Extracting Inflation from Stock Returns to Test Purchasing Power Parity*

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Abstract

We provide a novel method for extracting estimates of *realized* pure price inflation from stock returns. The key is recognizing that pure price inflation should affect nominal returns of all traded assets by exactly the same amount. The popular Fama-French three-factor model is employed to purge stock returns of real economic factors. We uncover evidence that purchasing power parity holds quite well using the extracted inflation measures.

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Purchasing power parity (PPP) is the simple proposition that prices in different countries should be equal if they are converted to the same currency. The absolute version of PPP is based on the law of one price, which maintains that arbitrage should tend to equilibrate prices of the same good at different locations. If the composition of the basket of goods used for constructing price indices is identical across countries, PPP trivially follows from the law of one price.

However, frictions to goods arbitrage such as transportation costs and other impediments to trade (the extreme being non-tradable goods such as land), inhibits cross-country price equalization. Even with such frictions, the relative version of PPP, which maintains that the *change* in price levels across countries should be the same after adjusting for the change in the exchange rate, may still hold if *relative* price changes across countries are identical. For instance, a pure money shock will change nominal prices of all goods, services and assets but relative prices will remain constant and the relative version of PPP will hold.

Although simple, the PPP hypothesis has defied empirical confirmation for decades. There seems to be little agreement about why it fails so spectacularly when taken to data.[1] The PPP puzzle in its most basic form can be described as follows. If PPP holds, then changes in the exchange rate must equal the concurrent inflation differential between two countries. Empirically, changes in exchange rates are extremely volatile, with a yearly standard deviation typically of the order of 12-13% for developed countries, while inflation differ-

[1] Rather than provide a long list of relevant references here, we point the reader to an excellent survey by Rogoff, 1996.

entials have yearly standard deviations less than 1% (see Rogoff, 1998). Moreover, although PPP says that exchange rate changes should equal inflation differentials, they are usually only weakly correlated (Rogoff, 1996).

We suspected that the "true" price level is much more volatile at high frequencies than official price indices, which are known to be sluggish in responding to monetary shocks (Dornbusch, 1976). If nominal prices were free to adjust, they would adjust to current monetary shocks, but, as was argued by Mussa (1982), would also adjust to shocks that changes *expectations* of future price levels. Mussa suggests that "information that changes these expectations can have a profound effect...., even if the current observed change that embodies this information is seemingly not very large." Consistent with the arguments of Barr and Campbell (1997), it seems possible that high frequency variation in forward looking price levels may be driven by "inflation scares" – scares that inflation may jump to very high levels or even spiral into hyper-inflation. Inflation scares, however, may turn out to be justified only rarely and, as a result, the realized inflation series constructed with "sticky" goods' prices may not appear to be volatile over long periods. This is analogous to the "peso problem" in which spot prices of a pegged currency, interspersed with infrequent but large devaluations, may appear to be smooth for long periods; the high frequency fluctuations in forward prices of such a currency may be reflecting changing expectations of the size and possibility of devaluation.

Thus, one resolution of the PPP puzzle is to test PPP using long horizons. This is extensively discussed in the literature and indeed, PPP does tend to hold much better in the long run (Rogoff, 1996). We wondered if the PPP puzzle in the short run might also be resolved if, instead of using official inflation data such as CPI, PPI or WPI, high frequency inflation data were estimated from prices of financial assets such as stock prices that are not "sticky." We attempt to settle that question in this paper. We provide a novel method for extracting pure price inflation from stock returns. Our method exploits the fact that pure price inflation should affect nominal returns of all traded assets by exactly the same amount. Our pure price inflation measure avoids the problem inherent in using official inflation data to test PPP as changes in macroeconomic price indices include effects of pure price inflation as well as real effects leading to changes in relative prices.

Of course, asset returns are influenced by factors other than pure price inflation. To extract pure price inflation from returns they must be purged of other influences. This requires an asset pricing model. We adopt the popular Fama-French three-factor model (and an extension that includes momentum as a fourth factor) to describe the return generating process. Purging stock returns of real economic factors essentially amounts to finding the return on a zero-beta portfolio, *i.e.*, a portfolio that is insensitive to real factors while responding to pure inflation. Alternatively, one could have used the nominal return on a traded risk-free asset to estimate pure price inflation. Unfortunately, there are no obvious proxies for traded risk-free assets. Notice that yields on securities such as treasury bills measure *expected* returns on a default-free asset and are influenced by *expected* inflation. To study PPP, we require estimates of *realized* returns on a risk-free asset, which will be affected by realized pure price inflation.

1 Theoretical Framework and Empirical Methodology

In this section, we first describe the factor model for stock returns and derive the empirical methodology of extracting unexpected inflation from stock returns. In the second subsection, we examine the PPP test and its alternative in the context of our procedure. Finally, we provide some simulation evidence illustrating the effectiveness of our methodology.

1.1 Extracting (Unexpected) Inflation from Stock Returns

We assume that the real stock returns in an economy follow a three-factor model:

$$r_{it} - r_{ft} = \sum_{k=1}^{3} \beta_{ik} f_{kt} + \epsilon_{it}, \tag{1}$$

where r_{it} is the real return for asset i at time t, r_{ft} is the real risk free rate, the $f_{kt}, k = 1, 2, 3$ represent real factors that describe the return generating process for securities in the economy, ϵ_{it} is a spherical disturbance and the β's are fixed sensitivity coefficients. Following Fama and French (1995), the three factors can be approximated by

1. returns on the market index in excess of the risk-free rate, $r_{Mt} - r_{ft}$,
2. returns on the zero-investment SMB portfolio, $r_{St} - r_{Bt}$, where r_{St} (r_{Bt})is the return on a small (big) cap portfolio, and
3. returns on the zero-investment HML portfolio, $r_{Ht} - r_{Lt}$, where r_{Ht} (r_{Lt}) is the return on a high (low) book/market portfolio

so that

$$r_{Mt} - r_{ft} = f_{1t} + u_{1t},$$

$$r_{St} - r_{Bt} = f_{2t} + u_{2t},$$

$$r_{Ht} - r_{Lt} = f_{3t} + u_{3t},$$

where u's represent the unknown measurement errors in estimating the true return generating factors in the economy.

These factors are created using domestic stocks only. One might argue that with integrated world capital markets, return generating factors might include some world factors as well.[2] But if world factors were included, we would have to convert returns denominated in foreign currencies into domestic units. We want to avoid having exchange rates appear in both the dependent and independent variables in our PPP tests. Furthermore, Griffin (2002) finds that domestic Fama-French factors explain much more time-series variation in returns and generally have lower pricing errors than the (Fama-French) world factor model.

Let r denote real variables and R nominal variables, then

$$R = r + \pi$$

if π measures the pure price inflation. In positing that stock returns respond to pure price inflation in this simple way, we are not ignoring the possibility that inflation, particularly unexpected inflation, can have real affects in the economy and thus could influence real stock returns. Our assumption is that such real effects on stock returns are spanned by the three

[2]For example, Ang and Bekaert (2001) find that the short rate is the only robust predictor and the U.S. rate is a stronger instrument than the local ones in predictive regressions when the local excess returns are converted into U.S. returns using spot rates, lending support for a globally integrated market. We later check the impact of adding the U.S. short rate as an additional factor in our model and find that our PPP results are virtually unchanged. Zhang (2003) also finds that the international CAPM with foreign exchange risk performs better than the single beta international CAPM and the Fama-French international three factor model. We later check whether our time series regression residuals are related to a world factor proxied by the Morgan Stanley Composite Index (MSCI), which is almost perfectly correlated with the Datastream Global Index used by Zhang (2003).

Fama-French factors.[3]

The three factor model (1) can then be rewritten as:

$$R_{it} - R_{ft} = \beta_{i1}\left[R_{Mt} - R_{ft}\right] + \beta_{i2}\left[R_{St} - R_{Bt}\right] + \beta_{i3}\left[R_{Ht} - R_{Lt}\right] - \beta_{i1}u_{1t} - \beta_{i2}u_{2t} - \beta_{i3}u_{3t} + \epsilon_{it}. \quad (2)$$

Even though the realized nominal returns on all stocks and portfolios of equity securities are observable, we do not, in fact, observe the *realized* risk-free rate R_{ft}. One way to think about the realized nominal return on a risk-free security is that it is the return on a default-free security whose ex post return is indexed by realized inflation.[4] We can, however, observe the TBill rate, which measures the *expected* nominal risk-free rate. Thus,

$$R_{ft} = \mathrm{E}_{t-1}[R_{ft}] + u_{R_{ft}} = \mathrm{E}_{t-1}[r_{ft} + \pi_t] + u_{R_{ft}} = TBill_{t-1} + u_{R_{ft}},$$

where $u_{R_{ft}}$ represents the unexpected portion of the ex-post nominal risk-free return caused by unexpected rate of inflation and the unexpected real risk-free return. It is worth pointing out that the realized real return on a TBill, measured as the difference between the contractual return on these default-free securities and the realized inflation, $Tbill_{t-1} - \pi_t$, is *not* necessarily equal to the real *risk-free* return because unanticipated inflation could make Treasury Bills risky in real terms.

[3] Boudoukh, Richardson and Whitelaw (1994) argue that expected inflation may be correlated with future production and thus affect real stock returns. We will later check the impact of adding the expected inflation as an additional factor in the model.

[4] One should not use the realized returns of securities such as the Treasury Inflation Protected Securities (TIPS) as a measure of R_{ft} because these returns are linked to official inflation data such as the changes in the CPI.

Substituting in (2) above and rewriting,

$$R_{it} - TBill_{t-1} = \beta_{i1}\left[R_{Mt} - TBill_{t-1}\right] + \beta_{i2}\left[R_{St} - R_{Bt}\right] + \beta_{i3}\left[R_{Ht} - R_{Lt}\right] + \eta_{it}, \tag{3}$$

where

$$\eta_{it} \equiv (1 - \beta_{i1})u_{R_{ft}} - \beta_{i1}u_{1t} - \beta_{i2}u_{2t} - \beta_{i3}u_{3t} + \epsilon_{it} \tag{4}$$

and where the u's and ϵ are i.i.d. normal, cross-sectionally independent, serially-uncorrelated over time and are assumed to be homoscedastic.

