Sectoral Real Exchange Rate Adjustments in Europe*

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Abstract

This paper presents estimates of the half–life (the time required for a shock to the real exchange rate to dissipate by one half) of euro area sectoral producer price based real exchange rates. We take heterogeneity across countries into account using Random Coefficient Models and Mean Group estimators as well as cross–sectional dependence. Our estimates show that the half–life falls considerably when using sectoral data compared to the consensus view of three to five years. For 9 out of 24 sectors, the estimated half–life is less than a year. When comparing point estimates prior to and after the introduction of the euro we find a somewhat higher degree of persistence during the former period. On average, however, there is no significant change in the mean, median and variance of estimated persistence. The large differences in estimated half–lives across sectors can be attributed to the type of industry. Capital–intensive and marketing–driven industries with small price–cost margins typically adjust rapidly to shocks whereas labor–intensive and technology–drive industries with large price–cost margins adjust slowly.

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1 Introduction

Most economists believe in some variant of the Purchasing Power Parity (PPP) theory as an underlying theory of the exchange rate and as an anchor of long–run real exchange rates. However, the empirical support of this theory is not very solid, see the surveys by Rogoff (1996) and Taylor (2006). Real exchange rates appear to be, if not nonstationary, very persistent. The conventional view holds that deviations from PPP disappear very slowly with the half–life (defined as the time required for a unit shock to the real exchange rate to dissipate by one half) contained in the range of three to five years (Rogoff (1996)). This empirical observation seems not to be consistent with the very large short–term volatility of real exchange rates and, thus, constitutes a puzzle. Furthermore, the estimated half–life seems too long to be explained by sticky-price models.

One explanation to the PPP puzzle is aggregation bias both across products and over time. Taylor (2001) shows that the upward bias of estimated half–life is increasing in the degree of temporal aggregation whereas Imbs, Mumtaz, Ravn and Rey (2005) suggest that there is an upward bias if there is heterogeneity across products.[1] Chen (2004) examining European sectoral (six countries and 17 sectors) real exchange rates shows that estimated half–lives are much less persistent than the conventional view holds. Her estimates of the half–life is in the range of four months and two years. Cheung, Chinn and Fujii (2001) study 9 manufacturing sectors for 14 OECD countries and find that the half–lives are in the range of 2.8 months to 3.3 years with a mean of 1.48 years. They also find a substantial variation within sectors (across countries) and across sectors.

Recent findings also suggest that the key source of the slow mean reversion is adjustments of the nominal exchange rate and not adjustment of the relative price (Cheung, Lai and Bergman (2004)). The reason why real exchange rates adjust slowly is that even though nominal exchange rates move quickly, they often move in the wrong direction whereas relative prices move slowly but often in the correct direction.

In this paper we provide new measures of the half–life of deviations from purchasing power parity (PPP) at a sectoral level.[2] We use disaggregated sectoral producer price based real exchange rates at the 3–digit Nace Rev.1.1 level for euro area countries to examine whether the estimated half–life is affected by aggregation bias as suggested Imbs, Mumtaz, Ravn and Rey (2005), how the absence of nominal exchange rate adjustments affects the estimated half–life within the euro area countries and whether there are structural breaks in the estimated half–life coinciding with the introduction of the euro in 1999. All our estimates correct for two potential biases, the aggregation bias associated

[1] Chen and Engel (2005) argue that cross–sectional aggregation bias cannot explain the puzzle. They show, using the same data as Imbs et.al., that small–sample correction reduces the bias and that heterogeneity across products cannot explain the puzzle.

[2] The present study is not concerned with the aggregation bias issue as such. Instead, we take the finding that aggregation bias may be a potential problem when estimating the half–life as given and provide new estimates based on sectoral real exchange rates using heterogeneous panel data models.

with heterogeneous coefficients in the panel data and cross–sectional dependence. We estimate Random Coefficient Models and use the Mean Group estimator (Pesaran and Smith (1995), Hsiao and Pesaran (2006)) to handle the first bias and we use the method suggested by Bai and Ng (2002) to correct for the second bias.

The paper extends the existing literature in different directions. First, to our best knowledge this is the first study of sectoral real exchange rate adjustments in the euro area.[3] Second, we provide a comparison of half–lives between different periods allowing us to study the role of the common currency. Third, our paper is an extension of the paper by Chen (2004) who also studied domestic producer prices but only for the period prior to the introduction of the euro. We also study a larger set of sectoral real exchange rates. Fourth, following Cheung, Chinn and Fujii (2001) we explain differences in the degree of persistence across sectors by market imperfections, i.e., the pricing behavior of firms.

Our results show that the half–life falls considerably when using disaggregated prices compared to using aggregated data, thus supporting the view that aggregation bias can explain the PPP puzzle. However, the estimated half–life varies considerably between different types of goods, from a few months for some sectors to several years for other sectors. These differences in the degree of persistence can be explained by the industry type. Our results suggest that capital–intensive and marketing–driven industries with small price–cost margins adjust very rapidly whereas technology–driven and labor–intensive industries with large price–cost margins adjust slowly to shocks to the real exchange rate. These results hold both during the EMU period and during the 1990's for the euro area.

The remainder of the paper is structured in the following way. In section 2 we discuss the methodology used in the empirical analysis. Section 3 describes the data. Section 4 contains the empirical analysis and section 5 summarizes the paper.

2 Methodology

Assume that the sectoral real exchange rate for country i, q_{it}, is generated by the following autoregressive model of the first order (AR(1))

$$q_{it} = \alpha_i + \rho q_{it-1} + \varepsilon_{it} \quad i = 1, \ldots, N \;\text{ and }\; t = 1, \ldots, T \tag{1}$$

where $\mid \rho \mid < 1$ and the residuals ε_{it} are serially uncorrelated. The estimate of the half–life defined as the time required for a one unit shock to the real exchange rate to disappear by one half is given by

$$H(\rho) = \frac{\ln(0.5)}{\ln(\rho)}. \tag{2}$$

This expression is, however, not valid for higher order autoregressive processes. In such cases we can compute the half–life from the impulse response function. The half–life is

[3] There are studies looking at price level convergence in Europe and in the euro area, for example Foad (2005) and Rogers (2006). These studies are related to the present study but they do not provide estimates of the persistence of real exchange rates.

then given by the time required for the impulse response of the real exchange rate to a one unit shock to be equal to 1/2.

