17) | 1.8 | 0.085 | (0.18) | 0.222 | 1.4 | (0.017) |
| 9 | Restaurant meal | 0.544 | (0.12) | 3.4 | 0.284 | (0.17) | 0.416 | 2.4 | (0.010) | 0.623 | (0.11) | 4.4 | 0.312 | (0.15) | 0.449 | 2.6 | (0.001) |
| 10 | Hotel stay | 0.513 | (0.13) | 3.1 | 0.282 | (0.17) | 0.415 | 2.4 | (0.057) | 0.414 | (0.13) | 2.4 | 0.200 | (0.15) | 0.323 | 1.8 | (0.022) |
| 11 | Rent | 0.859 | (0.05) | 13.6 | 0.367 | (0.13) | 0.511 | 3.1 | (0.000) | 0.809 | (0.08) | 9.8 | 0.437 | (0.22) | 0.590 | 3.9 | (0.000) |
| 12 | Price of services | 0.415 | (0.15) | 2.4 | 0.118 | (0.19) | 0.231 | 1.4 | (0.001) | 0.507 | (0.13) | 3.1 | 0.281 | (0.16) | 0.414 | 2.4 | (0.000) |
| *Mean (all)* | | *0.598* | | *5.2* | *0.284* | | *0.426* | *2.7* | | *0.533* | | *4.0* | *0.191* | | *0.320* | *1.9* | |
| *Mean (tradables, 2-5)* | | *0.593* | | *5.2* | *0.316* | | *0.455* | *3.2* | | *0.522* | | *3.9* | *0.193* | | *0.315* | *1.9* | |
| *Mean (nontradables, 6-11)* | | *0.617* | | *5.7* | *0.302* | | *0.448* | *2.7* | | *0.529* | | *4.1* | *0.173* | | *0.302* | *1.9* | |
| BMI (Cumby) | | 0.930 | (0.04) | 9.5 | 0.360 | (0.20) | 0.520 | 1.1 | (0.000) | 0.915 | (0.04) | 7.8 | 0.310 | (0.16) | 0.461 | 0.9 | (0.000) |
| BMI (longest joint sample) | | 0.919 | (0.04) | 8.2 | 0.476 | (0.15) | 0.620 | 1.4 | (0.000) | 0.918 | (0.04) | 8.1 | 0.537 | (0.17) | 0.689 | 1.9 | (0.000) |

Notes: see Table 2.

## Table A 3: Exchange Rate Forecasts for other Reference Countries

**a) Germany**

| No | | N x T | Pooled data 3 months β | t(β) | 12 months β | t(β) | 24 months β | t(β) | Country effects 3 months β | t(β) | 12 months β | t(β) | 24 months β | t(β) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1 | Price level incl.rent | 104 | -0.01 | -0.41 | 0.01 | 0.13 | 0.14 | 1.25 | -0.01 | -0.19 | 0.11 | 0.90 | 0.51 | 3.30 |
| 2 | Clothing | 117 | 0.01 | 0.74 | 0.04 | 0.88 | 0.04 | 0.50 | 0.03 | 0.96 | 0.08 | 1.36 | 0.15 | 1.57 |
| 3 | Household appliances | 117 | 0.04 | 1.67 | 0.08 | 1.42 | 0.19 | 2.14 | 0.06 | 1.54 | 0.16 | 1.95 | 0.38 | 3.30 |
| 4 | Automobile | 117 | 0.00 | 0.08 | 0.04 | 0.65 | 0.17 | 2.21 | 0.00 | -0.12 | 0.06 | 0.70 | 0.25 | 2.31 |
| 5 | Food | 117 | 0.02 | 0.96 | 0.02 | 0.35 | 0.03 | 0.47 | 0.04 | 0.83 | 0.11 | 1.23 | 0.24 | 1.78 |
| 6 | Public transport | 117 | 0.00 | 0.39 | 0.01 | 0.65 | 0.03 | 0.84 | 0.01 | 0.61 | 0.04 | 0.97 | 0.09 | 1.41 |
| 7 | Taxi ride | 117 | 0.01 | 0.80 | 0.00 | 0.03 | 0.03 | 0.66 | 0.02 | 0.65 | 0.07 | 1.19 | 0.21 | 2.56 |
| 8 | Car service | 104 | 0.00 | 0.40 | 0.02 | 1.21 | 0.06 | 1.95 | 0.00 | 0.49 | 0.03 | 1.35 | 0.08 | 2.13 |
| 9 | Restaurant meal | 117 | 0.00 | -0.21 | 0.00 | 0.12 | 0.07 | 1.44 | 0.00 | 0.02 | 0.03 | 0.62 | 0.16 | 3.03 |
| 10 | Hotel stay | 117 | 0.01 | 0.56 | 0.03 | 0.66 | 0.10 | 1.44 | 0.01 | 0.47 | 0.07 | 1.09 | 0.19 | 2.12 |
| 11 | Rent | 117 | 0.00 | -0.12 | 0.01 | 0.70 | 0.01 | 0.21 | 0.02 | 0.93 | 0.06 | 1.36 | 0.13 | 1.80 |
| 12 | Price of services | 117 | 0.02 | 0.83 | 0.05 | 0.90 | 0.16 | 1.78 | 0.03 | 0.72 | 0.13 | 1.69 | 0.34 | 3.23 |
| *Mean (all)* | | | *0.01* | *0.47* | *0.03* | *0.64* | *0.08* | *1.24* | *0.02* | *0.57* | *0.08* | *1.20* | *0.23* | *2.38* |
| *Mean (tradables, 2-5)* | | | *0.02* | *0.86* | *0.04* | *0.83* | *0.11* | *1.33* | *0.03* | *0.80* | *0.10* | *1.31* | *0.25* | *2.24* |
| *Mean (nontradables, 6-11)* | | | *0.00* | *0.30* | *0.01* | *0.56* | *0.05* | *1.09* | *0.01* | *0.53* | *0.05* | *1.10* | *0.14* | *2.17* |
| BMI (Cumby) | | 130 | 0.01 | 0.48 | 0.04 | 1.16 | 0.09 | 2.05 | 0.09 | 1.64 | 0.39 | 2.83 | 0.72 | 5.66 |
| BMI (longest joint sample) | | 156 | 0.01 | 0.53 | 0.04 | 1.30 | 0.09 | 2.45 | 0.07 | 1.67 | 0.33 | 3.09 | 0.62 | 5.71 |

