Finnish Economic Papers – Volume 19 – Number 2 – Autumn 2006

IS A LOW-INFLATION ENVIRONMENT ASSOCIATED WITH REDUCED EXCHANGE RATE PASS THROUGH? *

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This paper investigates the extent of pass through from the US dollar exchange rate to consumer prices in the European Union. A relatively new line of empirical research is pursued that considers whether or not the extent of exchange rate pass through is related to the inflationary environment. In contrast to previous empirical studies, recently developed panel data cointegrating techniques to measure long run pass through are employed. While there is evidence that long-run pass through has declined since the 1970s, it actually increased during the early ERM years despite the presence of lower inflation. (JEL: E31, F41, F49)

1. Introduction

The extent of exchange rate pass through (ERPT) from the nominal exchange rate to domestic prices is seen by many as being indicative of the extent of economic and financial interdependence. In terms of monetary policy, high ERPT might be viewed as consistent with low autonomy and high integration. On the other hand, low ERPT might signify high independence and the ability of the central authorities to implement inflation targeting.[1] Several studies have observed that the response of consumer prices to changes in nominal exchange rates has been both incomplete and varied over time. Moreover, recent research has suggested that ERPT declines in response to a credible low inflationary regime (see, *inter alia*, Taylor, 2000; Choudhri and Hakura, 2001; Bailliu and Fujii, 2004). This paper assesses the extent to which long-run ERPT has varied against a changing inflationary environment in the European Union (EU) over the study period 1972–2004. In this study, ERPT is based on the relationship between the nominal US dollar exchange rate and national consumer price indices where high (low) ERPT is seen as being indicative of high (low) interdependence with respect to the US. This paper addresses the central research question of whether or not ERPT in the European Union (EU) has declined over the past thirty years where EU inflation rates have fallen considerably. Evidence of reduced ERPT would lend empirical support to the work of Taylor (2000) and others and suggest that EU monetary policy has become increasingly independent from that of the US.

* *I am grateful to the Editor and two anonymous referees for their helpful and constructive comments. I am also grateful to Peter Pedroni and Donggyu Sul for their advice. Any remaining errors are my own.*

[1] *Pass-through issues have also played a central role in debates over appropriate monetary policies and exchange rate regime optimality (see, for example, Smets and Wouters, 2002).*

In contrast to much of the existing literature, it is the extent of *long-run*, as opposed to *short-run*, ERPT that is the main focus of attention in this study. Long-run ERPT is measured by the relationship between the nominal and real US dollar exchange rates using cointegrating techniques across different inflationary regimes. However, a significant contribution to the literature from this study is the application of panel data cointegrating techniques. Rather than relying upon Engle-Granger and Johansen cointegration tests which can suffer from power deficiency under conditions of limited time series, this is the first study that employs nominal and real exchange rate data for a large sample of EU countries and tests the null hypothesis of non-cointegration within a panel framework. This novel approach, recently introduced and developed by Pedroni (1999, 2001, and 2004) uses more observations and exploits the cross-country variations of the data in estimation thereby yielding higher test power than alternative unit root and cointegration tests. In addition, this study estimates the long-run cointegrating panels using dynamic ordinary least squares (DOLS) which provides estimates of long-run ERPT and facilitates tests of restrictions.

The paper is organised as follows. The following section discusses the ERPT literature and the panel data cointegration methodology followed in this paper. Using this methodology, the following questions are addressed. First, is there evidence that long-run ERPT in the EU has varied over the last thirty years or so? Second, if ERPT has varied, is there evidence that the reduced ERPT is associated with low inflationary regimes? The third section describes the data. The full study period of April 1972 to June 2004 is divided into sub-periods that reflect the credibility of EU inflationary regimes based on the rate of inflation and underlying exchange rate arrangements that prevailed for EU members. The fourth section reports and discusses the results. Evidence of complete ERPT is lacking throughout. In addition to this, we find that fluctuations in the ERPT coefficient do not provide the expected match with the corresponding inflationary performance of EU members. The final section concludes.

2. Literature and methodology

The traditional ERPT literature is concerned with pass through from the exchange rate to import prices stressing the role of market power and price discrimination in international markets (pricing to market). Goldberg and Knetter (1997) survey a literature which argues that import price pass through is essentially determined by microeconomic factors such as demand elasticities and market structure. Much of this work builds almost exclusively on the concept of market segmentation where the discussion is set in an oligopolistic framework with imperfect competition and third degree price discrimination.[2] While debate over the nature of pass through has concerned the prevalence of producer-currency versus local-currency pricing of imports, the behavior of import prices ultimately feeds through into consumer prices.