Estimates of the half–life using the specification above are based on the assumption of homogeneity, i.e., the autoregressive parameter is constant across countries. This may not be the case in empirical applications, therefore we rewrite the AR(1) model above as the following Random Coefficients Model (RCM)

$$q_{it} = \alpha_i + \rho_i q_{it-1} + \varepsilon_{it} \tag{3}$$

where the coefficients α_i and ρ_i have constant means and variance–covariances (Pesaran and Smith (1995), Hsiao and Pesaran (2006)). If there is heterogeneity in the model then a pooled estimate of the autoregressive parameter ρ is upward biased implying an upward bias in the estimated half–life, see Pesaran and Smith (1995) or Imbs, Mumtaz, Ravn and Rey (2005).

Consider now the problem of estimating the mean of α_i and ρ_i, i.e., $\overline{\rho}$ and $\overline{\alpha}$. Let $\theta_i = (\alpha_i, \rho_i)'$ and assume that $\theta_i = \overline{\theta} + \upsilon_i$ where $\mathrm{E}[\upsilon_i] = 0$. Assume that

$$\mathrm{E}\left[\upsilon_i \upsilon_j\right] = \begin{cases} \Delta & \text{if } i = j \\ 0 & \text{if } i \neq j \end{cases}.$$

The BLUE estimator of $\overline{\theta}$ is then the GLS estimator

$$\hat{\overline{\theta}}_{GLS} = \sum_{i=1}^{N} W_i \hat{\theta}_i \tag{4}$$

where the weights W_i are given by

$$W_i = \left\{ \sum_{i=1}^{N} \left[\Delta + \sigma_i^2 \left(Q_i' Q_i \right)^{-1} \right]^{-1} \right\}^{-1} \left[\Delta + \sigma_i^2 \left(Q_i' Q_i \right)^{-1} \right]^{-1} \tag{5}$$

where $Q_i = \left(q_{i,-1}, 1 \right)$ with $q_{i,-1} = (q_{i0}, q_{i1}, \ldots, q_{iT-1})'$ and

$$\hat{\theta}_i = \left(Q_i' Q_i \right)^{-1} Q_i' q_i. \tag{6}$$

The covariance matrix for the GLS estimator is given by

$$\mathrm{Cov}\left(\hat{\overline{\theta}}_{GLS} \right) = \left\{ \sum_{i=1}^{N} \left[\Delta + \sigma_i^2 \left(Q_i' Q_i \right)^{-1} \right]^{-1} \right\}^{-1} \tag{7}$$

where

$$\hat{\Delta} = \frac{1}{N-1} \sum_{i=1}^{N} \left(\hat{\theta}_i - \frac{1}{N} \sum_{j=1}^{N} \hat{\theta}_j \right) \left(\hat{\theta}_i - \frac{1}{N} \sum_{j=1}^{N} \hat{\theta}_j \right)' - \frac{1}{N} \sum_{i=1}^{N} \hat{\sigma}_i^2 \left(Q_i' Q_i \right)^{-1} \tag{8}$$

and

$$\hat{\sigma}_i^2 = \frac{\hat{\varepsilon}_i' \hat{\varepsilon}_i}{T-K} \tag{9}$$

with K equal to the number of explanatory variables.

Pesaran and Smith (1995) have suggested an alternative consistent estimator of $\overline{\theta}$, the so called Mean Group (MG) estimator. Instead of weighting the individual OLS estimators of the parameters, i.e. θ_i, they suggest the simple average

$$\hat{\overline{\theta}}_{MG} = \frac{1}{N}\sum_{i=1}^{N}\hat{\theta}_i \tag{10}$$

with covariance matrix

$$\text{Cov}\left(\hat{\overline{\theta}}_{MG}\right) = \frac{1}{N}\hat{\Delta}. \tag{11}$$

This estimator is asymptotically identical to the RCM estimator of the mean as shown by Hsiao and Pesaran (2006).

Both models and methods outlined above are based on the assumption of no cross–sectional dependence, i.e., the error terms ε_{it} are independent. To account for possible dependence in the error terms we use the method developed by Bai and Ng (2002) as suggested by Pesaran (2006). Assume that the error terms have the following structure

$$\varepsilon_{it} = \lambda_i F_t + \epsilon_{it} \tag{12}$$

where F_t is a vector of common factors, λ_i is the factor loading and ε_{it} is the idiosyncratic component of the errors. We may then rewrite the model in equation (3) as

$$q_{it} = \alpha_i + \rho_i q_{it-1} + \lambda_i F_t + \epsilon_{it}. \tag{13}$$

The problem now is how to specify the common factors. Pesaran (2006) suggests using cross section aggregates to eliminate the effects of the unobserved common factors. Bai and Ng (2002), on the other hand, suggest that the common factors F_t can be constructed from the estimated errors ε_{it} using principal components analysis. We choose to follow the latter approach.

For each cross–section i we have

$$\underline{\varepsilon}_i = F\lambda_i + \underline{\epsilon}_i \tag{14}$$

where $\underline{\varepsilon}_i$ is a $T \times 1$ vector, F is a $T \times k$ matrix with k common factors, λ_i contains the factor loadings and $\underline{\epsilon}_i$ is the idiosyncratic component. Adding all cross–sections in our panel we find that

$$\varepsilon = F\Lambda' + \epsilon \tag{15}$$

where ε now is a $T \times N$ matrix, Λ is a $k \times N$ matrix with factor loadings and ϵ is the $T \times N$ matrix with idiosyncratic components.