**b) UK**

| No | | N x T | Pooled data 3 months β | t(β) | 12 months β | t(β) | 24 months β | t(β) | Country effects 3 months β | t(β) | 12 months β | t(β) | 24 months β | t(β) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1 | Price level incl.rent | 104 | -0.01 | -0.24 | -0.03 | -0.71 | 0.04 | 0.53 | 0.00 | -0.01 | 0.00 | -0.05 | 0.22 | 1.78 |
| 2 | Clothing | 117 | 0.02 | 1.01 | 0.01 | 0.21 | 0.08 | 1.24 | 0.04 | 1.19 | 0.03 | 0.50 | 0.19 | 2.76 |
| 3 | Household appliances | 117 | -0.01 | -0.37 | 0.02 | 0.36 | 0.06 | 0.97 | -0.03 | -0.61 | 0.04 | 0.53 | 0.14 | 1.41 |
| 4 | Automobile | 117 | -0.01 | -0.59 | -0.02 | -0.59 | 0.02 | 0.32 | -0.03 | -0.75 | -0.03 | -0.61 | 0.01 | 0.17 |
| 5 | Food | 117 | 0.00 | 0.13 | -0.05 | -1.57 | -0.06 | -1.12 | -0.01 | -0.12 | -0.07 | -1.06 | -0.01 | -0.05 |
| 6 | Public transport | 117 | 0.01 | 0.97 | 0.02 | 1.46 | 0.02 | 0.65 | 0.03 | 1.25 | 0.09 | 3.09 | 0.11 | 2.21 |
| 7 | Taxi ride | 117 | 0.01 | 0.70 | 0.00 | 0.15 | -0.01 | -0.33 | 0.02 | 0.57 | 0.06 | 1.46 | 0.06 | 1.01 |
| 8 | Car service | 104 | -0.02 | -2.26 | -0.02 | -1.49 | 0.01 | 0.56 | -0.03 | -2.31 | -0.03 | -1.34 | 0.03 | 0.91 |
| 9 | Restaurant meal | 117 | -0.01 | -0.48 | -0.04 | -1.80 | -0.01 | -0.34 | 0.00 | -0.23 | -0.05 | -1.44 | 0.04 | 0.84 |
| 10 | Hotel stay | 117 | 0.01 | 0.56 | 0.00 | 0.00 | 0.04 | 0.66 | 0.01 | 0.48 | 0.03 | 0.62 | 0.13 | 1.77 |
| 11 | Rent | 117 | -0.01 | -0.91 | -0.01 | -0.85 | -0.02 | -0.87 | -0.02 | -0.61 | -0.03 | -0.92 | -0.03 | -0.53 |
| 12 | Price of services | 117 | -0.01 | -0.45 | 0.01 | 0.19 | 0.00 | 0.02 | -0.02 | -0.62 | 0.07 | 1.04 | 0.10 | 1.10 |
| *Mean (all)* | | | *0.00* | *-0.16* | *-0.01* | *-0.38* | *0.01* | *0.19* | *0.00* | *-0.15* | *0.01* | *0.15* | *0.08* | *1.12* |
| *Mean (tradables, 2-5)* | | | *0.00* | *0.05* | *-0.01* | *-0.40* | *0.02* | *0.35* | *0.00* | *-0.08* | *-0.01* | *-0.16* | *0.08* | *1.07* |
| *Mean (nontradables, 6-11)* | | | *0.00* | *-0.24* | *-0.01* | *-0.42* | *0.00* | *0.06* | *0.00* | *-0.14* | *0.01* | *0.24* | *0.06* | *1.04* |
| BMI (Cumby) | | 130 | 0.01 | 0.61 | 0.04 | 1.16 | 0.08 | 1.92 | 0.12 | 1.72 | 0.37 | 2.57 | 0.68 | 3.45 |
| BMI (longest joint sample) | | 156 | 0.01 | 0.58 | 0.03 | 1.09 | 0.08 | 1.87 | 0.06 | 1.17 | 0.21 | 1.73 | 0.37 | 2.16 |