Equation (3) is the basis for our empirical analysis. The construct required for testing PPP is $u_{R_{ft}}$. To extract it from nominal stock returns, we implement a Fama-MacBeth cross-sectional regression based on (3). To mitigate estimation errors in the betas, we employ industrial portfolios for both domestic and foreign stocks. The first step is a time-series regression in which the nominal industrial portfolio returns in excess of the treasury bill rate are regressed on the excess market portfolio returns and the two Fama-French factor returns. This step produces the estimate of the beta's and the residual η_{it}. Notice that the regressors are not orthogonal to the residual term because of the presence of measurement error terms u's in both the regressors and the residuals. This will make the estimates of beta coefficients biased. The direction of the bias, however, cannot be ascertained very easily in a multiple regression.

The second step is a cross-sectional regression, without an intercept term, based on equation (4), which is carried out at each time t. In this step, the residual from the time-series regression for each portfolio i, η_i, is the dependent variable, and $\left(1 - \hat{\beta}_{i1}\right)$ and $-\hat{\beta}_{ik}$ $(k = 1, 3)$ are the explanatory variables. The parameter estimate associated with $\left(1 - \hat{\beta}_{i1}\right)$

is then an estimate of $u_{R_{ft}}$ for each t. By suppressing the intercept, this regression is not degenerate even though $\left(1-\hat{\beta}_{i1}\right)$ and $\hat{\beta}_{i1}$ are perfectly related.

We then construct the series $TBill_{t-1}+\hat{u}_{R_{ft}} = \hat{R}_{ft} = \hat{r}_{ft}+\hat{\pi}_t$ for both foreign and domestic countries. The difference in the two series provides an estimate of the ex post inflation differential plus the difference between the ex post real risk-free rates which we assume is random and iid.[5] This assumption can be defended in two ways. First, in Irving Fisher's world of monetary neutrality and most general equilibrium models, the *real* interest rates are generally determined by the fundamentals such as the productivity of the economy and are independent of realized pure inflation and foreign exchange rates. Second, although several recent papers find positive relationship between real interest rate differentials and the spot rate changes over the long run (3-5 years),[6] there is virtually no covariation between the two variables in the horizon less than one year. Since we are interested in examining the PPP in the short run at the monthly, bi-monthly and quarterly horizon, the contamination of the realized inflation measure by the realized real interest rate differentials will not bias our estimate.

1.2 Testing PPP

The final step is the test of the PPP hypothesis. The relative PPP hypothesis implies that

$$\Delta s_t = \Delta \pi_t \equiv \pi_t^* - \pi_t \tag{5}$$

[5] Adler and Lehnman (1983) assumes that ex ante real interest rate differential is constant.
[6] See, for example, Edison and Pauls (1993), Baxter (1994), and Bleaney and Laxton (2003).

where s_t denote the logarithm of the nominal exchange rate defined as the price of domestic currency in units of foreign currency (e.g., yen per dollar) and π_t^* and π_t denote the change in log price indices and thus are the inflation rates in foreign and domestic countries respectively.

To test relative PPP, we write:

$$\Delta s_t = \delta_t + \gamma \Delta \pi_t \tag{6}$$

where δ_t is orthogonal to $\Delta \pi_t$. We pose PPP as the null hypothesis that $\delta_t = 0$ and $\gamma = 1$. The alternative $\delta_t \neq 0$ suggests that exchange rates move for other reasons as well which contribute to the higher volatility of exchange rate changes, and the alternative $\gamma > 1$ incorporates the possibility of exchange rate overshooting models (Dornbusch, 1976).

We test the relative PPP hypothesis by regressing the estimated inflation differential on the change in the nominal exchange rate, *i.e.*,

$$R^*_{ft} - R_{ft} = a + b\Delta s_t + \epsilon_t \tag{7}$$

which is equivalent to

$$\Delta \pi_t = a + b\Delta s_t + \epsilon'_t$$

where

$$\epsilon'_t \equiv \epsilon_t + (r_{ft} - r^*_{ft}).$$

Notice that because our method extracts an estimate of inflation with significant noise, the extracted inflation differential will be the dependent variable in our regression. The regression coefficient:

$$b = \frac{\text{Cov}(\Delta \pi_t, \Delta s_t)}{\text{Var}(\Delta s_t)} = \gamma \frac{\text{Var}(\Delta \pi_t)}{\text{Var}(\Delta s_t)}. \tag{8}$$

From (6), we know that

$$\text{Var}(\Delta s_t) = \text{Var}(\delta_t) + \gamma^2 \text{Var}(\Delta \pi_t),$$

which implies that

$$\gamma = \frac{\sigma(\Delta s_t)}{\sigma(\Delta \pi_t)} \sqrt{1 - \frac{\text{Var}\delta_t}{\text{Var}\Delta s_t}}. \tag{9}$$

From (8) and (9),

$$b = \frac{\sigma(\Delta \pi_t)}{\sigma(\Delta s_t)} \sqrt{1 - \frac{\text{Var}\delta_t}{\text{Var}\Delta s_t}} \leq \frac{\sigma(\Delta \pi_t)}{\sigma(\Delta s_t)}.$$

Under the null hypothesis that PPP holds, $b = 1$. Under the alternative of either exchange rate overshooting $\gamma > 1$, or excess exchange rate volatility unrelated to inflation differential $\text{Var}\delta_t > 0$, the regression coefficient $b < 1$. To estimate an approximate order of magnitude for b under these alternatives, assume that volatility of exchange rates is of the order of 10% per year and that volatility of inflation differential is of the order of 1% per year. In that case, $b \leq 0.1$.

Some readers may be wondering if our test of the PPP in (7) essentially amounts to testing the Uncovered Interest Rate Parity relation which states that the *expected* change in the exchange rates must equal the difference in nominal default-free interest rates. In our notation, this is equivalent to testing the following relation:

$$TBill^*_{t-1} - TBill_{t-1} = \text{E}_{t-1}\left[R^*_{ft} - R_{ft}\right] = \text{E}_{t-1}\Delta s_t.$$

The Uncovered Interest Rate Parity relation is usually tested by regressing the *realized* changes in the exchange rates on the nominal default-free interest rate differential (measured by the difference in TBill rates) which are determined ex-ante and thus incorporate only the *expected* inflation differential. Notice, however, that our test of the PPP in (7) regresses that the *realized* changes in the nominal risk-free rates, which incorporate *realized* inflation differential, on the *realized* changes in the exchange rates.

1.3 Some Simulation Evidence

If the true beta coefficients were regressors in the second-stage cross-sectional regressions, the estimate of $u_{R_{ft}}$ would be unbiased. But there are two potential problems. First, estimates of betas from the time series regressions could be biased. Second, even if the beta estimates were unbiased, they are estimated with some noise,which induces an errors-in-variables problem in the second stage cross-sectional regressions. We now provide simulation evidence about the extent of these potential difficulties.

Simulation procedure:

First notice that

$$r_{Mt} - r_{ft} = R_{Mt} - R_{ft}$$

is not observable because we don't observe the realized risk-free rate R_{ft} but instead we observe the expected risk-free rate represented by the TBill rate. If we let f'_{1t} denote the observable market excess return $R_{Mt} - Tbill_{t-1}$, then it is easy to see that

$$\begin{aligned} f'_{1t} &= R_{Mt} - TBill_{t-1} \\ &= R_{Mt} - R_{ft} + (R_{ft} - TBill_{t-1}) \\ &= r_{Mt} - r_{ft} + u_{R_{ft}} \\ &= f_{1t} + u_{1t} + u_{R_{ft}}, \end{aligned}$$

where $u_{R_{ft}}$ represents the unexpected movement in the nominal risk-free rate. We let f'_{1t} denote the proxy for the first factor f_{1t} in the return generating model (1). The proxies for the second and third factors respectively are the returns on the SMB and the HML portfolios which are observable, i.e.,

$$r_{St} - r_{Bt} \equiv f'_{2t} = f_{2t} + u_{2t},$$

$$r_{Ht} - r_{Lt} \equiv f'_{3t} = f_{3t} + u_{3t}.$$

Our goal in the simulation exercise is to create factors f_{kt}, and *independent* errors u_{kt} and $u_{R_{ft}}$ in such a manner that the means and variances of the factor proxies match those observed in data.

We create the three factors f_{1t}, f_{2t}, f_{3t} and the noise terms associated with each of these factors u_{1t}, u_{2t}, u_{3t} and the noise in the risk-free rate $u_{R_{ft}}$ in such a way that the volatility of the independent noise terms is a fraction $\{10\%, 30\%, 50\%, 70\%, 90\%\}$ of the volatility of the proxies for the three factors. Notice that some combinations of volatility of u_{1t} and $u_{R_{ft}}$ will not be feasible since the noise of the two noise terms cannot exceed 100% of the volatility of f_{1t}. The factors are then created as follows:

$$f_{1t} = \bar{f}'_{1t} + \sqrt{1 - \rho^2 - \omega^2} \left[f'_{1t} - \bar{f}'_{1t} \right],$$

$$f_{kt} = \bar{f}'_{kt} + \sqrt{1 - \rho^2} \left[f'_{kt} - \bar{f}'_{kt} \right], \quad k = 2, 3$$

where

$$\bar{f}'_{kt} = \frac{1}{T} \sum_{t=1}^{T} f'_{kt}, \quad k = 1, 2, 3,$$

ρ represents the volatility (standard deviation) of u_{kt} as a fraction of the volatility of f_{kt} and ω represents the volatility of $u_{R_{ft}}$ as a fraction of the volatility of f_{1t}. This ensures that the mean of the factor f_{kt} equals the mean of the observed proxy for that factor f'_{kt}, and that the volatility of $f_{1t}+u_{1t}+u_{R_{ft}}$ and $f_{kt}+u_{kt}, k=2,3$ matches the mean and (approximately) the volatility of the observed proxies for the three factors.

We create a series for the *realized* nominal risk-free rate in the economy as follows:

$$R_{ft} = TBill_{t-1} + u_{R_{ft}}.$$

We then simulate the realized nominal returns on 30 U.S. and 33 Japanese industry portfolios using the following return generating process:

$$R_{it} - R_{ft} = \alpha + \sum_{k=1}^{3} \beta_{ik} f_{kt} + \epsilon_{it},$$

where α and β_{ik} are estimates obtained by running the realized U.S. and Japanese market excess return on the thirty U.S. and 33 Japanese industry portfolios on the realized three Fama-French factor proxies and the volatility of ϵ_{it} is set to be the same as that of the estimate of volatility of the residual from these Fama-French industry regressions.