To estimate the common factors and to determine the number of factors to include (k), Bai and Ng (2002) suggest the following procedure. First, it is obvious that the number of common factors cannot exceed $\min\{N, T\}$. Second, let $\overline{\Lambda}^k$ be the $\sqrt{N}$ times the eigenvectors of $\varepsilon'\varepsilon$ corresponding to the k largest eigenvalues. The estimator of the common factors is then $\overline{F}^k = \varepsilon\overline{\Lambda}^k/N$ and the rescaled common factors are given by $\hat{F}^k = \overline{F}^k\left(\overline{F}^{k\prime}\overline{F}^k/T\right)^{1/2}$. Third, to estimate the number of factors to include, they suggest six different information criteria. These criteria have different properties and are therefore expected to perform differently in finite samples. However, in our empirical application below, we find that they in almost all cases indicate the same number of factors k. We therefore decide to use the following criteria.

$$IC_{p1} = \ln\left(V\left(k, \hat{F}^k\right)\right) + k\left(\frac{N+T}{NT}\right)\ln\left(\frac{NT}{N+T}\right)$$

where $V\left(k, \hat{F}^k\right) = \frac{1}{N}\sum_{i=1}^{N}\hat{\sigma}_i^2$ and $\hat{\sigma}_i^2 = \hat{\epsilon}_i'\hat{\epsilon}_i$.

Our empirical analysis then consists of the following steps. First, we estimate equation (3), save the residuals and estimate principal components in the errors and use the decision rule as suggested by Bai and Ng. This procedure determines the number of common factors to include in equation (12) above. The second step is to estimate equation (13) which will allow us to find the RCM and MG estimators conditioned on the existence of common factors, i.e., conditioned on cross–sectional dependence.

3 Data

We use domestic producer price index for manufacturing Nace rev.1.1 at the 3–digit level measured in domestic currency for EMU member states over the period 1991:1–2005:11. We focus on 24 manufacturing sectors and total manufacturing for 10 euro area countries: Austria, Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal and Spain. All data series are downloaded from Eurostat's homepage. The sectors are described in the Appendix. The empirical analysis is based on all unique bilateral real exchange rates for each sector. To convert domestic producer prices into common currencies before the introduction of the euro, we use monthly averages of daily nominal exchange rates from IFS.

4 Empirical results

We will start focusing on the EMU period 1999:1 to 2005:11. The first step in our empirical analysis is to test for unit roots in our panels of bilateral real exchange rates. We have used three different test procedures, the Levin and Lin (1993) (see also Levin, Lin and Chu (2002)), the Im, Pesaran and Shin (2003) test and the Moon and Perron (2004) test which is designed for the case where we have cross–sectional dependence in the panel. The results show that we can reject the null of non–stationarity very often, see the tables in the Appendix. Both the Levin–Lin and the Moon–Perron tests reject the null in almost all cases at conventional significance levels. The Im–Pesaran–Shin test, however, rejects the null less frequently, for 18 and 13 cases out of 25, respectively. The overall impression from these tests is that the sectoral real exchange rates do not contain unit roots with, maybe, a few exceptions. Given these results we assume stationarity and continue estimating the half–life using the autoregressive models discussed in the section above.

The estimates of the half–life of sectoral real exchange rates are presented in Table 1. To estimate the confidence bands of the estimated half–life we use a non–parametric boot strap (cross sectional resampling with replacement) with 500 replications. The first impression one gets from the estimates in the table is that the half–life varies considerably across sectors, from very rapid adjustments (around 1 year) to several years. Looking first at the RCM estimates we find that the half–life estimate for only one sector (office machinery and computers) lies in the range of 3 to 5 years as suggested by Rogoff (1996). The majority of half–lives are considerably smaller. For 9 out of 24 sectors, the estimated half–life is less than a year. The MG estimates are very similar but the point estimates of the persistence are very often smaller than the RCM estimates. According to these results, sectoral persistence is much smaller than the aggregate estimate. This interpretation is also supported by the FE estimates where a larger part of the sectors lie in the range of 3 to 5 years persistence. Aggregation across countries, assuming that the persistence of all bilateral real exchange rates are identical for each sector, thus indicate a higher degree of persistence than the heterogeneous estimates. In this respect our results lend support to the Imbs, Mumtaz, Ravn and Rey (2005) argument that aggregation may lead to an upward bias in the estimated degree of persistence.[4]

How reliable are the RCM and MG estimates? They usually indicate a more rapid adjustment of the sectoral real exchange rates than the FE estimate. To test the plausibility of pooling the data we apply a test suggested by Swamy (1971). The basic idea is to compare the RCM and MG estimate of the autoregressive parameter with the individual OLS estimates for each bilateral real exchange rate. The test is computed in the following way

$$\chi^2 = \sum_{i=1}^{N} \left(\theta_i^{OLS} - \hat{\hat{\theta}}_{GLS} \right)' \hat{V}_i^{-1} \left(\theta_i^{OLS} - \hat{\hat{\theta}}_{GLS} \right)$$

where $\hat{V}_i$ is the variance–covariance matrix of unit by unit OLS estimates. The test χ^2–distributed with $2(N-1)$ degrees of freedom. The same test can also be applied to the MG estimate.

The results when testing poolability are shown in Table 2. From this table we find that we can reject the null hypothesis that the individual OLS estimates are equal to the RCM and the MG estimate in 18 out of 25 cases and 17 out of 25 cases at the 10% level, respectively. These results, thus, indicate that it may not be appropriate to assume that

[4]Note that Imbs et.al. argue that aggregation across sectors for one bilateral real exchange rate leads to an upward bias. Our FE estimates aggregate over bilateral real exchange rates for each sector. However, the main argument is the same in both cases.

the individual estimates of the half–life are constant across our sample of bilateral real exchange rates.

The results above, thus, suggest quite rapid adjustments of sectoral real exchange rates, more rapid than the consensus view in the earlier literature. This is interesting and may indicate that the euro may play a role in explaining this result. Therefore, we now turn to the period prior to the introduction of the euro.[5] We use the same methods as above.

Prior to estimating the half–life, we perform unit root tests. The results are shown in the Appendix. Using the Levin–Lin test we almost always reject the null of unit roots. When applying the Im–Pesaran–Shin and the Moon–Perron tests we find that we cannot reject the null very often.[6] These results suggest that there could be a structural break in the degree of persistence. Prior to the introduction of the euro, sectoral real exchange rates contain a unit root whereas they are stationary after the introduction of the euro. One interpretation of these results is that the common currency in Europe has lead to a more rapid adjustment of relative prices, the degree of persistence is much lower under a common currency. A strict interpretation of the unit root tests is then that we cannot continue and estimate the half–life since we cannot apply the regression based methods outlined above when the sectoral real exchange rates are nonstationary. However, we also know that there are problems when using and interpreting unit root tests, in single–time series frameworks as well as in panels. We therefore continue our empirical analysis under the assumption that the sectoral real exchange rates are stationary.