Notes: see Table 3.

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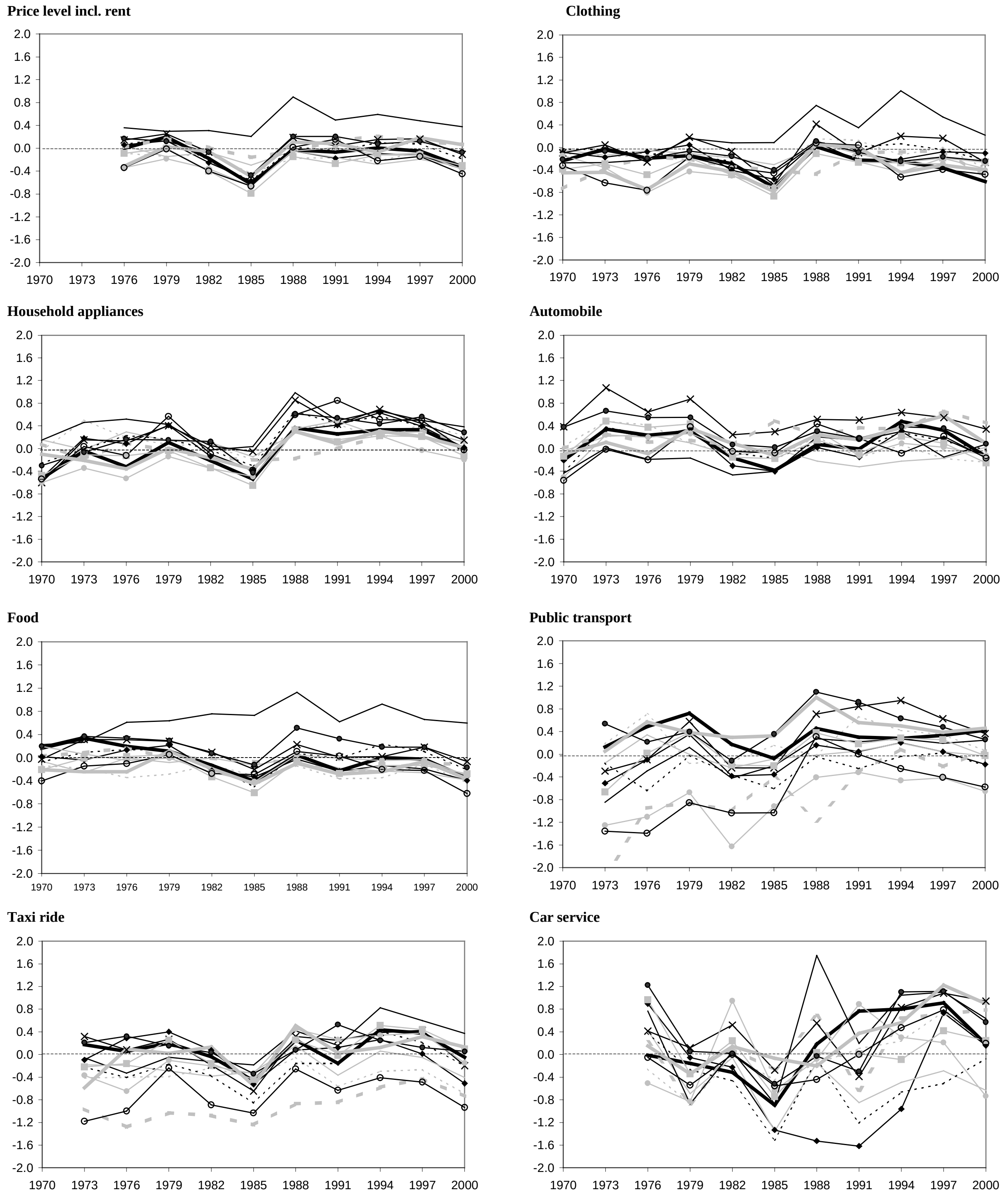
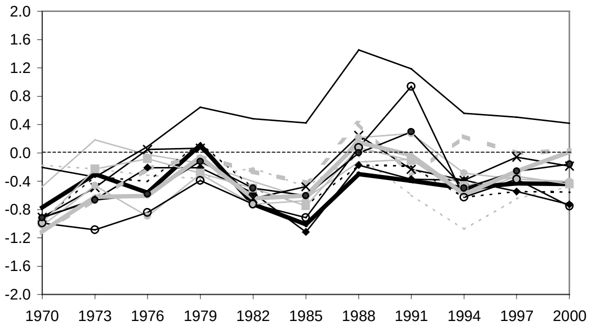
## Figure 1: Real exchange rates versus the US dollar (in logarithms)