Once the data are generated we extract an estimate of $u_{R_{ft}}$ and regress this estimate on the original $u_{R_{ft}}$ to see how well the method is able to extract the true value.

The above procedure is repeated ten times for each combination of the volatility specification, and the average slope coefficient estimate, its standard error, and the R^2 are reported in Table I.[7]

Notice that our empirical method is indeed quite effective in extracting unbiased estimates of $u_{R_{ft}}$ from stock returns when ρ and ω are low; the slope coefficients are close to 1 with small standard errors. The adjusted R^2s are not always high indicating that $u_{R_{ft}}$,

[7] We also simulated the process for one hundred times, and there is virtually no difference in the results.

although unbiased, is extracted with considerable noise. Notice also, that for any given ρ, the effectiveness first improves and then deteriorates as ω increases. Our method produces poor estimates only when both ρ and ω are high.

2 The Data

We use three sets of data to carry out the relative PPP test in the short run. The first set of data includes the industrial returns and the Fama-French three factors, namely, the excess market return, the return on a zero-investment portfolio of SMB, and the return on a zero-investment portfolio of HML, in the United States, the United Kingdom, Japan and Germany. The second set of data are the foreign exchange rates defined as the foreign currency (Yen, British Pound, and German Mark) per U.S. dollar. The third set of data are government CPI inflation measures.

All the U.S. industrial returns, the T-bill rate, and the three Fama-French factors are from Kenneth French's web-site[8] with the sample ranging from July 1926 to December 2000. The UK, German, and Japanese industrial returns are from Datastream. For the UK and Germany, total returns including dividends are available while only capital gains returns are available for Japan. The market returns for the UK, Japan, and Germany are constructed using the total (including dividend) market returns from Datastream, and the SMB and HML factor returns are constructed from the raw data from Datastream as well.[9] While the Tbill rates for the UK and Japan are from Datastream, the Tbill rate for Germany is from Bloomberg. The sample period for the UK is from January 1986 to December 1999 with 168

[8] http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/

[9] Xiaoyan Zhang and Kent Daniel kindly provided the SMB and HML data to us.

monthly observations, for Japan it is from May 1983 to December 1999 with 200 monthly observations, and for Germany it is from January 1988 to December 1999 with 144 monthly observations.

The change in the foreign exchange rate is calculated from the end of month to the end of month using the daily foreign exchange rate kindly provided by Pacific Foreign Exchange Rate Service.[10] We also calculated the foreign exchange rate changes from the beginning of the month to the beginning of the month, and the empirical results were virtually unchanged.

Table II provides the summary statistics of these three data sets. The market return for U.S., UK and Germany is at around 1.3-1.4% per month with a sample standard deviation of around 4-5% per month, but the Japanese market return has much lower mean and slightly higher volatility. The other two factors are much smaller in mean returns. The one-month treasury bill rates vary from a high of 0.71% per month for the UK to a low of thirty basis points in Japan. The official measures of inflation are calculated from the CPI data, and the monthly average rates are around 0.27%, 0.33%, 0.19% and 0.10% for the U.S., UK, Germany and Japan, respectively.

As is well known in the literature, there is a striking difference in the magnitude of the volatility of treasury bill rates, the CPI inflation rates, and the spot rate changes. For example, the sample volatility of treasury bill rates is smaller than thirty basis points and the CPI inflation rates has a sample volatility smaller than fifty basis points for all four countries. In contrast, the spot rate changes for the three currency pairs are all above 3% per month. This is not surprising if the spot rates move with not only actual realized inflation but also

[10]The foreign exchange rate is from http://pacific.commerce.ubc.ca/xr/.

"inflation scares," which could be driven by high frequency economic news or noise but may never materialize and get captured by the CPI inflation measures.

3 Empirical Results

Our empirical results, mainly based on the Fama-French three-factor model, are presented in the first half of this section. In the second half, we provide several robustness checks by examining the impact of adding a momentum factor, a world factor, or an expected inflation factor.

3.1 Results under the Three Factor Model

In this subsection, we provide empirical evidence for the relative PPP hypothesis in the short run using our estimate of "realized inflation" measure. First, a time series regression is carried out based on equation (2):

$$R_{it} - Tbill_{it} = \alpha_i + \beta_{i1}\left(R_{Mt} - Tbill_{it}\right) + \beta_{i2}R_{SMB} + \beta_{i3}R_{HML} + \eta_{it}$$

where i stands for individual industrial portfolio returns measured in local currency and R_M, R_{SMB} and R_{HML} are country-specific factors also measured in local currency. This regression is used to extract a time series of residuals η_i for each industry, and the procedure is repeated for the U.S., UK, Japan and Germany.

Time-series regressions of the industry portfolio returns are reported in Tables III-VI. The adjusted R^2 varies materially among industries, from as low as 17% to as high as 88%. Our simulation results indicate that efficiency in extracting unanticipated inflation increases with R^2. Instead of using industry returns, one could conceivably use returns on portfolios designed to minimize unexplained time series variation and thereby improve our results. To the extent that industry returns are not necessarily optimal, our procedure here is biased against finding a confirmation of the PPP hypothesis.

Recall that $\eta_{it} \equiv (1-\beta_{i1})u_{R_{ft}} - \beta_{i1}u_{1t} - \beta_{i2}u_{2t} - \beta_{i3}u_{3t} + \epsilon_{it},$. a cross-sectional regression is carried out next by regressing the unexplained industrial excess returns (residuals) at each point of time on the estimates of the betas from the time series regressions. The estimates for $u^*_{R_{ft}}$ and $u_{R_{ft}}$ are stored for each t. We find that both series exhibit extraordinarily high sample volatility, which is probably due to the estimation error in both the time series and the cross-sectional regressions.

The estimates for R^*_{ft} and R_{ft} are obtained by adding the treasury bill rates to $u^*_{R_{ft}}$ and $u_{R_{ft}}$, and the sample mean and volatility are reported in Table II. The sample means of R_{ft} of the four countries are exactly the same as the treasury bill rates because $u_{R_{ft}}$'s have zero means by construction. The sample volatility, however, is above 7% per month, which is a very large number even in comparison to the volatility of spot rate changes. This reflects the large measurement errors in our estimates of R_{ft} and is the main reason why we put the estimated nominal rate differentials on the left hand side of the PPP regression discussed below.

For further comparison, we report summary statistics for the extracted total inflation differentials, $R^*_{ft} - R_{ft}$, together with those for the CPI inflation differentials, $\pi^*_{CPI} - \pi_{CPI}$, and those for the spot rate changes, Δs, in Table VII. The correlation coefficients between the three variables are also provided in the table.

As is well known in the literature, the volatilities of official CPI inflation differentials is *too small* as compared to those of the spot rate changes: the former is only around one-fifth to around one-tenth of the latter. On the other hand, our extracted inflation is uniformly more volatile than the exchange rate changes with the extracted inflation differential volatilities around two to three times the exchange rate change volatilities. These larger extracted inflation differential volatilities, however, generally lie in plausible range, since they should be at least as large as the exchange rate change volatilities if PPP is valid and in addition, they reflect a sizable estimation noise. As expected, our extracted inflation differentials are more volatile than the official inflation differentials, which is consistent with the influence of changing expectations absent from the official series and thus resolves the tension between the long-observed grossly larger volatility of exchange rate changes as compared to official inflation differentials.

Despite the dramatic difference in their standard deviations, the correlation between our extracted and the official inflation differentials are mostly larger than 0.1. The correlations between the official series and the spot rate changes are positive for the Germany - U.S. pair, but are negative for the Japan - U.S. and the UK - U.S. pairs. In contrast, the correlations between our extracted series and the spot rate changes are all positive and uniformly larger: it is as high as 0.5 for the quarterly Germany - U.S. data.

Finally, we use our extracted "realized inflation measures" $\hat{R}_{ft}$ and $\hat{R}^*_{ft}$ to test the relative PPP hypothesis[11]

$$R^*_{ft} - R_{ft} = a + b\Delta s + \epsilon$$

[11] For comparison, we also used the CPI inflation differential as the dependent variable in the relative PPP regression. The intercepts are all significantly different from zero while slope coefficients are all close to zero and either not statistically significant or have the wrong sign. The adjusted R^2 are close to zero or negative. The detailed results are omitted from the paper for brevity and are available upon request.

where the null hypothesis $H_0 : a = 0$ and $b = 1$, i.e., relative PPP holds, is tested against the alternative $H_a : b < 0.1$ and in particular $H_a : b = 0$, i.e., spot rate changes are too volatile to be explained by inflation differentials.[12]

Before carrying out the PPP regression, we first test for unit roots in the foreign exchange rate changes and the unexpected inflation estimates. Both the Augmented Dickey-Fuller and the Phillips-Perron tests reject the null of unit roots under various specifications.

Tables VIII reports the relative PPP regression results for three different horizons: the monthly horizon, the bi-monthly horizon using only non-overlapping observations, and the quarterly horizon using non-overlapping observations. In order to adjust for the impact of heteroscedasticity and serial correlation, only the Newey-West adjusted standard errors are reported in the Table.

At the monthly frequency, the coefficients for the contemporaneous exchange rate changes are significantly different from zero at the 5% but not significantly different from one for the Japanese Yen - US $ and the German DM and US $ pair.

At the bi-monthly and quarterly frequencies, the results are broadly consistent with the PPP hypothesis for all three pairs of currencies. The intercepts in all regressions are not significantly different from zero. The slopes are all above 0.6 and significantly different from zero but not significantly different from one at the bi-monthly frequency. At the quarterly frequency, the UK-US pair has a slope estimate significantly different from zero at 10% but not from one at the same significance level. The adjusted R^2s, though not high, are not

[12]Note that the relative PPP hypothesis is tested against a specific alternative. The presence of a foreign exchange premium, as documented in some empirical studies, is also inconsistent with relative PPP in the sense that under this alternative $b \neq 1$. However, we do not understand what value of the slope b is implied in the presence of a foreign exchange premium and thus a failure to reject the null of $b = 1$ may not shed much light.

very different from our simulation results especially for the German DM - US $ pair. This is consistent with the our observation that there are large measurement errors contained in our nominal rate estimates.

