NBER WORKING PAPER SERIES

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Working Paper 7408
http://www.nber.org/papers/w7408

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 1999

We thank Mike Dooley, Daniel Friedman, K. C. Fung, Linda Goldberg, Ken Kletzer, Manuel Pastor, Jr., Donald Wittman, the participants of the Purchasing Power Parity session of the 1999 AEA Annual Meetings, and the Economics Brown Bag Seminar at UCSC for their comments on earlier drafts. This research is supported by funds from the UC Pacific Rim Research Program. The views expressed herein are those of the authors and not necessarily those of the National Bureau of Economic Research.

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NBER Working Paper No. 7408
October 1999
JEL No. F31, F40, L16

ABSTRACT

We examine the relationship between market structure and the persistence of U.S. dollar-based sectoral real exchange rates for fourteen OECD countries. Our empirical results based on disaggregated data suggest that differences in market structure significantly determine the rates at which deviations from sectoral purchasing power parity decay. Specifically, industries with a larger price-cost margin are found to exhibit slower parity reversion of their sectoral real exchange rates. Further, as the degree of intra-industry trade activity increases, sectoral real exchange rate persistence becomes more pronounced. These findings imply that an imperfectly competitive market structure contributes to the well-documented persistence in real exchange rates.

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1 Introduction

The last few years have borne witness to a remarkable transformation in the profession's views on the purchasing power parity (PPP) phenomenon. In contrast to the view prevailing in the 1980s, there now appears to be a consensus that long run PPP holds. Nevertheless, the slow *rate* of parity reversion remains a puzzle (Rogoff, 1996). Hence, we believe that the research agenda should no longer be directed solely toward detecting real exchange rate stationarity, but rather move toward isolating the empirical determinants of the rate of reversion from a microeconomic, market structure, perspective. This paper represents an initial effort in this direction.

The evidence for real exchange rate stationarity comes from several sources. A set of studies appeals to long-spans of data which encompass several exchange rate regimes.[1] For post-Bretton Woods data, evidence regarding PPP is usually derived from panel-based unit root tests (Levin and Lin, 1992). By pooling observations across different countries, panel data unit root tests attain a better power to uncover PPP behavior (Wei and Parsley,1995; Frankel and Rose,1996; Oh, 1996; Wu, 1996; Engel, Hendrickson and Rogers, 1997).[2]

One intriguing empirical regularity is the extremely slow rate at which PPP deviations decay (Rogoff, 1996). Figure 1 displays the half-lives of PPP deviations reported in some recent studies. Most of these studies use the autoregressive (AR) coefficient as a sufficient statistic to characterize the time profile of the effects of a shock to PPP. The oft-cited 3.5 to 5.5 year half-life corresponds to the earlier panel studies (Frankel and Rose, 1996; Wei and Parsley, 1995). More recent panel studies (Wu, 1996; Papell, 1997) find somewhat more rapid reversion, with half-lives on the order of 2 to 2.5 years. Even these estimates appear to imply more sluggishness

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[1] Examples include Abuaf and Jorion (1990), Culver and Papell (1995), Diebold, Husted, and Rush (1991), Glen (1992), and Lothian and Taylor (1996).

[2] See Taylor and Sarno (1998) for the limitation of panel data unit root tests. Non-panel studies using the post-Bretton Woods data are typically less favorable to long-run PPP (Meese and Rogoff, 1988; Mark, 1990). An exception is Cheung and Lai (1998). Also, see Engel (1999) for an alternative view.

than one can attribute entirely to nominal rigidities alone. What then accounts for such slow parity reversion? Despite the plethora of PPP literature in the past several years, studies attempting to answer this question are rather scarce.

In examining the relationship between PPP deviations and trade volume deviations among G7 countries, Campa and Wolf (1997) find that greater geographical proximity and a larger market size accelerate the rate of PPP reversion. Surprisingly, they find that a greater bilateral trade share leads to *slower* reversion, which seems to contradict the goods arbitrage-based view of long-run PPP. Adopting the macroeconomic perspective, Cheung and Lai (1999) examine variables such as inflation, productivity growth, trade openness, and government expenditure to account for the differences in real exchange rate persistence across 94 countries. Although lower inflation and larger government spending are found to be associated with slower parity reversion, a substantial portion of the cross-country differences in the real exchange rate persistence remains unexplained.

One potential source of the real exchange rate persistence that is not considered in the preceding studies is the discriminatory pricing behavior of firms with market power, termed pricing to market (PTM) (Krugman, 1987). When markets are segmented, a monopolistically competitive firm's optimal pricing behavior can create a wedge between common currency prices of the same good destined to different markets, and consequently, violate the law of one price (LOP) which is a building block of PPP. Empirical evidence of PTM includes Giovannini (1988), Knetter (1989, 1993), Marston (1990), and Ohno (1989) among others.[3] Implications of PTM for PPP deviations can be quite substantial. For instance, in examining the post-Bretton Woods real exchange rates between the U.S. dollar and Canadian dollar, German mark, British pound, and Japanese yen, Feenstra and Kendall (1997) find that a significant portion of the observed PPP deviations is attributable to the incomplete exchange rate pass-through due to PTM. Further, using a dynamic general equilibrium model, Faruqee (1995) shows that PTM

[3] Goldberg and Knetter (1997) provide an excellent literature survey of PTM and the closely related subject of exchange rate pass-through.

behavior intensifies the degree of persistence in the real exchange rate under nominal rigidities. The findings of those studies suggest that discriminatory pricing behavior may explain, at least in part, the commonly observed excessive persistence in real exchange rates.

Since discriminatory pricing behavior requires an imperfectly competitive market structure under which firms behave as price setters, it is quite conceivable that differences in market structure across industries and/or countries play an important role in determining the persistence of PPP deviations. This is the theme of the current study. We use data on nine manufacturing sectors of the U.S. and fourteen other OECD countries, and empirically test if differences in sectoral real exchange rate persistence systematically arise from differences in market structure. Specifically, we consider the hypothesis that industries with less competitive market structure have more persistent sectoral real exchange rates.

The remainder of this manuscript is organized as follows. Section 2 discusses the linkage between market structure and real exchange rate persistence. Section 3 describes the data, estimates the mean reversion coefficients of the sectoral real exchange rates, and constructs two proxies for market structure - the price-cost margin and the intra-industry trade index. In section 4 we analyze the effects of market structure on the persistence of sectoral PPP deviations. Section 5 presents additional analyses based on alternative measures of market structure to check the robustness of the results obtained in section 4. Some concluding remarks are provided in section 6.

2 Market Structure and Real Exchange Rate Persistence

In an early paper relating market structure to PPP, Dornbusch (1987) examines the adjustment of relative prices to exchange rate movements. His analysis suggests that the response of relative prices critically depends on the following three factors: market integration or separation; substitution between domestic and foreign variants of a product; and market structure (or market organization in Dornbusch's nomenclature). When markets are segmented and the price elasticities of demand are not constant, a monopolistic firm's optimal pricing behavior in response to exchange rate changes leads to price discrimination by market destinations.[4] Such pricing behavior was described as PTM by Krugman (1987). In a recent study, Feenstra and Kendall (1997) find PTM contributing substantially to the post-Bretton Woods PPP deviations among G5 countries. For instance, for the dollar/yen and dollar/sterling real rates, their estimates suggest that almost one third of the total PPP deviations are attributable to PTM.