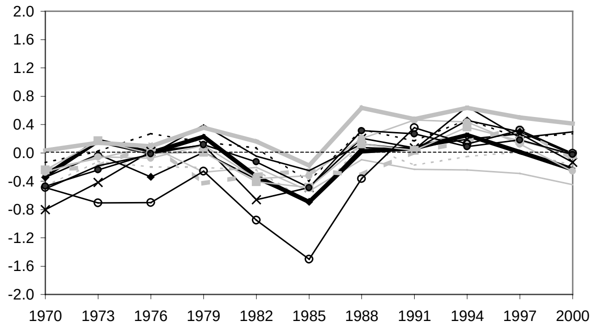


# Figure 1: continued

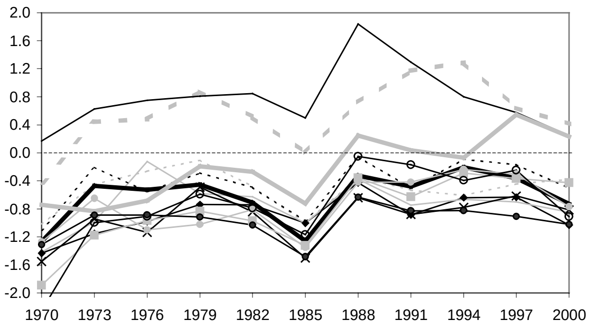
**Restaurant meal**



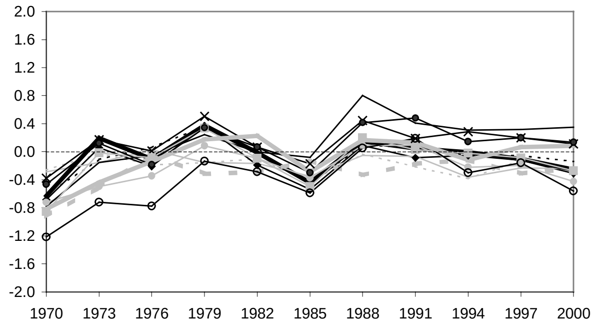

**Hotel stay**



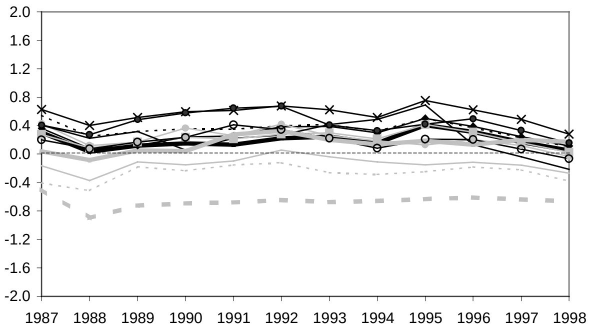

**Rent**




**Price of services**




**Big Mac Index**



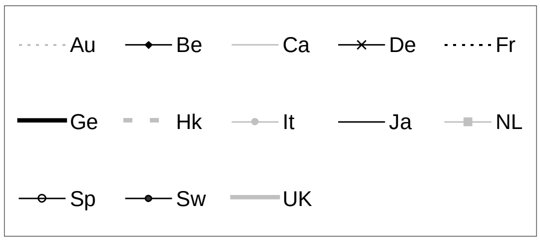
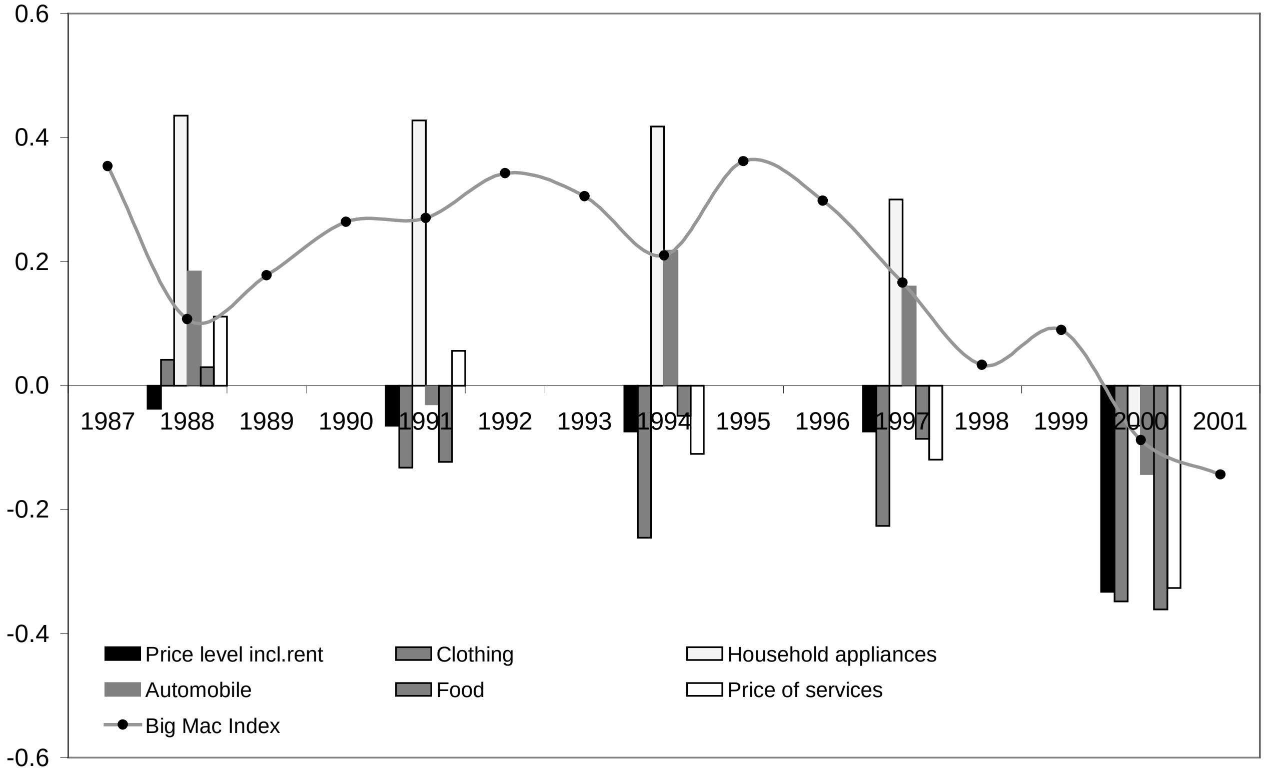



**Figure 2: European Real Exchange Rates versus the US dollar (in logarithms)**




Notes: The real exchange rates shown here averages for Belgium, France, Germany, Italy, Netherlands and Spain. The BMI average has the following missing values: Belgium (1999-2001) and Netherlands (2000-2001).

**Table 1: IPS Panel Unit Root Tests (US = reference country)**

| | | | | 1 | 2 | 3 | 4 | 5 | 6 |
|---|---|---|---|---|---|---|---|---|---|
| | *Adj. for cross-sectional correlation:* | | | *none* | | *demeaned* | | *bootstrap* | |
| | *Specification:* | | | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* |
| *No* | *Price series* | *T* | *N* | | | | | | |
| 1 | Price level incl. rent | 9 | 13 | -2.52 | -2.56 | -2.33 | -2.19 | (0.080) | (0.279) |
| 2 | Clothing | 11 | 13 | -3.01 | -3.14 | -2.55 | -2.95 | (0.013) | (0.071) |