For each individual country pair, a F test is carried out to test the joint hypothesis: $a = 0$ and $b = 1$. Consistent with the individual t-tests, the null for the UK-US pair at the monthly frequency is strongly rejected with a p-value well below 1%, the null for the Japan-US pair at the monthly frequency only has a p-value of around 7% so is also rejected at the 10% significance level, but all the other seven tests have quite large p-values and fail to reject the null at a comfortable margin.

Comparing to the current empirical results in the literature from testing the short run PPP hypothesis, out point estimates are quite close to one in magnitude and generally significantly different from zero but not from one. Since the strongest evidence in favor of the relative PPP comes from regressions using bi-monthly and quarterly data with relatively fewer observations, we also carry out a bootstrap simulation and report the estimates and the bootstrapped standard errors in the same table.[13] Evident in both the point estimates and the standard errors, there is virtually no difference between the bootstrap results and the OLS estimates and the Newey-West standard errors: in the UK-US pair, the bootstrap standard errors are slightly larger and thus push the slope estimate from marginal significance to insignificance at the quarterly frequency, while the standard errors in the other two country pairs are smaller and thus strengthen our earlier results.

[13] In the bootstrap procedure, we randomly draw data from the original sample with replacement and carry out the regression using the new sample. The procedure is repeated 1000 times, and the mean and the standard deviation across these 1000 estimates are reported as the bootstrap point estimate and the bootstrap standard error in the table. Please refer to Efron and Tibshirani (1994) for an excellent description of bootstrap procedure.

We do not report the results for three other cross currency pairs. The coefficient estimates are similar to those found in the results reported for currency pairs involving US $ but they are not statistically significant. It is worth mentioning here that the efficacy of these PPP tests depends critically on how well the Fama-French three factor model describes the real stock returns and spans the real impact of inflation. The reason that real factors left over in the residuals may *bias against* the PPP hypothesis is because there is no a priori theory or economic intuition that the spot rate changes move one-to-one with real factors across different countries. To the extent that real factors may affect different economies at different times and with different significance, the spot rate changes are unlikely to move with the real factor differentials. The Fama-French three factor model seems to fit the U.S. data quite well and has been subjected to intensive study; other countries, however, have not been scrutinized as thoroughly.

3.2 Robustness Checks

3.2.1 Controlling for Momentum Factor

Some scholars have argued that the three factor model of Fama and French does not adequately capture the time series variation in stock returns and in fact a fourth real factor, Momentum, explains a significant portion of stock returns.[14] We obtained the data for the fourth Momentum factor for the U.S. from Kenneth French's web-site and repeated our analysis. For brevity, we are not reporting the time-series regressions of the industrial portfolio returns.

Table IX reports the PPP regression results when the U.S. real return has a four-factor

[14] Carhart (1997) uses a 4-factor model to evaluate mutual fund performance and argues that the 4-factor model noticeably reduces the average pricing errors relative to both the CAPM and the 3-factor model.

model. The results support the PPP hypothesis even more strongly. In particular, the slope coefficients are closer to one and have smaller standard errors, and the R^2s are higher for all frequencies.

The slope estimates for the Japan-US and UK-US pairs have improved to, respectively, 0.87 and 0.74, at the bi-monthly frequency, and as close to one as 0.95 and 0.76 at the quarterly frequency. Although the point estimate $\hat{b}$ for the Japan-US pair remains at the mid way between zero and one, the joint F-test now has a p-value of around 14% as compared to 7% before and fails to reject the null hypothesis at 10% significance level. The point estimate $\hat{b}$ for the UK-US pair remains low at only 0.33, but it is now different from zero at 10% significance level. The estimate of b is still significantly different from one in this case as indicated by both the individual t ratio and the F-statistics. On the other hand, the results remain virtually unchanged for the German DM - US $ pair for which the original results in Table VIII were already quite strong. Similar to the earlier observation, the bootstrap results are virtually the same as those of the OLS regressions.

3.2.2 Regression Using System of Equations

Although our individual country-pair results taken together are broadly in favor of the relative PPP hypothesis, one important country pair, namely the UK-US, consistently rejects the PPP, in particular at the monthly frequency. A joint test statistic across all three country pairs may help us interpret the results in a more unified way. We use a system of equations in which we stack all three country pairs in one regression, to serve this purpose.

First, an unconstrained seemingly unrelated regression (SUR) is carried out:

$$\begin{bmatrix} R^*_{ft,JP} - R_{ft,US} \\ R^*_{ft,UK} - R_{ft,US} \\ R^*_{ft,Ger} - R_{ft,US} \end{bmatrix} = \begin{bmatrix} a_1 \\ a_2 \\ a_3 \end{bmatrix} + \begin{bmatrix} \Delta s_{t,JP} & 0 & 0 \\ 0 & \Delta s_{t,UK} & 0 \\ 0 & 0 & \Delta s_{t,Ger} \end{bmatrix} \begin{bmatrix} b_1 \\ b_2 \\ b_3 \end{bmatrix} + \epsilon_t.$$

Table X reports the SUR feasible generalized least square (FGLS) regression results in Panel A for the monthly, bi-monthly and the quarterly horizon where the contemporaneous correlation between different country pairs is taken into account not only in the standard errors but also in the point estimates.[15] Panel B of the table reports the F-statistics and the corresponding p-values for joint tests across equations. There are five null hypothesis tests in the table: (1) the intercepts are zero for all equations; (2) the slope coefficients are zero for all equations; (3) the slope coefficients are the same across all equations; (4) the slope coefficients are all equal to one; and (5) the intercepts are zero and the slopes are one for all country pairs.

Comparing the results in Panel A Table X with those in Tables VIII and IX, we note that the FGLS slope coefficients, although still significant for the JP-US (at 10% level) and GM-US (at 5% level) at the monthly horizon, are uniformly smaller in magnitude and closer to zero than to one when system of equations are used. The results for the bi-monthly horizon, however, are much more encouraging. Although the FGLS point estimates for the three slopes are also smaller than those reported in Tables VIII and IX, they are mostly above 0.5

[15] Cochrane (2001) argues that the GLS procedure is more efficient only if the error covariance matrix is correctly modelled and the regression is perfectly specified. Otherwise, the GLS is less robust than OLS especially when the variance-covariance matrix has to be estimated. In particular, the FGLS can put unreasonable weight on slightly misspecified area of the model and give misleading point estimates. The OLS estimates under SUR are the same as those under individual country-pair regressions reported in Tables VII and VIII.

and statistically different from zero but not from one except for the UK-US pair under the three-factor model. The result for Germany - US pair remains the strongest support for the relative PPP where the slope estimate is around 0.97 under the three-factor and 0.94 under the four-factor model. The results for the quarterly horizon show some slight improvement over the bimonthly results and are also in favor of the relative PPP hypothesis as illustrated by the individual t-ratios for the slope estimates.

The joint F tests reported in Panel B of Table X lend strong support for the relative PPP hypothesis at the bi-monthly and quarterly horizons as well. The p-values for the hypothesis that intercepts are zero for all equations are above 0.9 at all three horizons so that the null is accepted. The null hypothesis that the slopes are all zero is rejected at 5% or better significance level. One can easily accept the null that the slopes are all equal across the three equations, and the test fails to reject the null that the slopes are all equal to one as well except for the monthly horizon. Finally, the null that the intercepts are all zero and the slopes are all one is not rejected either at the bi-monthly and quarterly horizons.

Since the hypotheses that the intercept, a_i, and the slope, b_i, are the same across equations are both accepted, these constraints are imposed in the following regression:

$$\begin{bmatrix} R^*_{ft,JP} - R_{ft,US} \\ R^*_{ft,UK} - R_{ft,US} \\ R^*_{ft,Ger} - R_{ft,US} \end{bmatrix} = a + \begin{bmatrix} \Delta s_{t,JP} \\ \Delta s_{t,UK} \\ \Delta s_{t,Ger} \end{bmatrix} b + \epsilon_t.$$

Imposition of the constraint provides a more precise estimate of the coefficients and more powerful tests of the hypothesis if the constraints are true.

The results are reported in Table XI, where Panel A contains results using the Fama-

French three-factor model while Panel B contains results where a four-factor model is applied to the U.S. data. The constrained intercept estimate, $\hat{a}$, is not significant, and the constrained slope estimate, $\hat{b}$, is significantly different from zero at all three horizons. Although $\hat{b}$ at the monthly horizon is also significantly different from one, $\hat{b}$ is quite close to one in magnitude and statistically insignificantly different from one. For example, $\hat{b}$ is above 0.9 at quarterly horizon under both the three-factor and the four-factor models. The joint F-test for H_0 : $a = 0$ and $b = 1$ is also reported. Consistent with the implications from the individual t-ratios, the null hypothesis is strongly rejected at the monthly horizon but is comfortably accepted at the bi-monthly and quarterly frequency.

Although the Newey-West standard errors adjust for heteroskedasticity and serial autocorrelation, they may still be under-estimated due to the possible contemporaneous covariation among the three country pairs. To the extent that the three country pairs are perfectly contemporaneously correlated, the standard errors are under-estimated by $\sqrt{3}$. A bootstrap simulation is carried out to obtain the contemporaneous-covariation-adjusted standard errors. For the system of equations with a total number of n observations, instead of randomly drawing n observations from the original sample with replacement, only $n_3 < n/3$, the total number of observations for the Germany-U.S. pair, observations are drawn from the original sample with replacement. The estimation is then carried out using the new sample. This procedure is repeated 2000 times, and the average and the standard deviation across these 2000 estimates are reported as the bootstrap estimate and standard errors in the table. The reported standard errors thus represent upper bounds. Note that the standard errors are substantially larger than the Newey-West ones, but the point estimates are very close to the

OLS estimates. Despite the much larger standard errors, the null hypothesis $b = 1$ is still not rejected at the bi-monthly and quarterly horizons while the null $b = 0$ is rejected at the 5% significance level for all three horizons.