Although the finding of Feenstra and Kendall (1997) suggests that market structure is important to PPP deviations, its implications for the *persistence* of PPP deviations are not clear. Faruqee (1995) provides some insight on the linkage between market structure and real exchange rate persistence utilizing a dynamic general equilibrium model under monopolistic competition and market segmentation.[5] Consider the real exchange rate dynamics implied by the model:

$$q_t = \varphi q_{t-1} + \omega_1(m_t - m_t^*) + \omega_2(m_{t-1} - m_{t-1}^*) \tag{1}$$

where

$$\varphi = \frac{1-\sqrt{\Pi+2\theta(1-\psi)}}{1+\sqrt{\Pi+2\theta(1-\psi)}} \quad ; \quad \Pi = \frac{\gamma-1}{\varepsilon(\gamma-1)+1} \quad ; \quad \theta = 1-\varepsilon\Pi \tag{2}$$

and q_t is the log of real exchange rate, m_t (m^*_t) is the domestic (foreign) money supply, $\varepsilon > 1$ is the constant elasticity of substitution between any two varieties from the same industry, $(\gamma-1) > 0$

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[4] As Dornbusch (1987) points out, if demand curves have constant price elasticities in both foreign and domestic markets, a monopolistically competitive firm will follow a constant markup pricing rule, and the relative price of its product will remain constant as the exchange rate fluctuates even if markets are effectively segmented. On the other hand, any demand curve less convex than a constant elasticity curve will result in PTM. See also Marston (1990) for a more detailed discussion and a comparative static analysis.

[5] The model assumes two countries resided by representative consumer/producer agents engaging in either inter-sectoral or intra-sectoral trade in the presence of menu costs and staggered price adjustment. With inter-sectoral trade, countries specialize and trade at the industry level, while they trade at the variety level under intra-industry trade.

measures marginal disutility with respect to output, and $.5 < \psi < 1$ is the expenditure share for home goods.[6] A key implication of this model is that the AR coefficient φ increases as the elasticity of substitution between varieties from the same industry (ε) rises. The intuition is that as the elasticity of substitution rises exporting firms become more concerned with maintaining their prices in line with domestic competitors, leading to increased price rigidities in local currency terms. On the other hand, the real exchange rate persistence is reduced as the expenditure share on imported goods ($1-\psi$) increases since it makes the domestic price level more susceptible to inflation induced by exchange rate depreciation, and hence, encourages more frequent price adjustment.

The above studies highlight several industry-specific factors that may significantly determine the degree of real exchange rate persistence. These factors are the imperfectly competitive market structure, market segmentation, substitutability between domestic and foreign variants within an industry, and exposure to international trade. The objective of the current study is to empirically document the effects of these factors and the determinants of sectoral real exchange rate persistence. In the subsequent sections, we first construct empirical measures of sectoral real exchange rate persistence and the industry-specific market structure. Then, we test if differences in market structure across industries are indeed systematically related to differences in sectoral real exchange rate persistence.

3 Data and Preliminary Analysis

3.1 Data Description

Annual data on nine manufacturing sectors of the U.S. and fourteen other OECD countries are examined over the 1970-1993 period. The sectors (two-digit international standard industrial

[6] In Faruqee (1995), PTM occurs despite the constant elasticity demand functions since the cost is assumed to be separable for domestic and export markets.

classification (ISIC) codes in parentheses) are: food (31); textiles, apparel, and leather (32); wood products and furniture (33); paper, paper products, and printing (34); chemical products (35); non-metallic mineral products (36); basic metal industries (37); fabricated metal industries (38); and other manufacturing (39).[7] The country sample consists of Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Portugal, Sweden, the United Kingdom (U.K.), and the U.S.[8] The industry and country coverage is determined by the availability of the data.

The OECD Structural Analysis (STAN) Industrial Database and the International Sectoral Database (ISDB) contain data on value added in current and 1985 constant prices, gross output, labor compensation, imports, exports, number engaged, and gross capital stock for each sector.[9] Sectoral price deflator series are obtained by dividing the value added in current prices by that in constant prices. The data on bilateral nominal exchange rates vis-à-vis the U.S. dollar are obtained from IMF's International Financial Statistics (IFS).

3.2 Estimating the Rate of Parity Reversion: Unit Root Tests and the Mean Reversion Coefficient

Using the nominal exchange rate series and sectoral price deflators, we define the sectoral real exchange rate in logarithm of industry i between the U.S. and country j as

$$q_{i,t}^{j} = s_t^{j} + p_{i,t}^{j} - p_{i,t}^{us} \tag{3}$$

where s^j_t is the log of nominal exchange rate measured in dollars per unit of j's currency, $p^j_{i,t}$ and $p^{us}_{i,t}$ denote the log of sectoral price indexes of sector i of country j and of the U. S.,

[7] See the data appendix for a more detailed description of the classifications.

[8] Due to incomplete data coverage in the STAN database, we omitted Austria, Korea, Mexico, New Zealand and Spain from the analysis.

[9] The "number engaged" refers to the number of employees as well as self-employed, owner proprietors, and unpaid family workers of a given industry.

respectively, and t is the time subscript. With the nine industries for i and the fourteen countries for j, a total of 126 dollar-based bilateral sectoral real exchange rates are constructed.

In order to estimate the speed of parity reversion, we first identify the series that revert to the parity using unit root tests. It is well known that standard unit root tests, such as the augmented Dickey-Fuller (ADF) tests, possess low power against the alternative of a stationary but persistent process. While there is no strictly uniformly most powerful invariant test for the unit root hypothesis, a modified ADF test called the ADF-GLS test developed by Elliott, Rothenberg, and Stock (1996) is approximately uniformly most powerful invariant against the local alternatives. The superior performance of this test procedure is documented, for instance, by Pantula, Gonzalez-Farias and Fuller (1994) and Stock (1994). We therefore test the sectoral real exchange rates for a unit root using the ADF-GLS test.

The ADF-GLS$^\tau$ test which allows for a linear time trend is based on the following regression (for which the industry subscript and the country superscript are suppressed for brevity):

$$(1-L)q_t^\tau = \alpha_0 q_{t-1}^\tau + \sum_{k=1}^{p} \alpha_k (1-L) q_{t-k}^\tau + \varepsilon_t \qquad \textbf{(4)}$$

where q_t^τ is the locally detrended process under the local alternative of $\bar{\alpha}$ and is given by

$$q_t^\tau = q_t - \tilde{\gamma}' z_t \qquad \textbf{(5)}$$

with $z_t = (1, t)'$. $\tilde{\gamma}$ is the least squares regression coefficient of $\tilde{q}_t$ on $\tilde{z}_t$ where $(\tilde{q}_1, \tilde{q}_2 \ldots\ldots \tilde{q}_T) = (q_1, (1-\bar{\alpha}L)q_2, \ldots, (1-\bar{\alpha}L)q_T)$ and $(\tilde{z}_1, \tilde{z}_2 \ldots\ldots \tilde{z}_T) = (z_1, (1-\bar{\alpha}L)z_2, \ldots, (1-\bar{\alpha}L)z_T)$ with L being the lag operator. The local alternative $\bar{\alpha}$ is defined by $\bar{\alpha} = 1 + \bar{c}/T$ for which $\bar{c}$ is set to -13.5. The ADF-GLS$^\mu$ test, which does not allow for a linear time trend, involves the same procedure as the ADF-GLS$^\tau$ test, except that q^τ_t is replaced by the locally demeaned series q^μ_t, which is obtained by setting $z_t = 1$ and $\bar{c}$ to -7. The ADF-GLS test statistic is given by the usual t-statistic for $a_0 = 0$ against the alternative of $a_0 < 0$, and its statistical significance is evaluated using the finite sample critical values tabulated by Cheung and Lai (1995). The lag parameter p is determined in the following procedure. The maximum AR lag is set to 4 and the Akaike information criterion (AIC) is used to determine the first estimate of p. Then, the residuals from the selected model are checked for serial correlations. If there is no significant serial correlation in the estimated residuals, the number of lags determined by the AIC is used to conduct the test. Otherwise, the lag parameter will be increased by one until the resulting specification successfully removes serial correlation in the residuals.