| 3 | Household appliances | 11 | 13 | -2.70 | -2.80 | -3.35 | -3.56 | (0.031) | (0.156) |
| 4 | Automobile | 11 | 13 | -2.90 | -3.44 | -2.40 | -3.06 | (0.021) | (0.035) |
| 5 | Food | 11 | 13 | -2.41 | -2.70 | -2.22 | -2.22 | (0.064) | (0.184) |
| 6 | Public transport | 10 | 13 | -2.75 | -2.93 | -2.58 | -3.01 | (0.019) | (0.101) |
| 7 | Taxi ride | 10 | 13 | -2.61 | -3.03 | -2.24 | -2.98 | (0.042) | (0.102) |
| 8 | Car service | 9 | 13 | -2.66 | -3.76 | -2.56 | -3.00 | (0.027) | (0.017) |
| 9 | Restaurant meal | 11 | 13 | -2.96 | -3.01 | -2.22 | -2.86 | (0.010) | (0.076) |
| 10 | Hotel stay | 11 | 13 | -2.50 | -2.70 | -2.72 | -3.23 | (0.049) | (0.193) |
| 11 | Rent | 11 | 13 | -3.32 | -3.35 | -2.17 | -2.42 | (0.014) | (0.055) |
| 12 | Price of services | 11 | 13 | -3.95 | -3.63 | -2.82 | -3.23 | (0.001) | (0.031) |
| *Mean (all)* | | | | *-2.86* | *-3.09* | *-2.51* | *-2.89* | *0.031* | *0.108* |
| *Mean (tradables, 2-5)* | | | | *-2.76* | *-3.02* | *-2.63* | *-2.95* | *0.032* | *0.111* |
| *Mean (nontradables, 6-11)* | | | | *-2.80* | *-3.13* | *-2.41* | *-2.92* | *0.027* | *0.091* |
| BMI (Cumby) | | 10 | 13 | -2.52 | -4.35 | -2.52 | -2.70 | (0.041) | (0.005) |
| BMI (longest joint sample) | | 12 | 13 | -2.02 | -2.36 | -2.08 | -2.46 | (0.184) | (0.356) |