3.2.3 Could this be evidence of a missing world factor?

We did not include a world factor in the asset pricing model for the reasons given in Section 2.1. This means that the factor generating model could possibly be mis-specified in the sense that the residuals η_{it} from the time-series regressions in (2) or our estimates of $u_{R_{ft}}$ contain a missing world factor. The same world factor, expressed in units of foreign currencies, could also be contained in the corresponding estimates of $u_{R^*_{ft}}$. This suggests the possibility that our PPP tests might merely be detecting a much weaker phenomenon; viz., the law of one price holds for the world factor. To check on this possibility, we performed the following procedure.

The Morgan Stanley Composite Index (MSCI),[16] is a widely-used proxy for a world factor. We regressed the residuals from the time series industry returns and our estimates of $u_{R_{ft}}$ on MSCI returns. The results (not reported here for brevity) reveal that neither the residuals from time series regressions nor the estimates of $u_{R_{ft}}$ for any country are related to the MSCI returns.

3.2.4 Are the results driven by real effects of inflation?

In our model (2), there is an implicit assumption that either the inflation does not have real effects or its real effects are spanned by the Fama-French three factors. Boudoukh, Richardson and Whitelaw (1994) argue that expected inflations may be correlated with

[16]The data was downloaded from the MSCI web-site http://www.msci.com/

future production and thus affect real stock returns. They find inflation beta patterns across industries in a regression without any other factors, but none of their beta estimates are statistically significant. We also examine their conjecture by augmenting our factor models with an additional expected inflation factor. The results are omitted from the paper for brevity. In general, we find that most of the expected inflation betas are statistically insignificant, and the PPP test results improve marginally.

3.2.5 Should the extracted inflation track the official inflation measure?

One might wonder, what relation, if any, exists between our extracted pure price inflation estimates and inflation estimates constructed from macroeconomic data. A direct comparison is difficult for five reasons. One, our extracted variables $u_{R,ft}$ and $\hat{R}_{ft}$ contain both the pure price inflation estimates and the realized real interest rates. That's why the sample means of our variables are higher than those of CPI inflation measures as reported in Table II. Two, our $\hat{R}_{ft}$'s are estimated with considerable noise, as illustrated by their large sample volatilities. Three, the CPI is constructed from goods prices sampled throughout the month, it provides a closer measure of inflation from mid-month to mid-month while the extracted inflation is the end-of-month measure, leading to some mis-matching of time period. Four, inflation measures using CPI or PPI data would include not only pure price inflation but also effects of relative price changes. This reason is one of the motivations for constructing a pure inflation measure. Finally, as we argued earlier, our extracted price inflation measure should respond to changes in expectations about future prices whereas official indices may be sticky.

Nevertheless, we should expect some association between our $\hat{R}_{ft}$ and the CPI inflation measure. Table XII reports results from regressions of our estimates of total inflation and unanticipated inflation on estimates of total inflation and unanticipated inflation respectively from the CPI and PPI data for the U.S. The results, in general, indicate a positive and significant relation between our measures and official measures. However, at monthly intervals our extracted measures are not only more volatile but also appear to be multiples of official inflation. The slope coefficients range from roughly three to eight. This might be induced by the sluggish revisions in official price indices, a supposition supported by the decline in the slope coefficients as the data interval is lengthened to quarterly, semiannual and annual. The power of the results naturally decreases as interval lengthens from a month to a year because the number of non-overlapping observations decreases considerably.

4 Concluding Remarks

Our paper makes two contributions. First, we provide a novel method for extracting estimates of *realized* pure price inflation, which involves extracting estimates of unexpected inflation, from stock returns by exploiting the fact that pure price inflation should affect nominal returns of all traded assets by exactly the same amount.

Second, we provide compelling evidence indicating that the purchasing power parity hypothesis holds quite well in the short run when we use the extracted inflation measures from stock prices. Our results complement the current consensus that PPP holds in the long run. This is in sharp contrast to the poor performance of the PPP hypothesis documented in the extensive literature surveyed in Rogoff (1996) where the inflation estimates are obtained from macroeconomic series such as the CPI.

The strong confirmation of the PPP hypothesis using our estimates suggests that perhaps the “true” price level is much more volatile than what has been historically measured using official price indices. If that is indeed true, it obviates the need for reliance on some form of exchange rate “overshooting”, as in the seminal paper by Dornbusch (1976), to explain exchange rate volatility, particularly given the fact that implications of his overshooting model do not seem to be supported by empirical evidence (Rogoff, 2002).

Table I

Simulation Results of the Effectiveness of Extracting $u_{R_{ft}}$ from Stock Returns

This table reports the simulation results of the effectiveness of extracting the unexpected risk-free rate $u_{R_{ft}}$ for different combinations of ρ and ω where ρ represents the volatility of u_{kt} as a fraction of the volatility of f_{kt} and ω represents the volatility of $u_{R_{ft}}$ as a fraction of the volatility of f_{1t}. The extracted $\hat{u}_{R_{ft}}$ is regressed on the simulated input $u_{R_{ft}}$. The procedure is simulated ten times, and the average slope coefficient estimate, its standard error and the adjusted R^2 are reported.

		U.S.					Japan				
ω	ρ:	10%	30%	50%	70%	90%	10%	30%	50%	70%	90%
10%	Slope	0.97	0.96	0.72	1.05	0.92	0.95	1.15	0.94	0.88	0.96
	S.E.	(0.12)	(0.34)	(0.56)	(0.67)	(0.46)	(0.09)	(0.29)	(0.45)	(0.60)	(0.46)
	$\bar{R}^2$	0.27	0.05	0.01	0.02	0.03	0.38	0.09	0.03	0.02	0.03
30%	Slope	0.98	1.01	0.97	0.70	0.89	0.99	0.99	1.01	0.78	0.88
	S.E.	(0.05)	(0.14)	(0.21)	(0.24)	(0.16)	(0.04)	(0.11)	(0.17)	(0.22)	(0.15)
	$\bar{R}^2$	0.66	0.24	0.11	0.05	0.16	0.76	0.33	0.17	0.07	0.17
50%	Slope	0.95	0.97	0.88	0.68	n.a.	0.97	0.98	0.96	0.72	n.a.
	S.E.	(0.06)	(0.12)	(0.17)	(0.19)	n.a.	(0.05)	(0.09)	(0.14)	(0.16)	n.a.
	$\bar{R}^2$	0.58	0.28	0.14	0.07	n.a.	0.65	0.42	0.22	0.10	n.a.
70%	Slope	1.00	0.85	0.56	0.48	n.a.	1.00	0.92	0.73	0.61	n.a.
	S.E.	(0.09)	(0.14)	(0.18)	(0.16)	n.a.	(0.08)	(0.11)	(0.16)	(0.15)	n.a.
	$\bar{R}^2$	0.40	0.17	0.06	0.06	n.a.	0.48	0.29	0.11	0.09	n.a.
90%	Slope	0.92	0.30	n.a.	n.a.	n.a.	0.93	0.43	n.a.	n.a.	n.a.
	S.E.	(0.19)	(0.20)	n.a.	n.a.	n.a.	(0.16)	(0.18)	n.a.	n.a.	n.a.
	$\bar{R}^2$	0.12	0.01	n.a.	n.a.	n.a.	0.16	0.03	n.a.	n.a.	n.a.

Table II
Data Summary Statistics

This table reports sample mean and sample volatility for the three Fama-French factor returns for the U.S., UK, Germany and Japan. The sample mean and volatility for the change in the foreign exchange rate are also reported. The sample period for the U.S. and Japan is from May 1983 to December 1999, for UK is from January 1986 to December 1999, and for Germany is from January 1988 to December 1999. The numbers are percent per month.

Variable	Mean	Standard Deviation
US Market	1.36	4.24
US SMB	-0.20	2.67
US HML	0.15	2.72
US Tbill	0.48	0.16
US CPI Inflation	0.27	0.20
US Extracted Inflation	0.48	7.15
UK Market	1.39	4.83
UK SMB	0.05	4.11
UK HML	0.26	2.55
UK Tbill	0.71	0.26
UK £- US $	-0.07	3.14
UK CPI Inflation	0.33	0.47
UK Extracted Inflation	0.71	8.22
German Market	1.34	5.15
German SMB	-0.44	4.41
German HML	0.28	3.26
German Tbill	0.44	0.17
German DM - US $	0.15	3.10
German CPI Inflation	0.19	0.26
German Extracted Inflation	0.44	7.37
Japan Market	0.75	5.92
Japan SMB	0.03	3.34
Japan HML	0.13	3.07
Japan Tbill	0.30	0.21
Japanese Yen - US $	-0.42	3.52
Japanese CPI Inflation	0.10	0.47
Japanese Extracted Inflation	0.30	8.73

Table III
Time Series Regressions of Excess Industrial Returns on the Fama-French Three Factors
(U.S.)

This table reports the regression results of excess industrial portfolio returns on the Fama-French three factors for the U.S. with sample period from January 1986 to December 1999. The Newey-West adjusted standard errors are reported in parentheses to the right of the coefficient estimates.