Table 3 shows the estimated half–life of sectoral real exchange rates in the period prior to the introduction of the euro. Confidence intervals around each estimated half–life are constructed using a bootstrap based on cross sectional resampling with 500 replications are also reported in the table. Given our discussion above and the results from the unit root tests it is surprising to note that the point estimates of the degree of persistence is lower than for the more recent period, compare tables 1 and 3. Focusing on the RCM estimates we find that the degree of persistence has fallen after the introduction of the euro in 4 cases out of 25 and is constant in one case. For all other sectoral real exchange rates, the degree of persistence has increased in the most recent period. When also comparing the confidence bands we find that the bands are wider in the earlier sample in 13 out of 21 cases. Thus, even though the point estimates are smaller, the confidence bands are wider in the earlier period. It is therefore, not straightforward to draw strong conclusions regarding changes in the degree of persistence.

We have also performed tests of poolability. Table 4 reports the results of these tests. There is, in general, less support for using models allowing for heterogeneity in this sample. We reject the null of constant parameters across the cross–sections in 10 and 13 out of

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[5] Data are not available for 4 of the sectors we have analyzed above.

[6] Note that we cannot reject the null of a unit root for any sector when we use the Moon–Perron test and only include an intercept.

the 21 cases. Heterogeneity is, therefore, less evident in this period compared to the more recent sample.

Even though we have argued above that it may not be appropriate to compare point estimates of the degree of persistence we now examine whether the estimated half–life has changed over time. One should keep in mind that the confidence bands are much wider in the earlier sample which could bias our interpretations when looking at only point estimates.

In particular, we test whether the mean, the median and the variance are constant across the two samples. To test constancy of the mean we use the Anova F–test, the Wilcoxon signed rank test is used to test for constancy of the median and the Brown–Forsythe test is used to test the null that the variance of the estimated degree of persistence is constant over time.

The results are shown in Table 5. According to these tests, we cannot reject the null that the mean, median or the variance of estimated half–life using the RCM model or the MG estimator. Pairwise comparison of the sectors also reveal that there are very small differences in the estimated half–life. However, turning to the FE estimates we again note a very large difference, we can now always reject the null hypotheses.

Our interpretation of these results together with our results when testing the poolability assumption is that the estimated half–life in the two periods with different exchange rate regimes, on average, do not differ. This is expected since firms may have to change their prices in case a shock affects the sector regardless of the exchange rate regime. Under fixed exchange rates, all real exchange rate adjustments must be carried out by price changes. In a flexible exchange rate regime, the nominal exchange rate might change and given the consensus view in the earlier literature that the half–life of nominal exchange rates is sizable (often exceeding the half–life of relative prices), relative prices adjust more rapid to compensate for the very slow adjustment of the nominal exchange rate. When the exchange rate is fixed, relative prices are more persistent.

These results can be compared to estimated half–lives for other comparable countries, European countries that are not members of the monetary union. Data are available for Sweden and the United Kingdom but unfortunately not for Denmark. Therefore, we focus on the adjustments of sectoral Pound/SEK real exchange rates for the two samples. Data exists for 15 of the 24 sectors we have studied earlier and for total manufacturing.

Given that we only have one real exchange rate for each sector we now rely on univariate estimates of the half–life using the same basic regression as in our earlier analysis. One potential problem when using univariate methods is that least squares estimate of the autoregressive parameter is downward biased for near unit root processes. Several different approaches to handle this potential problem have been suggested in the literature. For example, Stock (1991) show how confidence bands can be constructed for the largest autoregressive root in local–to–unity autoregressive models, Andrews (1993) suggests that confidence bands should be constructed from median–unbiased estimates of the autoregressive parameter, Andrews and Chen (1994) show how median–unbiased estimates of persistence and half–lives can be obtained from general AR(p) processes, Hansen (1999) suggests a grid bootstrap method (which under the assumption of i.i.d. Gaussian errors corresponds to the method suggested by Andrews (1993). We will apply Hansen's method and use the grid bootstrap method. The grid in our application has 300 steps with increment 0.001 starting with an autoregressive parameter equal to 0.8. The number of bootstraps is set to 1999 following Hansen (1999).

In Table 6 we report estimated half–lives for the Pound/SEK exchange rate. The left panel reports results for the earlier pre–EMU period and the right hand panel shows results for the EMU period. Looking first at the point estimates of the half–life during the latter period 1999–2005 and comparing to the results in Table 1 we find that the half–lives are comparable, there seems to be only minor differences. Exceptions are the basic precious and non–ferrous metals sector (dj274) where the estimated half–life is considerably larger and the office machinery and computers sector (dl300) where the half–life is smaller for the Pound/SEK real exchange rate.

For the earlier period we find that the half–life of the Pound/SEK real exchange rates are much larger than the EMU member countries point estimates reported in Table 3 for 10 out of the 13 comparable cases. Contrary to the case of the EMU countries, we find a significant fall in the the estimated half–life of the Pound/SEK rate (for 13 out of the 16 cases). For the EMU countries we found no significant change.

These results must, however, be interpreted with caution since the confidence bands always include infinity suggesting that the sectoral real Pound/SEK exchange rates are nonstationary in both samples. If the Pound/SEK rate is nonstationary, then the estimated half–lives of the sectoral real exchange rates for EMU countries are always smaller implying a more rapid adjustment.

4.1 Industry grouping and persistence

In this subsection we examine whether the differences in the estimated half–life across sectors can be explained by the industry type, i.e., the market structure. We focus on the grouping of industries suggested by Peneder (2001) for the Nace rev.1 categories where he distinguishes between five different types: mainstream, labor–intensive, capital–intensive, marketing–driven and technology–driven industries. In Table 7 we show the mean of the estimated half–life for each industry group for the two samples. The last row in the table reports the p–value of an Anova F–test of the null hypothesis that the mean of estimated half–lives for each industry type is constant across industry types.

From Table 7 we note that there are significant differences across industry groups, the Anova tests always reject the null that the averages are equal across the five groups at the 5 percent level. Technology–driven and labor–intensive industries seem to have the largest half–life whereas capital–intensive industries have the smallest half–life. This result holds for both periods. Note also that we obtain very similar ranking of the industry groups with respect to estimated half–lives for all three statistical methods. It does not matter if we pool the data or if we use heterogeneous panel data methods. Since this ranking is also constant over time, we find that the introduction of the euro has played no role for the ranking of industry groups and the estimated half–lives.