Notes: *T* and *N* denote the time series and cross-sectional dimension of each panel, while *c* and *c,t* refer to the versions with constant, and constant and time trend, respectively, of the Dickey-Fuller regressions (eq. (2) in the text). The entries in columns 1 – 4 are the IPS *t*-bar statistics (unadjusted), given in (3). The most appropriate percent critical values given in IPS (1997, Table 4) are for $N = T = 10$. For a 10% significance level they are –1.93 for the test with constant and –2.59 for the test with constant and trend. All shaded entries are significant at the 10% level or below. 'Demeaned' implies that the cross-section means for each time period were subtracted from the data prior to estimation. Columns 5 and 6 show *p*-values based on a parametric bootstrap to generate the distribution of the IPS test statistics specific to this sample. They are based on 5,000 replications of the test with simulated data that was generated using the sample variance-covariance matrix for the residuals across the country-level equations. See the text for more details.

**Table 2: Estimates of the Speed of Convergence (US = reference country)**

| | | | Pooled data | | | Country effects | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| *No* | *Price series* | *N x T* | $\rho$ | s.e. ($\rho$) | Half life (years) | $\rho$ | s.e. ($\rho$) | $\rho$ adj. | Half life (years) | p-value CE's |
| 1 | Price level incl. rent | 104 | 0.461 | (0.20) | 2.7 | 0.010 | (0.28) | 0.136 | 1.0 | (0.000) |
| 2 | Clothing | 130 | 0.381 | (0.17) | 2.2 | 0.062 | (0.20) | 0.169 | 1.2 | (0.001) |
| 3 | Household appliances | 130 | 0.348 | (0.18) | 2.0 | 0.193 | (0.22) | 0.315 | 1.8 | (0.065) |
| 4 | Automobile | 130 | 0.387 | (0.15) | 2.2 | 0.085 | (0.18) | 0.195 | 1.3 | (0.001) |
| 5 | Food | 130 | 0.743 | (0.11) | 7.0 | 0.174 | (0.20) | 0.294 | 1.7 | (0.000) |
| 6 | Public transport | 117 | 0.633 | (0.09) | 4.5 | 0.281 | (0.16) | 0.429 | 2.5 | (0.001) |
| 7 | Taxi ride | 117 | 0.604 | (0.13) | 4.1 | 0.154 | (0.22) | 0.285 | 1.7 | (0.000) |
| 8 | Car service | 104 | 0.313 | (0.18) | 1.8 | 0.107 | (0.19) | 0.247 | 1.5 | (0.057) |
| 9 | Restaurant meal | 130 | 0.448 | (0.14) | 2.6 | 0.193 | (0.17) | 0.315 | 1.8 | (0.002) |
| 10 | Hotel stay | 130 | 0.454 | (0.17) | 2.6 | 0.292 | (0.21) | 0.426 | 2.4 | (0.001) |
| 11 | Rent | 130 | 0.701 | (0.10) | 5.9 | 0.178 | (0.17) | 0.298 | 1.7 | (0.000) |
| 12 | Price of services | 130 | 0.284 | (0.15) | 1.7 | 0.074 | (0.17) | 0.182 | 1.2 | (0.000) |
| *Mean (all)* | | | *0.480* | | *3.3* | *0.150* | | *0.274* | *1.6* | |
| *Mean (tradables, 2-5)* | | | *0.465* | | *3.3* | *0.129* | | *0.243* | *1.5* | |
| *Mean (nontradables, 6-11)* | | | *0.526* | | *3.6* | *0.201* | | *0.333* | *1.9* | |
| BMI (Cumby) | | 117 | 0.903 | (0.05) | 6.8 | 0.198 | (0.19) | 0.335 | 0.6 | (0.000) |
| BMI (longest joint sample) | | 143 | 0.907 | (0.04) | 7.1 | 0.395 | (0.19) | 0.528 | 1.1 | (0.001) |

Notes: Panel corrected standard errors are shown in parentheses. '$\rho$ adj.' is based on the Nickell (1981) formula to adjust for the downward bias of $\rho$ in fixed effects regressions. Estimated half-lives are based on the estimated $\rho$ for the first two sets of estimates and $\rho$ adj. for the last two. The estimated half-lives in years were calculated as $\ln 0.5/\ln\rho$ for the annual data and as $3(\ln 0.5/\ln\rho)$ for the UBS series.