Table 1 presents the summary of the ADF-GLS test results. According to the preliminary analysis, 31 of the 126 sectoral real exchange rates exhibit a significant deterministic time trend. Among these 31 series the ADF-GLS$^{\tau}$ test rejects the unit root hypothesis in 17 cases. Of the 95 sectoral real exchange rates without a significant time trend, the unit root null hypothesis is rejected in 51 cases. The rejection rate ranges from 35% for the basic metal industries and the other manufacturing to 78% for the non-metallic mineral products.

As a measure of the parity reversion rate for the I(0) sectoral real exchange rates, we define the mean reversion coefficient for industry i of country j as

$$MRC_i^j \equiv 1 + \hat{\alpha}_0 \qquad \textbf{(6)}$$

where $\hat{\alpha}_0$ is the estimated coefficient from the ADF-GLS equation (4).[10] The closer MRC^j_i is to unity, the more persistent is the sectoral real exchange rate, and thus, the slower is the speed of parity reversion. The last two columns of Table 1 report the ranges of MRC^j_i of the I(0) sectoral real exchange rate series. Among all I(0) sectoral real exchange rates, the value of MRC^j_i ranges from .053 (food, Canada) to .811 (fabricated metal products, Italy) with the

[10] AR(1) coefficients are often used as a proxy to capture the persistence of a time series. For instance, Campa and Wolf (1997) utilizes the AR(1) coefficient from a Dickey-Fuller test as their measure of the speed of parity reversion.

sample mean equal to .579.[11] Note the wide range of the rates at which the sectoral real exchange rates revert to the parity condition. The variation within sectors (across countries) is also fairly substantial in many cases, and is most pronounced for the food industry with the range from .053 (Canada) to .722 (Belgium). The basic metal industry has the narrowest range of MRC^j_i values from .411 (U.K.) to .504 (Japan). In section 4 we investigate the empirical relationships between the parity reversion rate, MRC^j_i, and market structure.

3.3 The Price-Cost Margin

Next we devise a proxy for market structure. Our first measure is the price-cost margin (PCM) which approximates the profitability of an industry. Define the PCM for industry i of country j in period t as

$$PCM^j_{i,t} = \frac{V^j_{i,t} - M^j_{i,t} - W^j_{i,t}}{V^j_{i,t}} = \frac{VA^j_{i,t} - W^j_{i,t}}{VA^j_{i,t} + M^j_{i,t}} \qquad (7)$$

where $V^j_{i,t}$ is the value of total production, $M^j_{i,t}$ is the cost of materials, $W^j_{i,t}$ is labor compensation, $VA^j_{i,t}$ (= $V^j_{i,t}$ - $M^j_{i,t}$) is the value added of industry i in country j in period t. Since PCM can be directly observed from accounting data, it is widely utilized as a measure of market structure (Campa and Goldberg, 1995; Domowitz, Hubbard, and Peterson, 1986 and 1987). In section 5 we will use an estimate of industry price markup over marginal cost (Hall, 1988) as an alternative measure of market structure to check the robustness of our findings.

The STAN Industrial database contains data on gross output, value added, and labor compensation. The cost of materials is calculated by subtracting nominal value added from nominal gross output. The information on PCM data is summarized in Table 2. To conserve space we report only the mean and standard deviation of the calculated PCMs for each

[11] In terms of half-lives, the range corresponds to 2.8 months to 3.3 years with the mean equal to 1.48 years.

industry in each country over the sample period. The data indicate that there is much variation in PCMs both across industries and countries.

3.4 The Intra-Industry Trade Index

Another way to capture the market structure of an industry is to characterize the nature of competition via the degree of product differentiation. For instance, an industry is better characterized as monopolistically competitive than perfectly competitive if domestic and foreign firms supply a variety of differentiated products that are imperfect substitutes for each other. The idea of utility gain from product differentiation (Dixit and Stiglitz, 1977) provides a plausible explanation for the predominance of intra-industry trade (IIT) among developed countries, and is an essential ingredient of the modern approach to international trade (Helpman, 1981; Grossman and Helpman, 1991; Krugman, 1995). In Dornbusch (1987) the monopolistic firm's pricing power is determined to be a function of the demand elasticity, which in turn depends crucially on the substitutability among varieties within an industry. Further, in Faruqee (1995) an increase in the elasticity of substitution between domestic and imported varieties intensifies the real exchange rate persistence.

We utilize the intra-industry trade index (Grubel and Lloyd, 1975) as our second measure of market structure to reflect the market power due to product differentiation. The IIT index of sector i in country j in period t is defined as

$$IIT_{i,t}^{j} \equiv 1 - \frac{|EX_{i,t}^{j} - IM_{i,t}^{j}|}{(EX_{i,t}^{j} + IM_{i,t}^{j})} \qquad \textbf{(8)}$$

where $EX^{j}_{i,t}$ and $IM^{j}_{i,t}$ represent sectoral exports and imports, respectively. A large value of the IIT index is interpreted as a high level of market power due to product differentiation. Table 3 presents the sample means and standard deviations of the IIT indexes. Not surprisingly, we observe fairly large IIT index values, particularly for the European countries. For many countries, the IIT index varies substantially across sectors, indicating heterogeneity in the market structure across the various manufacturing sectors.

4 The Persistence of Sectoral PPP Deviations

With the empirical measures constructed in the preceding section, we analyze the relationship between market structure and sectoral real exchange rate persistence. The first regression specification is

$$\begin{aligned} MRC_i^j = \beta_0 + \delta PCM_i^j + \beta_1 OPEN_i^j + \beta_2 INF_i^j + \beta_3 GOV^j \\ + \beta_4 SVAR^j + \beta_5 DIST^j + \eta_i^j \end{aligned} \tag{9}$$

where MRC^j_i is the mean reversion coefficient defined in section 3.2, PCM^j_i is the index of PCM defined as the sum of the sector i average PCM of the U.S. and that of country j, and η^j_i is the disturbance term. The second through the sixth regressors are included as control variables, and are defined and discussed below.

The fundamental idea of long-run PPP is that goods arbitrage ensures the parity condition across a range of individual goods over a certain time horizon.[12] Accordingly, trade activity affects the PPP adjustment rate. Also, in Faruqee (1995) an increase in openness encourages more frequent price adjustment by firms, and thus, reduces the real exchange rate persistence. To control for the effect of openness, we include the variable $OPEN^j_i$, which is defined as the sum of the sample average ratios of the imports plus exports to the total production in sector i of the U.S. and of country j.

The speed of parity reversion depends crucially on how quickly goods prices are adjusted. Given the existence of nominal rigidities, a higher inflation rate may lead to more rapid price adjustment (Ball and Mankiw, 1994). In accord with this view, empirical evidence indicates that PPP holds well for high inflation countries (Frenkel, 1978; McNown and

[12] Campa and Wolf (1997) dispute this view by reporting that large real exchange rate deviations and large trade deviations are not systematically related.

Wallace, 1989). Further, in their cross-country analysis Cheung and Lai (1999) find that a higher inflation is associated with lower real exchange rate persistence. These studies suggest that differences in sector-specific inflation may partly explain differences in sectoral real exchange rate persistence. Hence equation (9) includes INF^j_i, which is defined as the sum of the average sectoral inflation rates of industry i of the U.S. and of country j.[13]

Some structural models of PPP deviations consider government spending as an important demand-side factor in the short-run for creating a home goods bias (Frenkel and Razin, 1987; Froot and Rogoff, 1991; Rogoff, 1992). Froot and Rogoff (1995) and Gagnon and Rose (1995) document some empirical evidence for this effect. Also, Cheung and Lai (1999) find that government spending is positively correlated with real exchange rate persistence. We therefore include the variable GOV^j which denotes the average of the ratios of government consumption to gross domestic product (GDP) of country j, to control for the country-specific demand-side effect.