Industry	Constant		$R_m - Tbill$		R_{SMB}		R_{HML}		Adj. R^2
Food	-0.017	(0.27)	0.967	(0.06)	-0.444	(0.12)	0.044	(0.17)	0.633
Beer	0.103	(0.27)	0.798	(0.08)	-0.360	(0.11)	-0.268	(0.16)	0.500
Smoke	0.176	(0.54)	0.926	(0.11)	-0.195	(0.23)	0.154	(0.33)	0.271
Games	0.099	(0.31)	1.174	(0.07)	0.336	(0.17)	-0.103	(0.22)	0.699
Books	-0.216	(0.19)	1.088	(0.06)	0.087	(0.07)	0.196	(0.12)	0.757
Household	0.198	(0.14)	1.018	(0.04)	-0.389	(0.06)	-0.144	(0.07)	0.851
Apparel	-0.872	(0.33)	1.180	(0.08)	0.462	(0.12)	0.232	(0.22)	0.648
Health	0.247	(0.24)	0.902	(0.06)	-0.404	(0.09)	-0.456	(0.18)	0.710
Chems	-0.164	(0.20)	1.129	(0.06)	0.051	(0.07)	0.424	(0.17)	0.685
Textiles	-0.582	(0.36)	1.116	(0.08)	0.854	(0.12)	0.681	(0.18)	0.615
Cnstr	-0.324	(0.20)	1.192	(0.04)	0.245	(0.07)	0.307	(0.09)	0.825
Steel	-0.355	(0.30)	1.149	(0.07)	0.533	(0.11)	0.407	(0.15)	0.610
FabPr	-0.291	(0.25)	1.163	(0.04)	0.626	(0.08)	0.286	(0.10)	0.776
ElcEq	0.195	(0.34)	1.076	(0.06)	0.538	(0.12)	-0.501	(0.25)	0.761
Autos	-0.296	(0.26)	1.262	(0.05)	0.216	(0.08)	0.716	(0.13)	0.708
Carry	-0.300	(0.28)	1.147	(0.09)	0.081	(0.10)	0.339	(0.17)	0.640
Mines	-0.355	(0.43)	0.795	(0.11)	0.790	(0.16)	0.371	(0.19)	0.308
Coal	-0.712	(0.36)	0.974	(0.08)	0.634	(0.16)	0.559	(0.20)	0.455
Oil	-0.035	(0.28)	0.854	(0.07)	0.051	(0.13)	0.540	(0.12)	0.447
Util	-0.165	(0.22)	0.611	(0.06)	-0.284	(0.11)	0.580	(0.14)	0.462
Telcm	0.437	(0.29)	0.928	(0.07)	-0.156	(0.09)	-0.022	(0.13)	0.627
Servs	0.638	(0.18)	1.004	(0.05)	0.262	(0.10)	-0.882	(0.13)	0.848
BusEq	0.266	(0.31)	0.985	(0.07)	0.305	(0.14)	-0.565	(0.19)	0.696
Paper	-0.273	(0.18)	1.088	(0.06)	0.105	(0.09)	0.267	(0.16)	0.682
Trans	-0.507	(0.24)	1.162	(0.06)	0.325	(0.10)	0.496	(0.10)	0.702
Wholesale	-0.317	(0.17)	1.025	(0.05)	0.426	(0.06)	0.016	(0.11)	0.870
Retail	0.200	(0.23)	1.085	(0.07)	0.163	(0.08)	-0.168	(0.10)	0.746
Meals	-0.302	(0.24)	1.079	(0.06)	0.284	(0.12)	0.091	(0.14)	0.735
Finance	-0.232	(0.14)	1.159	(0.05)	-0.160	(0.07)	0.474	(0.07)	0.881
Other	-0.775	(0.32)	1.199	(0.06)	0.258	(0.09)	0.265	(0.21)	0.696

Table IV
Time Series Regressions of Excess Industrial Returns on the Fama-French Three Factors (U.K.)

This table reports the regression results of excess industrial portfolio returns on the Fama-French three factors for the U.K. with sample period from January 1986 to December 1999. The Newey-West adjusted standard errors are reported in parentheses to the right of the coefficient estimates.

Industry	Constant		$R_m - Tbill$		R_{SMB}		R_{HML}		Adj. R^2
Mining	0.166	(0.42)	1.183	(0.12)	0.212	(0.20)	0.187	(0.27)	0.426
OilGas	0.229	(0.28)	0.878	(0.09)	-0.028	(0.08)	0.507	(0.18)	0.518
Chemcals	-0.316	(0.32)	1.060	(0.06)	0.111	(0.09)	-0.168	(0.18)	0.631
Cnstr	-0.749	(0.28)	1.270	(0.06)	0.141	(0.08)	0.543	(0.13)	0.718
ForestryPaper	-0.024	(0.54)	1.171	(0.13)	0.392	(0.19)	0.073	(0.30)	0.303
Steel	-0.008	(0.45)	1.300	(0.10)	0.308	(0.15)	0.507	(0.21)	0.402
AerospaceDefense	-0.377	(0.38)	1.117	(0.10)	0.164	(0.16)	-0.046	(0.22)	0.526
Divs. Industrials	-0.610	(0.27)	1.076	(0.05)	0.050	(0.06)	-0.042	(0.15)	0.660
ElcEq	0.240	(0.35)	1.091	(0.13)	0.449	(0.24)	-0.497	(0.27)	0.529
EngMachinery	-0.346	(0.28)	1.223	(0.07)	0.291	(0.09)	-0.005	(0.17)	0.691
Autos	0.084	(0.38)	1.256	(0.09)	0.142	(0.11)	-0.106	(0.19)	0.568
Textiles	-1.260	(0.41)	1.202	(0.09)	0.284	(0.14)	0.222	(0.22)	0.640
Beverages	-0.128	(0.25)	0.928	(0.07)	-0.226	(0.11)	-0.242	(0.14)	0.685
Food	-0.183	(0.23)	0.810	(0.07)	-0.047	(0.11)	-0.092	(0.17)	0.651
Health	-0.294	(0.21)	0.943	(0.05)	0.183	(0.06)	-0.348	(0.16)	0.670
Pack	-0.347	(0.34)	1.128	(0.07)	0.178	(0.12)	-0.070	(0.17)	0.558
Prsnl.Care	-0.323	(0.36)	0.826	(0.08)	-0.203	(0.10)	0.014	(0.19)	0.389
PharmBiotech	0.782	(0.30)	0.840	(0.08)	-0.097	(0.09)	-0.901	(0.17)	0.573
Tobacco	0.346	(0.47)	0.776	(0.16)	-0.308	(0.17)	0.214	(0.25)	0.326
Dist.	-0.409	(0.28)	1.182	(0.06)	0.307	(0.08)	0.015	(0.11)	0.693
GenRetail	-0.516	(0.24)	0.904	(0.05)	-0.174	(0.08)	0.209	(0.14)	0.664
Ent.Hotels	-0.312	(0.28)	1.178	(0.08)	0.075	(0.08)	0.316	(0.11)	0.743
Media	0.064	(0.22)	1.272	(0.06)	0.371	(0.10)	-0.184	(0.10)	0.758
ResPub	-0.144	(0.26)	0.819	(0.08)	-0.154	(0.06)	0.269	(0.11)	0.620
Support	-0.096	(0.20)	1.085	(0.05)	0.375	(0.07)	-0.138	(0.08)	0.725
Transport	-0.261	(0.25)	1.023	(0.05)	0.058	(0.07)	0.200	(0.15)	0.732
FoodDrugRetailers	-0.225	(0.33)	0.684	(0.06)	-0.126	(0.09)	0.186	(0.18)	0.385
Telecom	0.605	(0.38)	0.856	(0.09)	-0.134	(0.11)	-0.200	(0.17)	0.498
Banks	0.440	(0.25)	1.117	(0.08)	-0.346	(0.08)	0.436	(0.19)	0.720
Insurance	-0.387	(0.25)	1.073	(0.07)	-0.166	(0.08)	0.181	(0.11)	0.679
LifeAssurance	0.479	(0.28)	0.925	(0.06)	-0.122	(0.07)	0.067	(0.14)	0.557
InvFirm	-0.100	(0.15)	1.066	(0.05)	0.165	(0.06)	0.031	(0.07)	0.870
RealEstate	-0.761	(0.23)	1.003	(0.08)	0.109	(0.07)	0.943	(0.13)	0.718
OtherFin	-0.054	(0.25)	1.298	(0.07)	0.297	(0.07)	0.106	(0.18)	0.748
IT Hardware	1.367	(1.14)	1.412	(0.27)	0.845	(0.49)	-0.904	(0.82)	0.253
Software	0.649	(0.47)	1.136	(0.11)	0.857	(0.22)	-0.461	(0.24)	0.505

Table V

Time Series Regressions of Excess Industrial Returns on the Fama-French Three Factors (Germany)

This table reports the regression results of excess industrial portfolio returns on the Fama-French three factors for Germany with sample period from January 1988 to December 1999. The Newey-West adjusted standard errors are reported in parentheses to the right of the coefficient estimates.

Industry	Constant		$R_m - Tbill$		R_{SMB}		R_{HML}		Adj. R^2
Autos	-0.343	(0.34)	1.118	(0.09)	-0.287	(0.14)	0.179	(0.19)	0.736
Banks	-0.220	(0.23)	0.911	(0.07)	-0.255	(0.08)	0.378	(0.12)	0.782
Chemcals	-0.058	(0.25)	0.825	(0.08)	-0.239	(0.09)	0.308	(0.10)	0.698
Media	1.126	(0.67)	0.853	(0.14)	0.454	(0.16)	-0.656	(0.26)	0.229
BSC Resources	0.082	(0.27)	1.120	(0.08)	0.463	(0.09)	0.223	(0.12)	0.663
FoodBevrge	-0.205	(0.17)	0.851	(0.06)	0.612	(0.07)	-0.034	(0.07)	0.618
Technology	-0.010	(0.31)	1.081	(0.07)	-0.090	(0.09)	0.018	(0.11)	0.698
Insurance	0.075	(0.29)	1.010	(0.08)	-0.167	(0.11)	-0.234	(0.17)	0.712
TRSPT&LGISTC	-0.048	(0.43)	1.190	(0.13)	0.280	(0.17)	0.305	(0.13)	0.463
Machinery	-0.554	(0.25)	1.244	(0.06)	0.481	(0.08)	0.215	(0.09)	0.753
Industrial	1.252	(0.43)	1.071	(0.11)	-0.010	(0.11)	-0.251	(0.24)	0.584
Construction	-0.483	(0.39)	1.475	(0.10)	1.003	(0.17)	0.036	(0.11)	0.586
PharmHlth	0.440	(0.26)	0.878	(0.06)	0.274	(0.11)	-0.261	(0.12)	0.555
Retail	-0.376	(0.37)	1.247	(0.13)	0.593	(0.10)	-0.185	(0.14)	0.546
Software	2.578	(0.83)	1.434	(0.15)	0.467	(0.22)	-0.087	(0.25)	0.308
Telecom	0.284	(0.67)	0.738	(0.13)	0.076	(0.18)	-0.397	(0.27)	0.170
Utilities	0.063	(0.23)	0.689	(0.07)	-0.009	(0.09)	0.222	(0.08)	0.594
Financial SRV	0.552	(0.33)	0.935	(0.13)	0.518	(0.14)	-0.123	(0.12)	0.408
Consumer Cyclical	-0.541	(0.40)	1.188	(0.08)	0.604	(0.11)	0.082	(0.14)	0.545

Table VI
Time Series Regressions of Excess Industrial Returns on the Fama-French Three Factors (Japan)

This table reports the regression results of excess industrial portfolio returns on the Fama-French three factors for Japan with sample period from January 1988 to December 1999. The Newey-West adjusted standard errors are reported in parentheses to the right of the coefficient estimates.