**Table 3: Exchange Rate Forecasts (US = reference country)**

| | | | **Pooled data** | | | | | | **Country effects** | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | 3 months | | 12 months | | 24 months | | 3 months | | 12 months | | 24 months | |
| *No* | | *N x T* | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) |
| 1 | Price level incl.rent | *104* | 0.05 | 1.33 | 0.14 | 1.55 | 0.30 | 2.02 | 0.09 | 1.59 | 0.30 | 2.29 | 0.64 | 3.35 |
| 2 | Clothing | *117* | 0.05 | 2.15 | 0.12 | 1.63 | 0.19 | 1.65 | 0.09 | 2.53 | 0.20 | 2.02 | 0.37 | 2.44 |
| 3 | Household appliances | *117* | 0.04 | 1.55 | 0.09 | 1.18 | 0.17 | 1.36 | 0.06 | 1.51 | 0.13 | 1.24 | 0.25 | 1.54 |
| 4 | Automobile | *117* | 0.01 | 0.24 | 0.07 | 0.98 | 0.19 | 1.66 | 0.01 | 0.12 | 0.10 | 0.92 | 0.28 | 1.64 |
| 5 | Food | *117* | 0.03 | 1.36 | 0.06 | 0.96 | 0.11 | 1.08 | 0.08 | 1.35 | 0.26 | 1.97 | 0.53 | 2.60 |
| 6 | Public transport | *117* | 0.01 | 0.72 | 0.00 | -0.08 | 0.01 | 0.16 | 0.02 | 0.88 | 0.01 | 0.22 | 0.06 | 0.54 |
| 7 | Taxi ride | *117* | 0.02 | 1.22 | 0.03 | 0.60 | 0.07 | 0.96 | 0.04 | 1.16 | 0.12 | 1.36 | 0.26 | 1.91 |
| 8 | Car service | *104* | 0.01 | 0.83 | 0.05 | 1.58 | 0.07 | 1.40 | 0.01 | 0.91 | 0.07 | 1.76 | 0.10 | 1.53 |
| 9 | Restaurant meal | *117* | 0.03 | 1.61 | 0.04 | 0.75 | 0.08 | 1.08 | 0.05 | 1.95 | 0.07 | 1.02 | 0.18 | 1.62 |
| 10 | Hotel stay | *117* | 0.03 | 1.00 | 0.06 | 0.81 | 0.15 | 1.31 | 0.03 | 0.96 | 0.09 | 1.00 | 0.22 | 1.56 |
| 11 | Rent | *117* | 0.01 | 1.14 | 0.03 | 1.42 | 0.04 | 1.08 | 0.06 | 2.13 | 0.14 | 1.86 | 0.22 | 1.94 |
| 12 | Price of services | *117* | 0.06 | 1.95 | 0.11 | 0.27 | 0.27 | 2.12 | 0.08 | 1.96 | 0.22 | 1.90 | 0.51 | 3.34 |
| *Mean (all)* | | | *0.03* | *1.26* | *0.07* | *0.97* | *0.14* | *1.32* | *0.05* | *1.42* | *0.14* | *1.46* | *0.30* | *2.00* |
| *Mean (tradables, 2-5)* | | | *0.03* | *1.33* | *0.08* | *1.19* | *0.16* | *1.44* | *0.06* | *1.38* | *0.17* | *1.54* | *0.36* | *2.06* |
| *Mean (nontradables, 6-11)* | | | *0.02* | *1.09* | *0.03* | *0.85* | *0.07* | *1.00* | *0.03* | *1.33* | *0.08* | *1.20* | *0.17* | *1.52* |
| BMI (Cumby) | | *130* | -0.01 | -0.32 | 0.02 | 0.64 | 0.07 | 1.72 | -0.06 | -0.78 | 0.19 | 1.30 | 0.51 | 3.11 |
| BMI (longest joint sample) | | *156* | -0.01 | -0.43 | 0.02 | 0.53 | 0.05 | 1.60 | -0.06 | -0.84 | 0.11 | 0.89 | 0.28 | 1.84 |

Notes: The *t*-ratios are based on panel-corrected standard errors. The shading implies significance at the five percent level or lower. See text for more details.