$SVAR^j$ in (9) represents the exchange rate variability measured by the standard deviation of first log differences of the nominal exchange rate between the U.S. and j. The variable is interpreted as a proxy for exchange rate uncertainty price-setters face. In his dynamic partial equilibrium model of a price setting firm with menu costs, Delgado (1991) shows that variability of the nominal exchange rate raises the level of uncertainty, and hence, intensifies price stickiness. In other words, firms become less willing to change their prices since the exchange rate may move back after the price change and another price change in the opposite direction may become necessary.

A popular view of PPP/LOP deviations is that transportation costs create a wedge between prices in two countries (Dumas, 1992; Obstfeld and Taylor, 1997; O'Connell and Wei, 1997; Wei and Parsley, 1995). It follows that a greater geographical distance can lead to

[13] We also considered country-specific inflation rates rather than sector-specific rates. For all of the specifications we estimated, the choice between the two different inflation rates does not alter the results significantly. The results based on the country-specific inflation rates are available upon request.

larger PPP deviations if transportation costs are proportional to distances (Wei and Parsley, 1995). In a recent study Campa and Wolf (1997) find that a greater geographical distance results also in *slower* PPP reversion. Thus, we add the variable $DIST^j$, which is the geographical distance in logarithm between the U.S. and country j to capture the transportation cost effect.[14]

Using the sample restricted to include only the I(0) sectoral real exchange rates, a truncated regression specification is employed to estimate equation (9). Maximum likelihood (ML) estimation results are summarized in Table 4.[15] For comparison purposes, the results of the ordinary least squares (OLS) estimations are also reported. In accordance with the hypothesis, the PCM term has a statistically significant positive effect on real exchange rate persistence. That is, sectors with a larger PCM, interpreted as a less competitive market structure, are associated with a slower rate of sectoral PPP reversion. The OLS estimation also yields a positive PCM effect. The effect of inflation is negative and significant, implying that industries with higher inflation rates experience faster sectoral real exchange rate parity reversion. This is also consistent with our prior. However, the effect of trade openness is puzzling. This variable has a significant positive effect indicating that the more open an industry is to international trade, the more persistent is its sectoral real exchange rate. The finding is counter-intuitive and contradicts the goods arbitrage view of PPP reversion. However, this result is not isolated; Campa and Wolf (1997) also report a similar result in their study of parity reversion among G7 currencies.[16] The trade openness effect presents a

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[14] We follow the common practice (Campa and Wolf, 1997; Wei and Parsley, 1995) of using the distances between national capitals as a proxy for the distances between countries.

[15] See Dhrymes (1984) and Maddala (1983). The Eicker-White method is used to calculate the asymptotic covariance matrix of the parameter estimates and standard errors. The method is a combination of analytical second derivatives and Berndt-Hall-Hall-Hausman method, and is robust to the distributional assumption, see White (1982).

[16] Campa and Wolf (1997) measures trade openness by bilateral trade shares.

new puzzle which needs to be addressed in future research. The two country-specific effects -- government consumption and nominal exchange rate volatility -- are also found to be statistically significant with the expected sign.

An increased share of government spending to GDP leads to a slower parity reversion; so, too, does increased exchange rate volatility. The effect of geographical distance is not significant and has a negative sign. The result is in contrast with that of Campa and Wolf (1997) who find that a greater distance is associated with a slower PPP reversion among G7 countries. However, theoretically the effect of geographical distance on the *speed* of reversion is ambiguous, unlike its effect on the *size* of PPP deviations.

To examine the effect of IIT, we estimate

$$\begin{aligned} MRC_i^j &= \beta_0 + \lambda IIT_i^j + \beta_1 OPEN_i^j + \beta_2 INF_i^j + \beta_3 GOV^j \\ &\quad + \beta_4 SVAR^j + \beta_5 DIST^j + \eta_i^j \end{aligned} \qquad \textbf{(10)}$$

where IIT^j_i is the sum of the sector i average IIT index of the U.S. and that of country j, and the other variables are as defined earlier. The last two columns of Table 4 summarize the results. Consistent with the prediction, IIT has a highly significant positive effect on real exchange rate persistence. In other words, sectors with substantial IIT activity, interpreted as an indication of imperfect competition due to product differentiation, tend to have more persistent sectoral PPP deviations. Although the effects of trade openness and government spending are essentially the same, replacing PCM with IIT leads to insignificant coefficients for inflation and exchange rate variability. The OLS estimation result is qualitatively similar to the ML one although the statistical significance level of the IIT effect declines to the 10 percent level.

In order to examine whether PCM and IIT are capturing different aspects of the market structure effect, we include both terms in the regression model simultaneously, and estimate (11) below.

$$
\begin{aligned}
MRC_i^j = {} & \beta_0 + \delta PCM_i^j + \lambda IIT_i^j + \beta_1 OPEN_i^j + \beta_2 INF_i^j + \beta_3 GOV^j \\
& + \beta_4 SVAR^j + \beta_5 DIST^j + \eta_i^j .
\end{aligned}
\tag{11}
$$

The results are reported in Table 5. The effects of PCM and IIT remain positive and significant at the 10% and 5% levels, respectively. This suggests that PCM and IIT are indeed capturing different aspects of the market structure effect and they both are positively associated with sectoral real exchange rate persistence. The effect of trade openness, although counterintuitive, is quite robust. While government spending indicates significant positive effect, the negative effect of sectoral inflation rates is not significant. Overall, the estimation results of (9), (10), and (11) suggest that market structure has a significant effect upon real exchange rate persistence. When the indicators of market imperfection, as measured by PCM and IIT index, increase, the corresponding sectoral real exchange rates become more persistent, and exhibit a slower reversion to sectoral PPP.

In order to see if the market structure variables as a group add to the explanatory power of distance, inflation, openness and government spending, we conducted a series of χ^2 tests. The hypothesis that the coefficients of PCM^j_i and IIT^j_i are jointly insignificant is rejected at the 5% marginal significance level. The effects of market structure on the sectoral real exchange rate persistence seem quite pervasive.

5 Additional Analyses

5.1 The Relative Price-Cost Margin

In this section, we examine a few more additional specifications to evaluate the robustness of the market structure effect reported in the previous section. In Table 2, we notice that some countries have relatively high average PCMs across sectors while others have relatively low values across sectors. A comparison between Australia and Belgium, or Portugal and Norway, serves as a good example. The high variation of PCMs across countries may imply that a direct comparison of sectors across countries based on the absolute values of PCMs does not capture the differences in market structure due to unobserved country-specific effects. To address this concern, we calculate the relative PCM (RPCM) as

$$RPCM^{j}_{i,t} = PCM^{j}_{i,t} - PCM^{j}_{m,t} \tag{12}$$

where $PCM^{j}_{i,t}$ is the PCM of sector i, and $PCM^{j}_{m,t}$ is the PCM of the total manufacturing sector of country j in period t.[17] Further let $RPCM^{j}_{i}$ denote the sum of the average sector i RPCMs of the U.S. and that of country j. Replacing PCM^{j}_{i} in (9) and (11) with $RPCM^{j}_{i}$, we re-estimate the models and report the results in Table 6. The results based on RPCM are qualitatively similar to the previous ones. The effect of RPCM is positive and statistically significant, and estimates of other coefficients are relatively unchanged. When both RPCM and IIT are included simultaneously, the effect of IIT dominates in terms of significance. The χ^2 test results indicate that RPCM and IIT together add significantly to the explanatory power of the model. The joint hypothesis of zero coefficients on both $RPCM^{j}_{i}$ and IIT^{j}_{i} is rejected at the 5% significance level. Overall, the effects of market structure on the sectoral real exchange rate persistence remain quite robust as we replace PCM with RPCM.