Why do we observe these differences across industry groups? The degree of competition in different sectors may explain the differences in the estimated half–life across these industry groups. It is likely that capital–intensive sectors are more competitive implying more flexible prices and therefore less persistence in real exchange rates. Technology–driven and labor–intensive industries very often operate on markets with monopolistic competition which implies more rigid prices and a higher degree of persistence in these sectoral real exchange rates. If these markets are affected by a shock it may not be necessary for a typical firm to adjust its prices immediately.

These conclusions are supported by estimates of the price–cost margin (PCM) presented in Table 8. We compute PCM as one minus the value added in production minus personnel costs over total value of production in each sector and then we aggregate over sectors belonging to a certain industry type as above. In the table we present the mean, minimum and maximum of these industry specific PCM's. As can be seen from the table there are not much variation, all PCM's are fairly close to unity indicating less than perfect competition across these industry types. However, looking more closely at the estimates and comparing the ranking of industry groups with respect to PCM and the ranking with respect to estimated half–lives, we find an exact match.[7] Technology–driven industries have the highest price–cost margin followed by labor–intensive industries. Capital–intensive and mainstream industries have the lowest PCM's. It is, therefore, likely that the former two groups operate under less competition than the latter two groups of industries.

Our empirical evidence thus suggests that the degree of persistence is not related to the exchange rate regime but to the type of industry. This explanation is consistent with earlier findings in Cheung, Chinn and Fujii (2001) and Chen (2004). Furthermore, Fabiani et.al. (2005) find, using survey data of more than 11000 companies in nine euro area countries, that firms operating in markets that are highly competitive tend to adjust their prices more frequently. Two–thirds of the companies included in the survey use state–dependent pricing rules, i.e., they adjust prices whenever there is a shock affecting the company whereas one–third of the companies use a time–dependent pricing rule, i.e., they adjust their prices on a periodic basis. It is likely that prices in sectors belonging to the latter group are more sticky leading to a higher degree of persistence. Their results also show, however, that there is no clear–cut relationship between the degree of competition and pricing rules.

Exchange–rate pass–through may also play a role in this respect. Campa, Gold-

[7] We only have 22 observations which prohibits a more formal analysis of the relationship using regression analysis.

berg and González–Mínguez (2005) find very little evidence on structural changes in the exchange–rate pass–through in the euro area countries. Point estimates suggest a decline in the pass–through but it is not statistically significant. There is in general a high but incomplete pass–through and they find differences across sectors (1 digit SITC classification) and countries. Therefore, it is not surprising that we find that the degree of persistence is not related to the exchange rate regime. Firms still have to adjust their prices but within a common currency area, they have to change prices more but less rapid compared to a situation where they also are faced with exchange rate changes that often move in the wrong direction in the short–run.

5 Conclusions

This paper examines sectoral real exchange rates in Europe during two distinct exchange rate regimes, the period prior to the introduction of the euro and the period after its introduction. We focus on 24 manufacturing sectors at the 3–digit level (Nace rev.1.1) and total manufacturing in 10 euro area countries. We use econometric panel data methods that allow for heterogeneity at a sectoral level and are corrected for cross–sectional dependence. For comparison we also provide pooled panel data estimates corrected for downward bias using the Nickell correction term.

Starting our analysis by looking at the euro period we find considerable variation in the estimated half–life, from less than a year for some sectors to more than 5 years. In this respect, we find that the Rogoff PPP puzzle can be explained by aggregation bias due to an assumption of homogeneity as suggested by Imbs, Mumtaz, Ravn and Rey (2005). Most estimates of persistence are based on aggregate data, typically, consumer price index with estimated half–life in the range of 3 to 5 years. We seldom find, in our sample of sectors, such slow adjustments of real exchange rates.

Comparing these estimates to the degree of persistence during the period prior to the introduction of the euro we find, surprisingly, that point estimates of the half–life is lower in general. However, we also find wider confidence bands implying that the uncertainty in the estimated degree of persistence is much higher. Formal tests of the equal mean, median and variance suggest, however, that there is no statistically significant difference across the two periods when allowing for heterogeneity.

The documented differences in the estimated half–life can be attributed to the type of industry such that capital–intensive and technology–driven sectors operating under less competition tend to have more rapid adjustment following a shock to the relative price. Capital–intensive sectors operating under more competition tend to have less persistent sectoral real exchange rates. These results hold for both exchange rate regimes and, thus, seems to be independent on whether there is a common currency or not, a result consistent with empirical evidence provided by Campa, Goldberg and González–Mínguez (2005) who show that there is no significant structural break in exchange–rate pass–through in the euro area. Fabiani et.al. (2005) show that price stickiness is higher when the degree of competition is lower. They also find that a high labor–share (labor–intensive industries) is associated with more price stickiness and therefore a higher degree of persistence in real exchange rates.

What can we learn from this study of sectoral real exchange rates. First, we have found that it is very important to allow for heterogeneity when estimating the degree of persistence, there is an aggregation bias not only across sectors as has been shown earlier but also across countries. Sectoral persistence is very often considerably lower than the consensus view of half–lives in the range of 3 to 5 years. Second, there is no statistically significant structural break in the degree of persistence associated with the introduction of the euro, a result consistent with earlier findings of no structural break in exchange–rate pass–through. Third, the degree of persistence is determined by the market structure such that less competition leads to less exchange–rate pass–through and a higher degree of persistence. It may well be that increased intra–industry trade within the euro area will lead to more competition and therefore less persistence in the future.

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Table 1: Estimates of half–life, 1999:1–2005:6.