**Table 4: Summary of Results across 3 Reference Currencies**

| | **US dollar** | **DM** | **£ sterling** |
|---|---|---|---|
| **<u>IPS Unit root tests</u>** | | *number of rejections (total)* | |
| *UBS* | 62 (72) | 50 (72) | 60 (72) |
| *BMI (Cumby)* | 6 (6) | 2 (6) | 4 (6) |
| *BMI (long)* | 2 (6) | 1 (6) | 1 (6) |
| **<u>Half-lives (with country dummies)</u>** | | *in years* | |
| *UBS (mean)* | 1.6 | 2.7 | 1.9 |
| *BMI (Cumby)* | 0.6 | 1.1 | 0.9 |
| *BMI (long)* | 1.1 | 1.4 | 1.9 |
| **<u>Exchange rate forecasts</u>** | | *number significant (total)* | |
| *UBS* | 22 (72) | 17 (72) | 5 (72) |
| *BMI (Cumby)* | 2 (6) | 3 (6) | 4 (6) |
| *BMI (long)* | 1 (6) | 4 (6) | 3 (6) |

## Table 5: Results for CPI Data

### a) IPS Unit root test

| | | *none* | | *demeaned* | | *bootstrap* | |
|---|---|---|---|---|---|---|---|
| | | *c* | *c,t* | *c* | *c,t* | *c* | *c,t* |
| **BMI periods** | *US $* | -2.10 | -2.42 | -1.88 | -2.23 | 0.14 | 0.99 |
| | *DM* | -1.77 | -1.97 | -1.92 | -2.21 | 0.31 | 0.64 |
| | *UK £* | -1.75 | -2.01 | -1.96 | -2.34 | 0.34 | 0.59 |
| **UBS periods** | *US $* | -2.73 | -2.94 | -2.00 | -3.19 | 0.06 | 0.29 |
| | *DM* | -2.68 | -3.24 | -2.13 | -3.20 | 0.02 | 0.04 |
| | *UK £* | -1.59 | -2.78 | -2.32 | -3.31 | 0.46 | 0.15 |

### b) The speed of convergence

| | | *Pooled data* | | | *Country dummies* | | | |
|---|---|---|---|---|---|---|---|---|
| | | $\rho$ | s.e. ($\rho$) | Half life | $\rho$ | s.e. ($\rho$) | $\rho$ adj. | Half life |
| **BMI periods** | *US $* | 0.79 | 0.11 | 3.0 | 0.44 | 0.16 | 0.58 | 1.3 |
| | *DM* | 0.81 | 0.12 | 3.3 | 0.54 | 0.17 | 0.69 | 1.9 |
| | *UK £* | 0.79 | 0.13 | 3.0 | 0.51 | 0.18 | 0.66 | 1.7 |
| **UBS periods** | *US $* | 0.34 | 0.17 | 1.9 | 0.10 | 0.22 | 0.21 | 1.3 |
| | *DM* | 0.55 | 0.12 | 3.5 | 0.31 | 0.17 | 0.45 | 2.6 |
| | *UK £* | 0.60 | 0.12 | 4.1 | 0.43 | 0.18 | 0.58 | 3.9 |

### c) Exchange rate forecasts

| | | *Pooled data* | | | | | | *Country dummies* | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | 3 months | | 12 months | | 24 months | | 3 months | | 12 months | | 24 months | |
| | | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) | $\beta$ | t($\beta$) |
| **BMI periods** | *US $* | 0.02 | 0.41 | 0.15 | 1.54 | 0.16 | 1.20 | 0.06 | 0.72 | 0.42 | 2.94 | 0.59 | 3.22 |
| | *DM* | 0.02 | 0.74 | 0.13 | 1.34 | 0.18 | 1.33 | 0.06 | 1.11 | 0.38 | 2.67 | 0.63 | 3.86 |
| | *UK £* | 0.03 | 1.10 | 0.15 | 1.65 | 0.24 | 1.68 | 0.06 | 1.25 | 0.36 | 2.71 | 0.62 | 3.37 |
| **UBS periods** | *US $* | 0.06 | 1.33 | 0.26 | 2.41 | 0.55 | 3.71 | 0.09 | 1.50 | 0.32 | 2.19 | 0.69 | 3.66 |
| | *DM* | 0.01 | 0.50 | 0.16 | 2.16 | 0.38 | 3.48 | 0.04 | 0.75 | 0.19 | 1.74 | 0.50 | 3.22 |
| | *UK £* | -0.02 | -0.69 | 0.08 | 1.25 | 0.24 | 2.31 | -0.03 | -0.59 | 0.05 | 0.56 | 0.24 | 1.69 |

Notes: The BMI data refers to the 'long' series from 1987-98. See the notes to Tables 1 – 3 for more details.