Industry	Constant		$R_m - Tbill$		R_{SMB}		R_{HML}		Adj. R^2
AirTransport	0.194	(0.57)	0.909	(0.11)	0.151	(0.12)	-0.049	(0.26)	0.348
Banks	0.315	(0.62)	0.964	(0.09)	-0.133	(0.11)	0.117	(0.15)	0.388
Chemical	0.010	(0.18)	1.076	(0.04)	0.164	(0.06)	-0.097	(0.09)	0.853
Communication	0.998	(0.63)	1.158	(0.10)	-0.105	(0.10)	-0.658	(0.19)	0.481
Construction	-0.504	(0.35)	1.065	(0.08)	0.241	(0.11)	0.344	(0.23)	0.605
ElcEq	0.007	(0.33)	0.950	(0.07)	-0.171	(0.09)	0.047	(0.34)	0.555
Utilities	-0.040	(0.35)	0.913	(0.09)	-0.539	(0.11)	0.698	(0.37)	0.467
Fisheries	-0.526	(0.30)	0.999	(0.07)	0.497	(0.10)	0.193	(0.17)	0.628
Foods	-0.008	(0.27)	0.894	(0.05)	0.218	(0.07)	-0.011	(0.10)	0.728
GlassCeramics	-0.039	(0.23)	1.057	(0.05)	0.105	(0.05)	-0.179	(0.09)	0.797
Insurance	0.337	(0.43)	1.107	(0.09)	-0.392	(0.10)	0.219	(0.19)	0.559
Steel	-0.382	(0.50)	1.131	(0.06)	-0.085	(0.11)	0.271	(0.17)	0.581
LandTransport	0.306	(0.35)	1.017	(0.08)	-0.013	(0.09)	0.179	(0.35)	0.522
Machinery	-0.202	(0.19)	1.071	(0.04)	0.358	(0.06)	-0.011	(0.11)	0.859
MarineTransport	-0.566	(0.50)	1.293	(0.09)	0.314	(0.14)	0.164	(0.26)	0.584
MetalProducts	-0.278	(0.24)	0.947	(0.06)	0.582	(0.08)	0.312	(0.12)	0.711
Mining	-0.241	(0.44)	1.093	(0.06)	0.605	(0.08)	-0.381	(0.13)	0.573
Non-ferrous Mets	-0.021	(0.22)	1.169	(0.05)	0.080	(0.06)	-0.185	(0.12)	0.774
OilCoal	-0.409	(0.41)	1.021	(0.09)	0.200	(0.11)	-0.009	(0.13)	0.533
OtherFinancials	-0.131	(0.40)	1.037	(0.06)	0.208	(0.10)	0.153	(0.12)	0.635
OtherProducts	0.251	(0.28)	0.832	(0.05)	0.140	(0.10)	0.016	(0.16)	0.649
Pharmaceutical	0.282	(0.32)	0.800	(0.06)	0.003	(0.07)	-0.136	(0.12)	0.531
PrecisionInstr.	0.174	(0.36)	0.922	(0.08)	0.107	(0.08)	-0.247	(0.30)	0.558
Paper	-0.287	(0.30)	0.824	(0.10)	0.235	(0.09)	0.171	(0.12)	0.524
RealEstate	0.081	(0.33)	1.267	(0.13)	-0.307	(0.16)	0.355	(0.36)	0.536
Retail	0.039	(0.24)	0.908	(0.05)	0.224	(0.07)	0.249	(0.10)	0.702
Rubber	0.437	(0.30)	1.033	(0.05)	0.133	(0.11)	0.068	(0.12)	0.675
Securities	-0.002	(0.52)	1.506	(0.17)	-0.391	(0.15)	0.347	(0.20)	0.621
Service	0.174	(0.34)	0.922	(0.06)	0.361	(0.09)	0.167	(0.13)	0.679
Textiles	-0.289	(0.22)	0.999	(0.04)	0.236	(0.08)	0.071	(0.13)	0.769
Transport Equip.	0.200	(0.21)	0.910	(0.04)	-0.302	(0.08)	0.183	(0.16)	0.716
Warehouse	-0.206	(0.31)	1.113	(0.08)	0.326	(0.11)	0.425	(0.27)	0.625
Wholesale	-0.359	(0.22)	1.079	(0.05)	0.095	(0.07)	0.133	(0.10)	0.830

Table VII

Summary Statistics for Extracted Inflation Differentials, CPI Inflation Differentials, and Foreign Exchange Rate Changes

This table reports the mean and standard deviation of extracted inflation differentials from the Fama-French Three Factor Model, official (CPI) inflation differentials, and foreign exchange rate changes. The inflation differential is the difference between a foreign and the U.S. extracted inflation rates. The foreign exchange rate is measured as foreign currency per U.S. dollar. The CPI inflation is calculated from the CPI index. The correlations between the variables are also reported.

		CPI Inflation Differentials				Extracted Inflation Differentials			Spot Rate Changes	
Currency Pair	frequency	mean	standard deviation	correlation with extracted counterpart	correlation with spot rate changes	mean	standard deviation	correlation with spot rate changes	mean	standard deviation
Japanese Yen-US $	monthly	-0.167	0.499	0.100	-0.065	-0.186	10.437	0.176	-0.424	3.520
	bi-monthly	-0.334	0.822	0.139	-0.170	-0.371	15.199	0.269	-0.847	5.364
	quarterly	-0.500	0.681	0.268	-0.090	-0.697	20.925	0.275	-1.290	6.775
UK £- US $	monthly	0.074	0.482	0.100	-0.029	0.275	8.874	0.100	-0.066	3.142
	bi-monthly	0.149	0.701	0.116	-0.052	0.550	12.813	0.263	-0.133	4.983
	quarterly	0.223	0.797	0.322	-0.075	0.825	17.084	0.201	-0.199	5.454
German DM- US $	monthly	-0.073	0.296	0.160	0.127	0.005	8.584	0.234	0.148	3.095
	bi-monthly	-0.146	0.430	0.064	0.170	0.011	11.720	0.419	0.296	4.648
	quarterly	-0.218	0.534	0.189	0.210	0.016	15.688	0.501	0.443	6.031

Table VIII
PPP Regression Results using Extracted Inflation from the Fama-French Three Factor Model

This table reports the results of testing the relative PPP hypothesis by regressing the estimated inflation differential on the change in the nominal exchange rate, $R^*_{ft} - R_{ft} = a + b\Delta s_t + \epsilon_t$. The change in the foreign exchange rate is measured as yen or British pounds or Deutche Marks per U.S. dollar. Newey-West adjusted standard errors are reported in parentheses. n stands for the number of observations. The bootstrap point estimates and the standard errors are reported as well. The p value refers to the p-value of the joint F-test for the null hypothesis: intercept=0 and slope=1.

			OLS Results			Bootstrap Results		$H_0: a=0$ and $b=1$	
Currency Pair	frequency	n	a	b	Adj. R^2	a	b	$F(2, n-2)$	p-value
Japanese Yen-US $	monthly	200	0.036	0.522	0.026	0.028	0.535	2.706	0.069
			(0.639)	(0.250)*		(0.700)	(0.220)*		
	bi-monthly	100	0.276	0.763	0.063	0.281	0.769	0.421	0.658
			(1.021)	(0.337)*		(1.473)	(0.294)*		
	quarterly	66	0.398	0.849	0.061	0.411	0.882	0.111	0.895
			(1.404)	(0.395)*		(2.343)	(0.382)*		
UK £- US $	monthly	168	0.294	0.283	0.004	0.292	0.275	5.534	0.005
			(0.732)	(0.189)		(0.669)	(0.219)		
	bi-monthly	84	0.640	0.677	0.058	0.632	0.672	0.494	0.613
			(1.436)	(0.150)*		(1.372)	(0.220)*		
	quarterly	56	0.950	0.631	0.023	0.892	0.578	0.820	0.444
			(2.254)	(0.352)**		(2.186)	(0.412)		
German DM- US $	monthly	144	-0.091	0.648	0.048	-0.087	0.651	1.230	0.295
			(0.667)	(0.227)*		(0.686)	(0.212)*		
	bi-monthly	72	-0.302	1.056	0.164	-0.318	1.073	0.047	0.954
			(1.209)	(0.276)*		(1.222)	(0.298)*		
	quarterly	48	-0.562	1.303	0.235	-0.631	1.302	0.440	0.647
			(1.587)	(0.380)*		(1.979)	(0.320)*		

*: significantly different from zero at 5%.

**: significantly different from zero at 10%.

Table IX

PPP Regression Results using Extracted Inflation Using Momentum as an Additional Factor

This table reports the results of testing the relative PPP hypothesis by regressing the estimated inflation differential on the change in the nominal exchange rate, $R^*_{ft} - R_{ft} = a + b\Delta s_t + \epsilon_t$. The change in the foreign exchange rate is measured as yen or British pounds or Deutche Marks per U.S. dollar. Newey-West adjusted standard errors are reported in parentheses. n stands for the number of observations. The bootstrap point estimates and the standard errors are reported as well. The p value refers to the p-value of the joint F-test for the null hypothesis: intercept=0 and slope=1.