Sector	N	RCM	MG	FE	Sector	N	RCM	MG	FE
d	45	17	16	20	dg243	21	17	13	30
		(11 27)	(11 25)	(14 35)			(11 66)	(9 26)	(15 138)
da154	15	14	10	47	dg245	15	3	3	3
		(8 315)	(6 24)	(18 ∞)			(2 6)	(1 5)	(1 14)
da155	21	27	19	28	dh251	21	14	11	33
		(14 ∞)	(10 ∞)	(12 ∞)			(9 43)	(7 23)	(18 196)
da158	28	20	16	44	dh252	28	13	10	30
		(13 30)	(10 21)	(27 106)			(9 33)	(7 18)	(18 58)
da159	15	17	14	26	di264	15	24	18	155
		(11 83)	(9 43)	(15 73)			(13 48)	(11 27)	(68 ∞)
db174	15	38	27	47	di265	15	11	9	51
		(18 ∞)	(13 ∞)	(22 ∞)			(6 16)	(5 13)	(28 153)
dc193	10	35	26	33	di266	21	20	15	37
		(21 ∞)	(18 ∞)	(17 186)			(11 111)	(9 40)	(19 ∞)
dd202	15	10	9	8	dj274	21	11	10	21
		(8 ∞)	(7 ∞)	(6 15)			(8 32)	(6 22)	(10 269)
dd203	21	14	10	15	dj286	15	11	9	40
		(8 53)	(6 19)	(9 49)			(6 23)	(5 16)	(17 232)
de211	21	14	12	11	dk297	21	9	7	12
		(11 ∞)	(9 ∞)	(7 21)			(6 23)	(5 14)	(8 25)
de212	21	10	8	24	dl300	15	137	65	136
		(6 25)	(5 15)	(13 62)			(55 ∞)	(34 ∞)	(59 ∞)
df232	21	6	5	8	dn361	21	27	17	55
		(4 9)	(4 7)	(5 15)			(16 ∞)	(11 80)	(33 213)
dg241	15	10	9	15					
		(7 18)	(6 15)	(9 41)					

Note: The first column shows the number of the sector and column two shows the number of bilateral real exchange rates in the panel. RCM denote Random Coefficient Model, MG is the Mean Group estimator and FE is the standard pooled fixed effect model. All three models include a correction for cross–section dependence using the Bai and Ng (2002) method. The number of common factors is determined by information criteria. The fixed effect estimates are also corrected for downward biases using the Nickell (1981) correction term. The confidence intervals shown in parentheses below each estimated half–life are constructed using a bootstrap based on cross–sectional resampling with 500 replications.

Table 2: Test for homogeneity, 1999:1–2005:6.

Sector	N	RCM	MG	Sector	N	RCM	MG
d	45	0.00	0.00	dg243	21	0.01	0.17
da154	15	0.00	0.00	dg245	15	0.00	0.00
da155	21	0.00	0.00	dh251	21	0.12	0.71
da158	28	0.00	0.00	dh252	28	0.00	0.00
da159	15	0.00	0.02	di264	15	0.00	0.00
db174	15	0.49	0.50	di265	15	0.00	0.00
db193	10	0.65	0.71	di266	21	0.00	0.00
dd202	15	0.01	0.04	dj274	21	0.02	0.03
dd203	21	0.00	0.00	dj286	15	0.00	0.01
de211	21	0.65	0.86	dk297	21	0.00	0.00
de212	21	0.00	0.00	dl300	15	0.12	0.44
df232	21	0.02	0.21	dn361	21	0.00	0.00
dg241	15	0.00	0.00				

Notes: N denotes the number of bilateral real exchange rates in the panel, RCM denotes a test of the null hypothesis that the autoregressive parameter in the RCM model is constant whereas MG denotes a test of the null hypothesis that the autoregressive parameter is constant in the MG model. The columns report the p–values.

Table 3: Estimates of half–life, 1991:1–1998:12.

Sector	N	RCM	MG	FE	Sector	N	RCM	MG	FE
d	15	12	9	17	dg245	15	19	14	25
		(7 157)	(6 23)	(8 167)			(11 ∞)	(8 49)	(13 165)
da154	6	8	6	21	dh251	15	9	7	15
		(1 ∞)	(1 ∞)	(11 78)			(5 30)	(4 14)	(8 48)
da155	15	13	11	17	dh252	21	14	11	23
		(8 141)	(7 33)	(9 51)			(9 37)	(7 19)	(13 65)
da158	15	11	9	17	di264	15	33	24	37
		(6 32)	(6 19)	(9 75)			(21 ∞)	(15 101)	(21 128)
da159	10	17	13	20	di265	15	13	11	22
		(10 ∞)	(8 49)	(10 86)			(7 20)	(7 15)	(11 76)
db174	10	17	14	19	di266	15	17	14	28
		(10 ∞)	(8 115)	(9 246)			(10 53)	(8 33)	(14 192)
de211	10	10	9	11	dj274	15	7	6	7
		(6 33)	(6 28)	(6 43)			(4 14)	(4 11)	(5 14)
de212	15	9	8	14	dj286	6	2	2	17
		(6 19)	(5 14)	(8 51)			(0 3)	(0 3)	(9 77)
df232	10	5	4	6	dl300	10	28	19	46
		(3 10)	(3 9)	(4 14)			(15 ∞)	(11 55)	(25 359)
dg241	10	8	8	8	dn361	10	16	12	28
		(5 34)	(5 25)	(5 20)			(8 66)	(7 26)	(12 ∞)
dg243	15	14	11	21					
		(8 ∞)	(6 34)	(10 ∞)					

Note: The first column shows the number of the sector and column two shows the number of bilateral real exchange rates in the panel. RCM denote Random Coefficient Model, MG is the Mean Group estimator and FE is the standard pooled fixed effect model. All three models include a correction for cross–section dependence using the Bai and Ng (2002) method. The number of common factors is determined by information criteria. The fixed effect estimates are also corrected for downward biases using the Nickell (1981) correction term. The confidence intervals shown in parentheses below each estimated half–life are constructed using a bootstrap based on cross–sectional resampling with 500 replications.

Table 4: Test for homogeneity, 1991:1–1998:12.

Sector	N	RCM	MG	Sector	N	RCM	MG
d	15	0.00	0.20	dg245	15	0.79	0.87
da154	6	0.01	0.02	dh251	15	0.00	0.00
da155	15	0.63	0.92	dh252	21	0.00	0.23
da158	15	0.81	0.91	di264	15	0.00	0.01
da159	10	0.00	0.00	di265	15	0.11	0.16
db174	10	0.92	0.81	di266	15	0.79	0.91
de211	10	0.69	0.47	dj274	15	0.76	0.66
de212	15	0.01	0.08	dj286	6	0.00	0.00
df232	10	0.00	0.00	dl300	10	0.00	0.00
dg241	10	0.52	0.31	dn361	10	0.24	0.33
dg243	15	0.03	0.35				

Notes: N denotes the number of bilateral real exchange rates in the panel. RCM is a test of the null hypothesis that the autoregressive parameter in the RCM model is constant whereas MG is a test of the null hypothesis that the autoregressive parameter is constant in the MG model. The columns report the p–values.