5.2 The Price Markup over Marginal Cost

The PCM measures the margin between price and average cost. Another measure of market power is the gap between price and *marginal* cost which is usually not directly observable from the data. Hall (1988) proposes a technique to estimate the industry price markup over marginal cost, and applies it to the U.S. manufacturing industries. The method is adopted by other studies including Domowitz, Hubbard, and Petersen (1988), Hall (1990), Harrison (1994), and Levinsohn (1993).[18]

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[17] PCM^{j}_{m} is calculated from the data on the total manufacturing (ISIC 30).

[18] See Bresnahan (1989) and Feenstra (1995) for a summary and discussion of Hall's (1988) method.

The Hall (1988) method estimates the price markup over marginal cost by

$$\Delta y_{i,t} - \alpha_{i,t}^{L}\Delta l_{i,t} - \alpha_{i,t}^{M}\Delta m_{i,t} = (1-\rho)\theta_i + \rho\Delta y_{i,t} + (1-\rho)\Delta a_{i,t} \tag{13}$$

where

$$\rho \equiv (P - MC)/P \tag{14}$$

$y_{i,t} \equiv \ln(Y_{i,t}/K_{i,t})$, $l_{i,t} \equiv \ln(L_{i,t}/K_{i,t})$, $m_{i,t} \equiv \ln(M_{i,t}/K_{i,t})$ and $a_{i,t} \equiv \ln A_{i,t}$, for which $Y_{i,t}$, $K_{i,t}$, $L_{i,t}$, $M_{i,t}$ and $A_{i,t}$ denote output, capital, labor, material input, and random productivity shock, respectively. θ_i represents the constant rate of Hicks-neutral technology progress. $\alpha^{L}_{i,t}$ and $\alpha^{M}_{i,t}$ denote the labor share and the material share of the output, respectively. P is the price and MC is the marginal cost. Under the assumption of constant returns to scale and perfect competition, the price is equal to the marginal cost, and hence, ρ in (13) equals 0. Under imperfect competition, however, price exceeds marginal cost, and therefore ρ takes a positive value.

To account for the endogeneity, we estimate (13) using two commonly-adopted instruments: the growth rate of crude fuel prices (Hall, 1988) and the current and lagged growth rate of GDP (Domowitz, Hubbard, and Peterson, 1988).[19] Again, the sectoral data are taken from the STAN Industrial Database and the ISDB. The real gross capital stock is used as a measure for the capital. The labor is proxied by the number of total employees. The share of labor in the output is measured by the ratio of total compensation to the total production. The annual real GDP data is obtained from IFS. The monthly data on crude fuel price for the manufacturing industry is obtained from DRI Basic Economics, and the annual series is derived from period averages. Then, the sector-specific deflator is used to calculate

[19] It is argued that changes in crude fuel price affect production decisions, and thus, changes in industry input and output. Further, changes in crude fuel prices are unlikely to cause the random component of productivity shocks in the short-run. GDP growth is included to capture the aggregate demand factor under two assumptions: first, there is no common element to productivity shocks across sectors; second, no sector is large enough to affect GDP (Domowitz, Hubbard, and Peterson, 1988).

the real sector-specific oil price. Unfortunately, due to the limited coverage of the capital stock data in ISDB, we are unable to estimate (13) for Australia, the Netherlands, and Portugal. In addition, there are seven other sectors for which markups are not estimated due to missing observations.[20] The reduction of the sample size further limits our subsequent analysis of the effect of price markup on sectoral real exchange rate persistence.

The estimated values of ρ in (13) are reported in Table 7. In 71 cases (including the U.S.) the estimates of ρ have a significantly positive value (evaluated at the 5% marginal significance level). A comparison of Table 7 and Table 2 reveals that the estimated markups are much more variable than the calculated margins. For the cases where the estimated ρ is not significantly different from zero, the price markup is assumed to be zero.

With the estimated price markup, we examine the effect of market structure on the sectoral real exchange rate persistence using

$$\begin{aligned} MRC_i^j &= \beta_0 + \varphi MKUP_i^j + \beta_1 OPEN_i^j + \beta_2 INF_i^j + \beta_3 GOV^j \\ &\quad + \beta_4 SVAR^j + \beta_5 DIST^j + \eta_i^j \end{aligned} \tag{15}$$

where $MKUP^j_i$ denotes the sum of the estimated sector i price markup over marginal cost of the U.S. and that of country j. Other variables are as defined earlier. The estimation results are presented in the second and third columns of Table 8. With the reduced sample, (15) fits the data rather poorly. None of the independent variables have a significant coefficient estimate. Since the lack of statistical significance may be simply the result of the reduced sample size, we add the other measures of market structure, PCM and IIT index, one at a time to (15) and examine their effects in this sample.

As presented in the fourth and fifth columns of Table 8, the effects of PCM and trade openness are both statistically significant as found in section 4. A similar result is obtained

[20] They are as follows: wood products and furniture for Belgium, Italy, and Japan; non-metallic mineral products for Norway; other manufacturing for Belgium, France, and Norway.

when the IIT index, instead of PCM, is added to (15). These results are reported in columns 6 and 7 of Table 8. Both the IIT index and trade openness have significant positive coefficients. The coefficient estimate of price markup is statistically insignificant, although it has a positive sign, as expected. It is likely that the insignificance of the price markup variable is related to the fact that it is a generated regressor, and is estimated with considerable inaccuracy.[21] Nonetheless, when all of the three measures of market structure are included in the model, the price markup shows up significant at the 10 percent level, while PCM and the IIT index are significant at the 5 and 1 percent levels, respectively. These results suggest that the price markup, PCM, and IIT variables are capturing different market structure effects. Hence, each variable can be thought of complementing, rather than substituting for, each other. Note that all three measures have signs that are consistent with the priors. That is, they all suggest that market imperfection is related to persistence in sectoral real exchange rates. Overall, the effects of market structure variables remain fairly robust.[22]

6 Concluding Remarks

One of the more intriguing aspects of the post-Bretton Woods period is the marked persistence exhibited by real exchange rates. While the slow speed of reversion to PPP is quite extensively documented, the determinants of this sluggish adjustment have not been identified. One set of factors that has not heretofore been examined is suggested by models of

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[21] Another possible reason for the insignificant results is the assumption required by the Hall method that the price markup is constant over time as implied by (15). The ability to price discriminate across markets as exchange rates fluctuate requires that imperfectly competitive firms can vary the price markup.

[22] The estimations of (17) through (20) are repeated by replacing the estimated price markups with the estimated *relative* price markups. These relative markups are obtained by subtracting the estimated markup for the entire manufacturing industry from the individual estimated markups. The results are qualitatively similar to those based on the estimated price markups, and are available upon request.

imperfect competition. These models suggest that markets characterized by trade in differentiated but substitutable goods and segmentation between countries will evidence slow reversion to PPP.

Using data on U.S. dollar-based sectoral real exchange rates for fourteen OECD countries, this study investigates the empirical relationship between several measures of market structure and real exchange rate persistence, taking into consideration the effects of macroeconomic variables commonly believed to affect PPP adjustment. The econometric results reveal considerable evidence for the hypothesis that market imperfection is associated with high PPP persistence. In general, the two measures of market imperfection, PCM and IIT index, are significant across different specifications and have a positive impact on real exchange rate persistence. The robustness of the market structure effects stands in stark contrast with the results pertaining to the macroeconomic variables. The coefficient estimates on these variables tend to vary across model specifications, and occasionally, have a sign different from that predicted by theory. Overall, our analysis uncovers some positive evidence of market structure effects on real exchange rate persistence.

The novelty of these results should be noted. In particular, the use of the IIT index to measure the degree of substitutability between differentiated product is, to our knowledge, quite new. Moreover, while the PCM has been linked to the degree of market imperfection in previous work, use of this variable as a determinant of reversion rates is, as has been remarked earlier, an innovation.