An interesting direction in the literature has recently argued that the extent of ERPT has declined with the transition towards a credible low inflationary environment. Against a background of staggered price setting and monopolistic competition, Taylor (2000) presents a dynamic general equilibrium open economy model where firms set prices for several periods in advance, but their prices respond more to price increases (due to exchange rate depreciation or other sources) if cost changes are perceived to be more persistent. Regimes with higher inflation appear to have more persistent costs and will therefore increase the extent of ERPT with firms experiencing greater pricing power. Referring to the low inflation episode of the 1990s, Taylor (2000) argues that the persistence of inflation declined in the US. This is because firms expected changes in costs or prices to be less per-

[2] *Given segmented markets, the seminal papers of Krugman (1987) and Dornbusch (1987) spawned oligopolistic models based on the variations in mark-ups in response to exchange rate changes. Another line taken in the literature has its roots in Baldwin (1988), Dixit (1989) and Baldwin and Krugman (1989). This line of literature emphasises hysteretic effects arising from the sunk costs of entering a market that firms cannot recoup when they leave the market. A third direction in the literature concentrates on institutional settings such as the effects of non-tariff barriers or the role of multinational corporations and intra-firm trade (see Menon, 1995, 1996).*

sistent therefore fewer exchange rate changes were passed through to domestic prices. Choudhri and Hakura (2001) emphasise a similar channel to Taylor (2000) where ERPT is reflected in the expected effects of monetary shocks on current and future costs, while Devereux and Yetman (2003) argue that the frequency of price changes of importing firms declines with the degree of credibility attached to monetary policy. Finally, Devereux and Yetman (2005) develop a theoretical framework to show how sticky prices represent a key determinant of exchange rate pass through to consumer prices.

A number of empirical studies indicate a reduction in ERPT. For example, event studies by Cunningham and Haldane (1999) of the 1992 depreciation and 1996 appreciation in the UK, the 1992 depreciation in Sweden, and the 1999 depreciation in Brazil show a remarkably small pass through of exchange rate changes to retail prices. Using a cross-sectional approach applied to various samples of countries, Choudhri and Hakura (2001) and Devereux and Yetman (2003) conduct a two-stage methodological approach. In the first stage, the ERPT coefficient is estimated for each country using time-series data. The second stage entails regressing the ERPT coefficients against explanatory variables that include inflation. Using data for the Bretton Woods period, these studies find that inflation significantly explains the differences in the ERPT coefficients. Gagnon and Ihrig (2002) use a similar two-stage approach for industrialised countries over the period 1971–2000. However, they subdivide their study period on the basis of inflationary experience and find that ERPT declined in the regime of low inflation. In most candidate countries the high inflation environment of the early 1990s gradually changed into a single-digit inflation rate episode. Consequently, we might expect this development to influence the pass-through relationship. McCarthy (2000) uses a vector autoregression model to show a decline in exchange rate pass through to consumer prices for all nine of the OECD countries examined in the period 1983–98 compared with the period 1976–82. According to these estimates, the pass through declined by 50% or more in the US, UK, France, and Japan and by a smaller amount in Germany, Belgium, Netherlands, Sweden, and Switzerland. Finally, Bailliu and Fujii (2004) employ a generalised method of moments panel data approach rather than cross sectional approach. Using data for eleven industrialised countries over the study period 1977–2001, they confirm the positive relationship between ERPT and inflation.

In this study, an alternative methodological approach is taken. This is the first study that investigates the relationship between the inflationary environment and long-run pass through from the US dollar exchange rate to EU consumer prices within a panel data cointegration framework. In contrast to much of the existing literature, this study investigates the long-run impact of nominal exchange rate movements on the real exchange rate. An assessment of how ERPT has altered over different inflationary regimes requires the use of sub-periods that are based on short spans of time series data with the possibility that the null hypothesis of non-cointegration might be accepted on the grounds of low test power. Pedroni (1999, 2004) proposes and discusses a number of statistics that can be used to determine the presence of cointegration in heterogeneous panels. These tests use more observations and exploit the cross-country variations of the data in estimation thereby yielding higher test power than alternative unit root and cointegration tests.[3] Consider the following relationship between domestic consumer prices, the nominal exchange rate and foreign consumer prices

$$(1) \quad p_{it} = \beta_{0i} + \beta_{1i} e_{it} + \beta_{2i} p_t^f + \mu_{it}$$

where there are $i = 1, 2, \ldots, N$ countries and $t = 1, 2, \ldots, T$ time periods, p_{it} is the natural logarithm of the consumer price index of country i, e_{it} is the natural logarithm of the nominal exchange rate (domestic price of foreign currency)

[3] In a recent contribution, Hatemi-j and Irandoust (2004) employ panel data cointegration techniques to examine how Swedish import prices across industries react to exchange rate changes. They confirm less than complete pass through. This study differs from Hatemi-j and Irandoust (2004) in three crucial ways. First, we consider the role of the inflationary environment. Second, we focus on a wide sample of EU countries. Third, we examine pass through with respect to consumer prices rather than import prices.