			OLS Results			Bootstrap Results		$H_0 : a = 0$ and $b = 1$	
Currency Pair	frequency	n	a	b	Adj. R^2	a	b	$F(2, n-2)$	p-value
Japanese Yen-US $	monthly	200	0.059	0.577	0.031	0.036	0.589	2.005	0.137
			(0.678)	(0.259)*		(0.740)	(0.228)*		
	bi-monthly	100	0.364	0.867	0.074	0.356	0.881	0.152	0.859
			(1.146)	(0.355)*		(1.489)	(0.308)*		
	quarterly	66	0.425	0.903	0.062	0.431	0.919	0.052	0.950
			(1.577)	(0.413)*		(2.450)	(0.425)*		
UK £- US $	monthly	168	0.296	0.326	0.007	0.305	0.326	4.791	0.010
			(0.744)	(0.196)**		(0.689)	(0.225)		
	bi-monthly	84	0.648	0.742	0.067	0.581	0.752	0.541	0.584
			(1.458)	(0.146)*		(1.394)	(0.212)*		
	quarterly	56	0.975	0.756	0.036	0.870	0.710	0.256	0.775
			(2.298)	(0.351)*		(2.323)	(0.430)**		
German DM- US $	monthly	144	-0.091	0.650	0.048	-0.077	0.654	1.214	0.300
			(0.672)	(0.227)*		(0.690)	(0.212)*		
	bi-monthly	72	-0.302	1.055	0.163	-0.309	1.056	0.046	0.955
			(1.223)	(0.277)*		(1.216)	(0.312)*		
	quarterly	48	-0.561	1.301	0.231	-0.565	1.311	0.428	0.650
			(1.587)	(0.380)*		(1.979)	(0.320)*		

*: significantly different from zero at 5%.

**: significantly different from zero at 10%.

Table X
Seemingly Unrelated PPP Regressions for Three Country Pairs

This table reports the results of testing the relative PPP hypothesis by using the seemingly unrelated system equations (SUR) across all three country-pairs. The estimated inflation differential is regressed on the change in the nominal exchange rate, $\begin{bmatrix} R^*_{ft,JP} - R_{ft,US} \\ R^*_{ft,UK} - R_{ft,US} \\ R^*_{ft,Ger} - R_{ft,US} \end{bmatrix} = \begin{bmatrix} a_1 \\ a_2 \\ a_3 \end{bmatrix} + \begin{bmatrix} \Delta s_{t,JP} & 0 & 0 \\ 0 & \Delta s_{t,UK} & 0 \\ 0 & 0 & \Delta s_{t,Ger} \end{bmatrix} \begin{bmatrix} b_1 \\ b_2 \\ b_3 \end{bmatrix} + \epsilon_t$. The change in the foreign exchange rate is measured as yen or British pounds or Deutsche Marks per U.S. dollar. Newey-West adjusted standard errors are reported in parentheses, and n stands for the number of observations for the system.

	Panel A: SUR FGLS Estimates								
			Three Factor Model			Momentum as a 4th Factor		
Currency Pair	frequency	n	a	b	System R^2	a	b	System R^2
Japanese Yen-US $	Monthly	512	-0.032	0.362	0.016	-0.020	0.391	0.017
			(0.730)	(0.192)**		(0.753)	(0.196)*	
UK £- US $			0.349	0.152		0.359	0.174	
			(0.675)	(0.203)		(0.682)	(0.203)	
German DM- US $			-0.067	0.530		-0.086	0.517	
			(0.692)	(0.219)*		(0.692)	(0.219)*	
Japanese Yen-US $	bi-monthly	256	0.094	0.542	0.067	0.124	0.584	0.065
			(1.477)	(0.252)*		(1.556)	(0.258)*	
UK £- US $			0.749	0.491		0.775	0.501	
			(1.341)	(0.252)**		(1.371)	(0.251)**	
German DM- US $			-0.241	0.968		-0.251	0.943	
			(1.262)	(0.271)*		(1.264)	(0.270)*	
Japanese Yen-US $	Quarterly	170	0.295	0.661	0.085	0.327	0.686	0.077
			(2.503)	(0.341)**		(2.641)	(0.350)*	
UK £- US $			1.051	0.450		1.074	0.497	
			(2.236)	(0.391)		(2.309)	(0.395)	
German DM- US $			-0.433	1.147		-0.462	1.083	
			(1.978)	(0.326)*		(1.986)	(0.323)*	

*: significantly different from zero at 5%.
**: significantly different from zero at 10%.

Table X

Seemingly Unrelated PPP Regressions for Three Country Pairs (continued)

	Panel B: Hypothesis Testing						
				Three Factor Model		Momentum as a 4th Factor	
H_0	Frequency	Df_1	Df_2	$F(Df_1, Df_2)$	p-value	$F(Df_1, Df_2)$	p-value
$a_1 = a_2 = a_3 = 0$	monthly	3	506	0.119	0.950	0.126	0.946
	bi-monthly	3	250	0.128	0.945	0.136	0.940
	quarterly	3	164	0.094	0.964	0.101	0.960
$b_1 = b_2 = b_3 = 0$	monthly	3	506	2.865	0.036	2.916	0.034
	bi-monthly	3	250	6.247	< 0.001	6.129	< 0.001
	quarterly	3	164	5.220	0.002	4.779	0.003
$b_1 = b_2 = b_3$	monthly	2	506	0.911	0.406	0.785	0.462
	bi-monthly	2	250	0.981	0.379	0.788	0.460
	quarterly	2	164	1.032	0.362	0.742	0.482
$b_1 = b_2 = b_3 = 1$	monthly	3	506	8.865	< 0.001	8.296	< 0.001
	bi-monthly	3	250	2.109	0.102	1.817	0.150
	quarterly	3	164	0.943	0.426	0.730	0.541
$a_1 = a_2 = a_3 = 0$ and $b_1 = b_2 = b_3 = 1$	monthly	6	506	4.507	< 0.001	4.213	< 0.001
	bi-monthly	6	250	1.126	0.352	0.982	0.448
	quarterly	6	164	0.527	0.790	0.423	0.867

Table XI
Constrained System of Equations PPP Regression for Three Country Pairs

This table reports the results of testing the relative PPP hypothesis of the constrained system equations across all three country-pairs. The estimated inflation differential is regressed on the change in the nominal exchange rate, $\begin{bmatrix} R^*_{ft,JP} - R_{ft,US} \\ R^*_{ft,UK} - R_{ft,US} \\ R^*_{ft,Ger} - R_{ft,US} \end{bmatrix} = a + \begin{bmatrix} \Delta s_{t,JP} \\ \Delta s_{t,UK} \\ \Delta s_{t,Ger} \end{bmatrix} b + \epsilon_t$. The change in the foreign exchange rate is measured as yen or British pounds or Deutsche Marks per U.S. dollar. Newey-West adjusted standard errors are reported in parentheses. The bootstrap point estimates and the standard errors are reported as well, and n stands for the number of observations for the system. For the bootstrap procedure, to guard against the possibility of underestimation of standard errors because of potential cross correlation among the three country pairs, only $n_3 < n/3$, the total number of observations for the Germany-U.S. pair, observations are drawn from the original sample with replacement; thus reported standard errors represent upper bounds. The p value refers to the p-value of the joint F-test for the null hypothesis: intercept=0 and slope=1.

Currency Pair	frequency	n	OLS Results a	OLS Results b	OLS Results Adj. R^2	Bootstrap Results a	Bootstrap Results b	$H_0: a=0$ and $b=1$ $F(2, n-2)$	$H_0: a=0$ and $b=1$ p-value
Panel A: Three Factor Model									
System of equations	monthly	512	0.089	0.482	0.026	0.034	0.497	8.644	0.0002
			(0.389)	(0.141)*		(0.769)	(0.246)*		
	bi-monthly	256	0.272	0.803	0.087	0.261	0.820	0.847	0.430
			(0.753)	(0.175)*		(1.499)	(0.321)*		
	quarterly	170	0.546	0.903	0.109	0.531	0.907	0.244	0.784
			(1.139)	(0.222)*		(2.291)	(0.396)*		
Panel B: Four Factor Model									
System of Equations	monthly	512	0.095	0.520	0.030	0.114	0.525	7.171	0.0008
			(0.400)	(0.144)*		(0.805)	(0.254)*		
	bi-monthly	256	0.291	0.869	0.095	0.276	0.878	0.394	0.675
			(0.781)	(0.182)*		(1.551)	(0.331)*		
	quarterly	170	0.527	0.988	0.121	0.557	0.981	0.094	0.910
			(1.189)	(0.226)*		(2.323)	(0.402)*		

*: significantly different from zero at 5%.
**: significantly different from zero at 10%.

Table XII
Regression of Extracted Inflation on Official Inflation Measures (U.S.)

This table presents the results of regressing the extracted total and unexpected inflation on the official inflation measures. The extracted unexpected inflation is $\hat{u}_{R_{ft}}$ and the extracted total inflation is $\hat{R}_{ft} = TBill_{t-1} + \hat{u}_{R_{ft}}$. We use both CPI and PPI inflation rates as the official total inflation measures. The official unexpected inflation measure is simply the CPI or PPI inflation rate minus the Tbill rate. The monthly, the non-overlapping quarterly, semi-annual and annual frequencies of the data are used in the regressions. Newey-West adjusted standard errors (S.E.) are reported in parentheses.

Price Index	Dependent Variable	Frequency	Constant	S.E.	Official Infl. Rate	S.E.	Adj. $\bar{R}^2$
PPI	$\hat{R}_{ft}$	monthly	0.035	(0.445)	3.087	(0.756)*	0.059
		quarterly	-0.143	(1.057)	3.713	(0.833)*	0.137
		semiannual	0.781	(1.972)	2.357	(1.472)	0.040
		annual	3.873	(3.963)	0.888	(0.904)	-0.060
CPI	$\hat{R}_{ft}$	monthly	-1.638	(0.625)*	8.061	(1.804)*	0.063
		quarterly	-4.808	(2.201)*	7.926	(2.796)*	0.075
		semiannual	-5.980	(4.568)	5.573	(2.534)*	0.024
		annual	-7.694	(6.828)	4.193	(2.004)*	0.007
PPI	$\hat{u}_{R_{ft}}$	monthly	0.824	(0.523)	2.674	(0.780)*	0.044
		quarterly	2.846	(1.590)	3.080	(0.908)*	0.093
		semiannual	2.369	(3.905)	1.282	(1.433)	-0.013
		annual	-2.330	(6.645)	-0.630	(0.912)	-0.070
CPI	$\hat{u}_{R_{ft}}$	monthly	0.992	(0.652)	5.464	(2.031)*	0.027
		quarterly	1.896	(2.563)	3.483	(3.422)	0.002
		semiannual	-1.509	(5.261)	-1.386	(4.093)	-0.034
		annual	-11.368	(7.926)	-5.220	(2.790)	0.058

*: significantly different from zero at 5%.

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