Future study of the relationship between market structure effects should benefit from the availability of better quality data on price and cost structure at finer detail. Unfortunately, at the current moment, it is difficult to compile a data set encompassing a wide number of countries when dealing with data at the ISIC 3-digit level of classification. An interesting future research project would entail the collection and construction of data that would allow for direct examination of these effects at the disaggregated industry level.

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Table 1: Summary of the Unit Root Test Results

ISIC code and classification description	No trend		Linear time trend		I(0) series		
	I(0)	I(1)	I(0)	I(1)	Total	Range of MRC^j_i	
						Min.	Max.
31: Food	4	7	3	0	7	.053	.722
32: Textiles, Apparel & Leather	4	4	4	2	8	.205	.778
33: Wood Products & Furniture	5	8	1	0	6	.383	.743
34: Paper, Paper Products & Printing	8	3	0	3	8	.283	.723
35: Chemical Products	8	5	0	1	8	.560	.733
36: Non-Metallic Mineral Products	4	2	7	1	11	.376	.739
37: Basic Metal Industries	4	6	1	3	5	.411	.504
38: Fabricated Metal Products	9	4	1	0	10	.343	.811
39: Other Manufacturing	5	5	0	4	5	.575	.752
Total	51	44	17	14	68	.053	.811

Notes: Each entry, except in the last two columns, presents the number of the sectoral real exchange rate series in the corresponding category. The columns under the heading "No trend" contain the results for the series without a significant deterministic trend at the 5% level. Similarly, the columns under "Linear time trend" contain the results for the series which exhibit a significant deterministic trend. I(0) and I(1) indicate the orders of integration as determined by

ADF-GLS tests. The10% finite sample critical values from Cheung and Lai (1995) is used to evaluate significance of ADF-GLS test statistics. The last three columns contain the total number of I(0) series in each industry and the corresponding range of the mean reversion coefficient (MRC^j_i) defined by equation (6).

Table 2: Sample Mean (Standard Deviation) of the Price-Cost Margin

ISIC codes	U.S.	Australia	Belgium	Canada	Denmark	Finland	France	Germany
31	.121	.150	.126	.122	.083	.087	.149	.172
	(.018)	(.015)	(.017)	(.021)	(.017)	(.010)	(.024)	(.014)
32	.072	.122	.074	.097	.100	.115	.102	.103
	(.009)	(.021)	(.014)	(.010)	(.011)	(.013)	(.013)	(.017)
33	.128	.151	.076	.080	.097	.101	.122	.104
	(.015)	(.034)	(.016)	(.030)	(.014)	(.036)	(.047)	(.019)
34	.148	.179	.086	.147	.082	.114	.103	.126
	(.022)	(.028)	(.022)	(.026)	(.012)	(.033)	(.032)	(.015)
35	.161	.186	.076	.097	.122	.153	.209	.167
	(.037)	(.034)	(.023)	(.024)	(.030)	(.024)	(.054)	(.031)
36	.103	.203	.029	.171	.146	.182	.137	.163
	(.027)	(.019)	(.046)	(.026)	(.027)	(.032)	(.041)	(.025)
37	.083	.142	.032	.077	.064	.092	.105	.085
	(.017)	(.022)	(.056)	(.022)	(.038)	(.030)	(.041)	(.017)
38	.089	.128	.058	.101	.085	.130	.120	.101
	(.012)	(.014)	(.019)	(.008)	(.011)	(.019)	(.029)	(.016)
39	.150	.166	1.345	.113	.198	.171	.140	.205
	(.040)	(.035)	(.565)	(.015)	(.052)	(.038)	(.042)	(.020)

(Table 2 continued)

ISIC codes	Italy	Japan	Netherlands	Norway	Portugal	Sweden	U.K.
31	.155	.222	.083	.064	.179	.069	.234
	(.018)	(.019)	(.016)	(.043)	(.036)	(.038)	(.025)
32	.152	.097	.070	.077	.139	.070	.031
	(.037)	(.020)	(.019)	(.022)	(.065)	(.034)	(.016)
33	.188	.110	.091	.099	.192	.110	.134
	(.023)	(.023)	(.024)	(.029)	(.042)	(.023)	(.020)
34	.148	.148	.129	.093	.235	.114	.055
	(.017)	(.022)	(.014)	(.015)	(.048)	(.027)	(.020)
35	.129	.161	.150	.109	.124	.144	.057
	(.013)	(.037)	(.050)	(.037)	(.062)	(.040)	(.017)
36	.203	.186	.166	.139	.257	.072	.094
	(.020)	(.029)	(.046)	(.023)	(.042)	(.060)	(.034)
37	.120	.148	.145	.122	.193	.041	.019
	(.052)	(.021)	(.048)	(.047)	(.086)	(.052)	(.031)
38	.151	.151	.107	.063	.171	.074	.032
	(.012)	(.016)	(.021)	(.014)	(.035)	(.021)	(.015)
39	.149	.284	.145	.102	.206	-.481	.158
	(.015)	(.020)	(.025)	(.034)	(.069)	(.137)	(.109)

Notes: Each entry and the number in the parentheses give the sample average and standard deviation of the price-cost margin for 1970-93, respectively. The ISIC codes denote the industry classifications as follows: food (31); textiles, apparel, and leather (32); wood products and furniture (33); paper, paper products, and printing (34); chemical products (35); non-metallic mineral products (36); basic metal industries (37); fabricated metal industries (38); and other manufacturing (39).

Table 3: Sample Mean (Standard Deviation) of the Intra-Industry Trade Index

ISIC codes	U.S.	Australia	Belgium	Canada	Denmark	Finland	France	Germany
31	.880	.319	.932	.890	.496	.774	.916	.780
	(.089)	(.113)	(.034)	(.073)	(.042)	(.123)	(.032)	(.137)
32	.476	.538	.945	.332	.785	.875	.861	.746
	(.140)	(.151)	(.051)	(.040)	(.062)	(.129)	(.068)	(.048)
33	.534	.463	.927	.335	.779	.180	.681	.856
	(.120)	(.234)	(.049)	(.085)	(.081)	(.083)	(.043)	(.080)
34	.890	.196	.850	.359	.546	.094	.762	.911
	(.071)	(.038)	(.031)	(.049)	(.082)	(.027)	(.031)	(.076)
35	.933	.570	.906	.875	.712	.626	.962	.861
	(.057)	(.107)	(.032)	(.094)	(.117)	(.156)	(.030)	(.056)
36	.730	.231	.781	.572	.860	.847	.956	.867
	(.147)	(.050)	(.063)	(.088)	(.083)	(.111)	(.044)	(.048)
37	.508	.403	.681	.664	.477	.805	.954	.894
	(.133)	(.101)	(.071)	(.071)	(.107)	(.074)	(.028)	(.074)
38	.849	.279	.945	.820	.923	.801	.911	.605
	(.072)	(.074)	(.037)	(.050)	(.059)	(.126)	(.058)	(.089)
39	.519	.425	.962	.431	.793	.800	.882	.930
	(.158)	(.126)	(.027)	(.056)	(.105)	(.100)	(.059)	(.055)

(Table 3 continued)

ISIC codes	Italy	Japan	Netherlands	Norway	Portugal	Sweden	U.K.
31	.613	.276	.675	.746	.858	.539	.676
	(.076)	(.132)	(.031)	(.073)	(.093)	(.089)	(.091)
32	.521	.728	.786	.260	.514	.494	.794
	(.057)	(.212)	(.079)	(.048)	(.076)	(.041)	(.119)
33	.776	.278	.470	.601	.176	.417	.284
	(.105)	(.214)	(.070)	(.166)	(.103)	(.088)	(.071)
34	.815	.826	.864	.691	.621	.215	.603
	(.064)	(.069)	(.029)	(.144)	(.105)	(.052)	(.047)
35	.865	.925	.709	.808	.532	.686	.909
	(.054)	(.052)	(.046)	(.104)	(.093)	(.159)	(.044)
36	.492	.416	.750	.528	.731	.783	.822
	(.065)	(.170)	(.074)	(.127)	(.121)	(.100)	(.128)
37	.833	.576	.948	.784	.286	.881	.909
	(.123)	(.228)	(.052)	(.057)	(.106)	(.067)	(.046)
38	.844	.264	.892	.638	.491	.866	.867
	(.068)	(.068)	(.030)	(.074)	(.107)	(.036)	(.109)
39	.424	.827	.798	.335	.726	.645	.932
	(.080)	(.097)	(.055)	(.060)	(.160)	(.151)	(.038)

Notes: Each entry and the number in the parentheses give the sample average and standard deviation of the intra-industry trade index for 1970-93, respectively. The ISIC codes denote the industry classifications as follows: food (31); textiles, apparel, and leather (32); wood products and furniture (33); paper, paper products, and printing (34); chemical products (35); non-metallic mineral products (36); basic metal industries (37); fabricated metal industries (38); and other manufacturing (39).