Table 5: Testing whether the mean, median and variance changes over time.

	RCM	MG	FE
Mean	0.194	0.158	0.034
Median	0.191	0.206	0.025
Variance	0.304	0.273	0.041

Notes: The test for equal mean is an Anova F–test, median equality test is the Wilcoxon signed rank test, and variance equality test is the Brown–Forsythe test. The table only reports p–values.

Table 6: Estimates of half–life, sectoral Pound/SEK real exchange rates.

Sample: 1991:1–1998:12				Sample: 1999:1–2005:11			
Sector	Half–life	Sector	Half–life	Sector	Half–life	Sector	Half–life
d	74	dg243	19	d	9	dg243	10
	(69 ∞)		(12 ∞)		(6 ∞)		(7 ∞)
da154	133	dh251	27	da154	7	dh251	3
	(157 ∞)		(20 ∞)		(4 ∞)		(3 ∞)
da155	21	dh252	53	da155	18	dh252	12
	(14 ∞)		(49 ∞)		(11 ∞)		(8 ∞)
dd202	8	di266	67	dd202	10	di266	6
	(5 ∞)		(61 ∞)		(6 ∞)		(4 ∞)
dd203	85	dj274	40	dd203	11	dj274	110
	(79 ∞)		(26 ∞)		(7 ∞)		(111 ∞)
de212	50	dk297	18	de212	9	dk297	13
	(51 ∞)		(11 ∞)		(5 ∞)		(8 ∞)
df232	5	dl300	11	df232	4	dl300	26
	(3 ∞)		(6 ∞)		(3 ∞)		(14 ∞)
dg241	6	dn361	80	dg241	5	dn361	12
	(4 ∞)		(84 ∞)		(3 ∞)		(7 ∞)

Notes: All estimated half–lives are based on univariate AR(1) models. Confidence bands reported in parentheses below each estimate are constructed using Hansen's (1999) grid bootstrap method with 300 grids with increment equal to 0.001 and with 1999 replications.

Table 7: Relationship between the estimated half–life and industry type.

	1999:1–2005:11			1991:1–1998:12		
Industry	RCM	MG	FE	RCM	MG	FE
Mainstream	14	11	28	13	10	20
Labor–intensive	23	16	56	22	17	28
Capital–intensive	10	9	21	9	8	11
Marketing–driven	18	14	32	12	9	20
Technology–driven	137	65	136	28	19	46
Anova F–test	0.000	0.000	0.019	0.012	0.029	0.000

Notes: In the table we report estimated mean of the half–life for each industry group. The last row reports the p–value of an Anova test of equal half–life across all industry groups.

Table 8: Price–cost margin in 2002 for industry type.

Industry type	$\overline{PCM}$	$\min(PCM)$	$\max(PCM)$
Mainstream	0.878	0.795	0.928
Labor–intensive	0.890	0.820	0.946
Capital–intensive	0.866	0.712	0.968
Marketing–driven	0.886	0.773	0.938
Technology–driven	0.949	0.878	0.997

Notes: PCM is computed as value added minus personnel costs over total value of production in each country for each sector in 2002.

Appendix

Table A.1: Sectors

NACE rev1. code	Description
d	Total manufacturing
da154	Vegetable and animal oils and fats
da155	Dairy products
da158	Other food products
da159	Beverages
db174	Made-up textile articles, except apparel
dc193	Footwear
dd202	Panels and boards of wood
dd203	Builders' carpentry and joinery
de211	Pulp, paper and paperboard
de212	Articles of paper and paperboard
df232	Refined petroleum products
dg241	Basic chemicals
dg243	Paints, coatings, printing ink
dg245	Detergents, cleaning and polishing, perfumes
dh251	Rubber products
dh252	Plastic products
di264	Bricks, tiles and construction products
di265	Cement, lime and plaster
di266	Articles of concrete, plaster, cement
dj274	Basic precious and non-ferrous metals
dj286	Cutlery, tools and general hardware
dk297	Domestic appliances n.e.c.
dl300	Office machinery and computers
dn361	Furniture

Table A.2: Unit root tests, 1999:1–2005:6.

Sector	N	trend	LL no	LL yes	Z_{tbar} no	Z_{tbar} yes	t^*_a no	t^*_b yes
d	45		9.30	39.93	−1.75	−1.55	2.87	3.08
			0.00	0.00	0.04	0.06	0.00	0.00
da154	15		19.79	34.84	0.72	0.23	−6.49	−3.43
			0.00	0.00	0.24	0.41	0.00	0.00
da155	21		13.00	28.35	3.21	2.91	2.99	2.61
			0.00	0.00	0.00	0.00	0.00	0.00
da158	28		4.58	56.54	1.54	3.97	2.34	2.21
			0.00	0.00	0.06	0.00	0.01	0.01
da159	15		10.95	32.20	1.79	3.31	0.58	0.45
			0.00	0.00	0.04	0.00	0.28	0.33
db174	15		7.16	34.97	2.49	1.04	0.63	0.59
			0.00	0.00	0.01	0.15	0.26	0.28
dc193	10		4.36	21.09	−0.14	2.30	2.20	2.38
			0.00	0.00	0.45	0.01	0.01	0.01
dd202	15		8.20	28.08	−1.76	−0.99	−5.41	−2.86
			0.00	0.00	0.04	0.16	0.00	0.00
dd203	21		7.62	45.76	2.42	1.68	4.52	3.92
			0.00	0.00	0.01	0.05	0.00	0.00
de211	21		11.41	35.57	−3.09	−1.37	−2.32	−1.95
			0.00	0.00	0.00	0.08	0.01	0.03
de212	21		10.00	34.04	−1.76	−1.17	−4.82	−3.09
			0.00	0.00	0.04	0.12	0.00	0.00
df232	21		20.78	47.48	−5.28	−2.39	−15.01	−8.61
			0.00	0.00	0.00	0.01	0.00	0.00
dg241	15		15.33	33.72	−1.15	−0.12	−9.77	−4.72
			0.00	0.00	0.13	0.45	0.00	0.00
dg243	21		19.96	58.41	2.60	0.25	2.19	1.84
			0.00	0.00	0.00	0.40	0.01	0.03
dg245	15		25.01	81.56	1.59	1.76	−23.72	−6.92
			0.00	0.00	0.06	0.04	0.00	0.00
dh251	21		11.19	42.26	2.22	−0.68	−17.53	−8.55
			0.00	0.00	0.01	0.25	0.00	0.00
dh252	28		16.02	38.75	1.62	1.24	−1.79	−1.57
			0.00	0.00	0.05	0.11	0.04	0.06
di264	15		−5.02	13.75	−0.39	0.25	6.83	10.03
			0.00	0.00	0.35	0.40	0.00	0.00
di265	15		1.65	23.96	2.62	1.44	1.73	1.95
			0.05	0.00	0.00	0.08	0.04	0.03