Table 4: Effects of Price-Cost Margin and Intra-Industry Trade on Sectoral Real Exchange Rate Persistence

	Price-Cost Margin		Intra-Industry Trade	
	ML	OLS	ML	OLS
Price-Cost Margin	.5579**	.7397**		
(PCM^j_i)	(.2393)	(.3196)		
Intra-Industry Trade			.1367***	.1373*
(IIT^j_i)			(.0433)	(.0716)
Trade Openness	.0700***	.0967***	.0720***	.0897**
($OPEN^j_i$)	(.0195)	(.0357)	(.0197)	(.0357)
Inflation	-1.0778**	-1.2310**	-.4112	-.5900
(INF^j_i)	(.4566)	(.5504)	(.5162)	(.5772)
Government Spending	.7884**	1.2704**	.6507**	.9712**
(GOV^j)	(.3322)	(.4958)	(.3234)	(.4863)
Exch. Rate Variability	1.6380**	1.6934*	.2612	.1996
($SVAR^j$)	(.6573)	(.9920)	(.7376)	(1.1274)
Geographical Distance	-.0157	.0128	.0193	.0600
($DIST^j$)	(.0244)	(.0376)	(.0255)	(.0411)
Constant	.3468	-.0806	.0868	-.3501
	(.2136)	(.3177)	(.2459)	(.3798)

Notes: The price-cost margin results are based on equation (9) in the text. The intra-industry trade results are based on equation (10) in the text. The entries under the heading "ML" present the maximum likelihood estimation results of the truncated regression. The entries under the heading "OLS" present the results obtained by ordinary least squares. ***, **, and * indicate 1, 5, and 10% statistical significance, respectively. The standard errors are provided in the parentheses.

Table 5: Combined Effects of Price-Cost Margin and Intra-Industry Trade on Sectoral Real Exchange Rate Persistence

	ML	OLS
Price-Cost Margin	.4468*	.6367*
(PCM^j_i)	(.2491)	(.3237)
Intra-Industry Trade	.1157**	.1074
(IIT^j_i)	(.0466)	(.0716)
Trade Openness	.0806***	.1045***
($OPEN^j_i$)	(.0200)	(.0357)
Inflation	-.6908	-.9053
(INF^j_i)	(.5023)	(.5865)
Government Spending	.7937**	1.2207**
(GOV^j)	(.3304)	(.4919)
Exch. Rate Variability	.6172	.8106
($SVAR^j$)	(.7232)	(1.1448)
Geographical Distance	.0095	.0399
($DIST^j$)	(.0254)	(.0414)
Constant	.0527	-.3771
	(.2429)	(.3714)

Notes: The results are based on equation (11) in the text. The entries under the heading "ML" present the maximum likelihood estimation results of the truncated regression. The entries under the heading "OLS" present the results obtained by ordinary least squares. ***, **, and * indicate 1, 5, and 10% statistical significance, respectively. The standard errors are provided in the parentheses.

Table 6: Effects of Relative Price-Cost Margin and Intra-Industry Trade on Sectoral Real Exchange Rate Persistence

	Relative Price-Cost Margin		Relative Price-Cost Margin and Intra-Industry Trade	
	ML	OLS	ML	OLS
Relative PCM	.5128**	.6691*	.4099	.5693
($RPCM^j_i$)	(.2491)	(.3509)	(.2555)	(.3516)
Intra-Industry Trade			.1220***	.1171
(IIT^j_i)			(.0464)	(.0718)
Trade Openness	.0623***	.0864**	.0752***	.0965***
($OPEN^j_i$)	(.0184)	(.0355)	(.0194)	(.0355)
Inflation	-.9222**	-1.0278*	-.5477	-.7039
(INF^j_i)	(.4545)	(.5459)	(.4938)	(.5740)
Government Spending	.5420*	.9306*	.5989*	.9279*
(GOV^j)	(.3265)	(.4872)	(.3265)	(.4807)
Exch. Rate Variability	1.5482**	1.5082	.4915	.5752
($SVAR^j$)	(.6610)	(.9954)	(.7260)	(1.1366)
Geographical Distance	-.0085	.0235	.0166	.0514
($DIST^j$)	(.0247)	(.03768)	(.0255)	(.0409)
Constant	.4545**	.0622	.1215	-.2823
	(.2052)	(.3166)	(.2416)	(.3772)

Notes: The relative price-cost margin results are based on equation (9) in the text with $RPCM^j_i$ replacing PCM^j_i. The relative price-cost margin and intra-industry trade results are based on equation (11) in the text with $RPCM^j_i$ replacing PCM^j_i. The entries under the heading "ML" present the maximum likelihood estimation results of the truncated regression. The entries under the heading "OLS" present the results obtained by ordinary least squares. ***, **, and * indicate 1, 5, and 10% statistical significance, respectively. The standard errors are provided in the parentheses.

Table 7: Price Markup Estimates

ISIC Codes	U.S.	Belgium	Canada	Denmark	France	Germany
31	.0815*	.6463**	.1974**	1.4609*	.0656	.1728*
	(.0330)	(.1860)	(.0514)	(.7087)	(.0660)	(.0683)
32	.3113**	.4413*	.3868**	1.0953**	.2673**	.2487
	(.0407)	(.2102)	(.0783)	(.3847)	(.0689)	(.1393)
33	.4371**	n.a.	.8000**	1.0909**	.2815**	.5499**
	(.0625)		(.2083)	(.1407)	(.0877)	(.1595)
34	.4534**	.6740*	.4362**	1.0255**	.2734**	.3925
	(.1237)	(.3379)	(.0975)	(.1115)	(.0825)	(.5301)
35	.1416	9.3557	.1244**	-0.3436	.0962*	-.0527
	(.0934)	(29.3259)	(.0423)	(5.3205)	(.0424)	(.1054)
36	.5515**	.8821	.6088**	.9131**	.6723**	.5466**
	(.0821)	(.7206)	(.0972)	(.1237)	(.1501)	(.1526)
37	.2997**	1.0092	.3347**	.9958**	.1017	.6037
	(.0578)	(.1823)	(.0786)	(.1216)	(.0891)	(.7790)
38	.3415**	.4374	.3937**	1.2012	.2673**	.3420**
	(.0294)	(.3036)	(.0575)	(.9164)	(.0887)	(.0647)
39	.4809**	n.a.	.3347*	3.6438	n.a.	.3946
	(.0963)		(.1435)	(14.0524)		(.2165)

(Table 7 continued)

ISIC Codes	Italy	Japan	Norway	Sweden	U. K.
31	.2617**	.2919**	.9739**	1.0387**	.0995
	(.0618)	(.0640)	(.0717)	(.2978)	(.0593)
32	.2394**	.0647	.7623**	.7848**	.2347**
	(.0368)	(.0878)	(.1734)	(.1517)	(.0667)
33	n.a.	n.a.	1.0680**	.9552**	.4278**
			(.2663)	(.1236)	(.0621)
34	.3943**	.2816**	.9738**	.9944**	.2410**
	(.0416)	(.0565)	(.1290)	(.0888)	(.0546)
35	.4487*	.1748**	1.0743**	.9242**	.0419
	(.2016)	(.0582)	(.1839)	(.3541)	(.0305)
36	.5931**	.4073**	n.a.	.7145**	.3677**
	(.1180)	(.1369)		(.1191)	(.1219)
37	.2128**	.3103**	.9972	1.1557**	.1155**
	(.0719)	(.0781)	(.2095)	(.2254)	(.0287)
38	.5313**	.3725**	.8213*	.7272**	.3018**
	(.0744)	(.1191)	(.4134)	(.1243)	(.0770)
39	.2699**	.3468*	n.a.	-0.3922	.0068
	(.0240)	(.1570)		(.8420)	(1.3262)

Notes: Each entry shows the estimated price markup. The numbers in parentheses are standard errors. ** and * indicate statistical significance at the 1% and 5% levels, respectively. The ISIC codes denote the industry classifications as follows: food (31); textiles, apparel, and leather (32); wood products and furniture (33); paper, paper products, and printing (34); chemical products (35); non-metallic mineral products (36); basic metal industries (37); fabricated metal industries

(38); and other manufacturing (39). "n.a." indicates that the corresponding price markup is not estimated due to incomplete data coverage. Due to the data limitation, price markup estimates are not available for Australia, Finland, and the Netherlands.

Table 8: Effects of Price Markup, Price-Cost Margin, and Intra-Industry Trade on Real

	Price Markup		Price Markup and Price-Cost Margin		Price Markup and Intra-Industry Trade		Price Markup, Intra-Industry Trade, and Price-Cost Margin	
	ML	OLS	ML	OLS	ML	OLS	ML	OLS
Price-Cost Margin (PCM^j_i)			.7380** (.3215)	1.0591** (.4705)			.6299** (.3123)	.8890* (.4563)
Intra-Industry Trade (IIT^j_i)					.1854*** (.0545)	.2078** (.0909)	.1649*** (.0520)	.1690* (.0880)
Price Markup ($MKUP^j_i$)	.0026 (.0402)	.0561 (.0740)	.0469 (.0438)	.1129 (.0748)	.0428 (.0424)	.1013 (.0730)	.0742* (.0407)	.1328** (.0650)
Trade Openness ($OPEN^j_i$)	.0442 (.0311)	.0621 (.0572)	.0904*** (.0346)	.1290** (.0621)	.0649** (.0316)	.0807 (.0550)	.0928** (.0369)	.1179* (.0589)
Inflation (INF^j_i)	-.5037 (.8559)	-.8797 (1.2149)	-.8383 (.8035)	-1.3340 (1.1736)	-.4670 (.7295)	-.7961 (1.1546)	-.6475 (.6878)	-.9680 (1.0915)
Government Spending (GOV^j)	.6510 (.4496)	.8971 (.7036)	.5658 (.4568)	.8279 (.6702)	.6025 (.4228)	.7051 (.6736)	.6061 (.4177)	.8457 (.6106)
Exch. Rate Volatility ($SVAR^j$)	.2098 (2.1220)	.9573 (2.7523)	1.8797 (1.9990)	3.8376 (2.9151)	-.7711 (2.0050)	-.2112 (2.6640)	.1520 (1.7047)	1.4348 (2.6541)
Geographical Distance ($DIST^j$)	.0415 (.0833)	.0414 (.1094)	-.0195 (.0774)	-.0704 (.1154)	.0616 (.0775)	.0752 (.1050)	.0309 (.0658)	.0140 (.1037)
Constant	.1334	-.0598	.2759	.2920	-.2435	-.5408	-.2277	-.4110

Notes: The price markup results are based on equation (15) in the text. The other results are obtained by adding the price-cost margin, intra-industry trade, and both price-cost margin and intra-industry trade variables to (15), respectively. The entries under the heading "ML" present the maximum likelihood estimation results of the truncated regression. The entries under the heading "OLS" present the results obtained by ordinary least squares. ***, **, and * indicate 1, 5, and 10% statistical significance, respectively. The standard errors are provided in the parentheses.

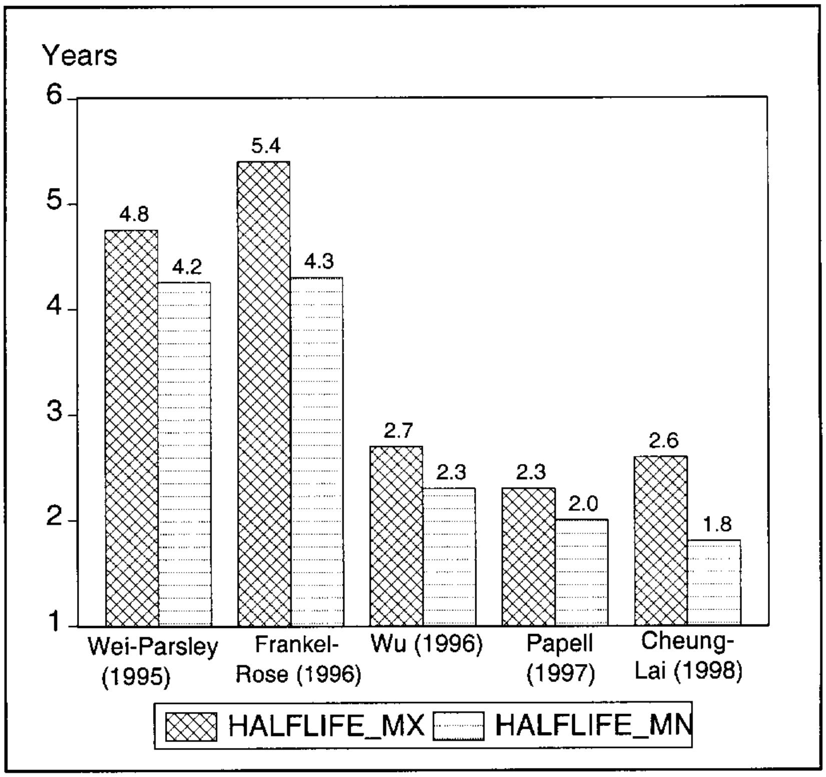

Figure 1: Half-lives of PPP deviations from various studies.

Data Appendix

1. Sectoral data source:

OECD Structural Analysis (STAN) Industrial Database (December 1994), and OECD International Sectoral Data base (ISDB) (version 95)

2. Countries:

the U.S., Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Portugal, Sweden, the U.K.

3. Industries:

ISIC Code	Classification Description	Products Contained in the Classification
31	Food, beverages, and tobacco	Food, beverages, and tobacco
32	Textiles, apparel, and leather	Textiles, wearing apparel, leather & products, footwear
33	Wood products and furniture	Wood products, furniture & fixtures
34	Paper, paper products, and printing	Paper & products, printing & publishing
35	Chemical products	Industrial chemicals, other chemicals, drugs & medicines, other chemical products, petroleum refineries, petroleum & coal products, plastic products (not elsewhere classified)
36	Non-metallic mineral products	Pottery, China etc., glass & product, non-metallic products (not elsewhere classified)
37	Basic metal industries	Iron & steel, non-ferrous metals
38	Fabricated metal products	Metal products, non-electrical machinery, office & computing machinery, machinery & equipment (not elsewhere classified), electrical machinery, radio, TV & communication equipment, electrical apparatus (not elsewhere specified), transport equipment, shipbuilding & repairing, radio equipment, motor vehicles, motor cycles & bicycles, aircraft, transport equipment (not elsewhere classified), professional goods
39	Other manufacturing	Other manufacturing (not elsewhere classified)