continued on next page

Sector	N	trend	LL no	LL yes	Z_{tbar} no	Z_{tbar} yes	t_a^* no	t_b^* yes
$di266$	21		−0.17	29.75	−0.84	3.16	1.59	1.76
			0.43	0.00	0.20	0.00	0.06	0.04
$dj274$	21		22.97	48.85	1.25	0.44	−5.91	−4.22
			0.00	0.00	0.11	0.33	0.00	0.00
$dj286$	15		12.43	40.78	4.06	1.15	−3.62	−2.12
			0.00	0.00	0.00	0.13	0.00	0.02
$dk297$	21		13.82	36.34	1.02	2.76	−1.81	−1.28
			0.00	0.00	0.15	0.00	0.03	0.10
$dl300$	15		1.92	34.24	7.06	6.68	4.49	9.32
			0.03	0.00	0.00	0.00	0.00	0.00
$dn361$	21		8.39	41.39	4.27	−1.25	5.46	8.55
			0.00	0.00	0.00	0.11	0.00	0.00

Note: LL is the Levin and Lin (1993) (see Levin. Lin and Chu (2002)) test, Z_{tbar} is the Im, Pesaran and Shin (2003) test (number of lags is 8), and t_a^* and t_b^* are Moon and Perron (2004) tests (number of factors is set to 2). All tests include intercepts. P–values are reported below each test–statistic.

Table A.3: Unit root tests, 1991:1–1998:12.

Sector	N	trend	LL no	LL yes	Z_{tbar} no	Z_{tbar} yes	t_a^* no	t_b^* yes
d	15		8.26	17.10	−2.55	1.20	0.03	0.89
			0.00	0.00	0.01	0.12	0.49	0.19
$da154$	6		1.87	12.09	−1.98	0.14	0.03	0.57
			0.03	0.00	0.02	0.45	0.49	0.28
$da155$	15		7.46	31.10	−2.24	0.65	0.03	0.96
			0.00	0.00	0.01	0.26	0.49	0.17
$da158$	15		9.04	24.67	−2.25	0.07	0.01	0.32
			0.00	0.00	0.01	0.47	0.50	0.37
$da159$	10		3.52	14.10	−1.63	1.56	−0.04	−1.06
			0.00	0.00	0.05	0.06	0.49	0.14
$db174$	10		7.78	21.32	−1.86	0.83	−0.00	−0.09
			0.00	0.00	0.03	0.20	0.50	0.47
$de211$	10		7.02	29.92	−2.26	−2.03	−0.05	−0.74
			0.00	0.00	0.01	0.02	0.48	0.23
$de212$	15		10.92	30.20	−3.54	−0.47	0.04	0.63
			0.00	0.00	0.00	0.32	0.49	0.27
$df232$	10		29.72	66.60	0.29	2.18	0.07	1.02
			0.00	0.00	0.39	0.01	0.47	0.16
$dg241$	10		9.83	36.80	−1.63	0.68	0.04	0.94
			0.00	0.00	0.05	0.25	0.48	0.17
$dg243$	15		6.39	19.71	−0.99	1.33	−0.02	−0.69
			0.00	0.00	0.16	0.09	0.49	0.24
$dg245$	15		5.01	16.62	−1.43	0.31	−0.08	−1.76
			0.00	0.00	0.08	0.38	0.47	0.04
$dh251$	15		15.35	35.46	0.19	−0.44	−0.03	−0.68
			0.00	0.00	0.43	0.33	0.49	0.25
$dh252$	21		9.47	22.65	−1.44	0.22	0.02	0.76
			0.00	0.00	0.07	0.41	0.49	0.22
$di265$	15		−1.38	13.60	−1.48	2.95	0.22	2.90
			0.08	0.00	0.07	0.00	0.41	0.00
$di265$	15		7.38	19.76	−1.10	1.33	−0.02	−0.55
			0.00	0.00	0.14	0.09	0.49	0.29
$di266$	15		5.16	16.53	−1.72	2.17	−0.07	−1.41
			0.00	0.00	0.04	0.02	0.47	0.08
$dj274$	15		18.64	35.04	−3.58	−1.49	0.01	0.23
			0.00	0.00	0.00	0.07	0.50	0.41
$dj286$	6		2.14	7.04	−1.41	−0.08	0.06	2.09
			0.02	0.00	0.08	0.47	0.47	0.02

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Sector	N		LL	LL	Z_{tbar}	Z_{tbar}	t_a^*	t_b^*
		trend	no	yes	no	yes	no	yes
$dl300$	10		0.56	15.96	−0.04	2.90	−0.14	−2.68
			0.29	0.00	0.49	0.00	0.44	0.00
$dn361$	10		3.45	10.30	−1.42	1.64	0.03	0.96
			0.00	0.00	0.08	0.05	0.49	0.17

Note: LL is the Levin and Lin (1993) (see Levin. Lin and Chu (2002)) test, Z_{tbar} is the Im, Pesaran and Shin (2003) test (number of lags is 8), and t_a^* and t_b^* are Moon and Perron (2004) tests (number of factors is set to 2). All tests include intercepts. P–values are reported below each test–statistic.