of country i, p_t^f is the natural logarithm of the foreign price index,[4] β_{0i} allows the cointegrating regression to include country-specific fixed effects, β_{1i} captures the degree of ERPT from e to p where $\beta_{1i} = 1$ indicates complete ERPT and $0 < \beta_{1i} < 1$ indicates incomplete ERPT. Likewise, $0 \leq \beta_{2i} \leq 1$ measures pass through from p^f to p. Following Monacelli (2005), a natural benchmark is where PPP holds exactly and there is complete pass through. Adding $-p_t^f$ and $-e_{it}$ to both sides of equation (1) yields

$$(2) \quad r_{it} = \beta_{0i} + \varphi_{1i} e_{it} + \varphi_{2i} p_t^f + \mu_{it}$$

where $r_{it} = (p_{it} - p_t^f - e_{it})$ is the real exchange rate for country i, while $\varphi_{1i} = (\beta_{1i} - 1)$ and $\varphi_{2i} = (\beta_{2i} - 1)$ measure the deviation from complete pass through stemming from e_{it} and p_t^f respectively. Equation (2) may be rewritten in vector form as follows

$$(3) \quad r_i = x_i \beta_i + \mu_i$$

where $x_i = \left(1, e_{it}, p_t^f\right)$ and $\beta_i = \left(\beta_{0i}, \varphi_{1i}, \varphi_{2i}\right)'$.

A necessary but not sufficient condition for long-run ERPT is that r_{it}, e_{it} and p_t^f are cointegrated. The procedure for computing the test statistics for panel data cointegration involves estimating the hypothesized cointegration regression described in (3) and using the residuals to estimate the appropriate autoregression. Pedroni advocates two statistics both based on a group-mean approach. *Group PP* is non-parametric and analogous to the Phillips-Perron t statistic and *Group ADF* is a parametric statistic and analogous to the ADF t statistic.[5] These two statistics are referred to as *between-dimension* statistics that average the estimated autoregressive coefficients for each country. Under the alternative hypothesis of cointegration, the autoregressive coefficient is allowed to vary across countries. This allows one to model an additional source of potential heterogeneity across countries.[6] Following an appropriate standardization, both of these statistics tend to a standard normal distribution as N, $T \to \infty$ diverging to negative infinity under the alternative hypothesis and consequently, the left tail of the normal distribution is used to reject the null hypothesis of non-cointegration.

The above panel procedures test the null hypothesis that all members of the panel are not cointegrated against the alternative that at least one member is cointegrated. A potential drawback with such a testing procedure is that one cannot determine which and how many panel members are responsible for any rejection of the null (Maddala and Kim, 1998). Drine and Rault (2005) offer an alternative perspective on these tests by pointing out that each member of the panel represents a draw from an underlying population where the panel simply represents a repeated sampling (N times) from an underlying population. In this case, they argue that the population data generation process (DGP) either is cointegrated or is not cointegrated. As the number of individuals of the panel is increased, one is simply accumulating information on whether or not the population DGP is cointegrated or is not cointegrated. In this case, the proper interpretation of the panel data cointegration test is in terms of a null hypothesis stipulating that the DGP is not cointegrated, and an alternative hypothesis stipulating that the DGP is cointegrated. This translates, for the panel, into the statement based on a null hypothesis where no individuals are cointegrated, and an alternative hypothesis that all individuals are cointegrated.

Following Pedroni (2001), a DOLS procedure can be employed to obtain the panel data estimates for φ_1 and φ_2. DOLS estimation involves augmenting the cointegrating regression with lead and lagged differences of the regressors to control for endogenous feedback effects. The DOLS regression for the i^{th} panel member may be written as

[4] *Devereux and Yetman (2003) employ the foreign (US) price level in their short-run equation used for the analysis of ERPT.*

[5] *This latter statistic is analogous to the Im, Pesaran and Shin (2003) test for a panel unit root applied to the estimated residuals of a cointegrating regression.*

[6] *Pedroni also proposes four* within-dimension *statistics (*panel v, panel ρ, panel t *and* panel ADF*) that effectively pool the autoregressive coefficients across different countries during the unit root tests. In these tests, a common value for the autoregressive coefficient is specified under the alternative hypothesis of cointegration.*

(4) $\tilde{r}_i = z_i \varsigma_i + \mu_i^*$

where $z_i = (e_{it} - \bar{e}_i, p_t^f - \bar{p}^f, \Delta e_{it-k}, ..., \Delta e_{it+k}, \Delta p_{t-k}^f, ..., \Delta p_{t+k}^f)$ contains both lead and lagged differences of the regressors, ς_i is a vector that contains the DOLS estimators for φ_{1i} and φ_{2i} along with coefficients on the lead and lagged differences of the regressors, and $\tilde{r}_i$ is the vector based on $r_{it}–\bar{r}_i$. The DOLS estimator for the i^{th} member of the panel may be written as

(5) $\hat{\varsigma}_{D,i} = (z_i' z_i)^{-1} (z_i' \tilde{r}_i)$

From this regression, we may construct the between-dimension group-mean panel DOLS estimator as

(6) $\hat{\varsigma}_{GD}^* = N^{-1} \left[\sum_{i=1}^{N} (z_i' z_i)^{-1} (z_i' \tilde{r}_i) \right]$

where the vector $\hat{\varsigma}_{GD}^*$ contains the group mean panel DOLS estimates $\hat{\varphi}_{1,GD}^*$ and $\hat{\varphi}_{2,GD}^*$ that respectively indicate the extent deviation from complete pass through stemming from e_t and p_t^f. If the long-run variance of the residuals from the DOLS regression for each country is given as $\sigma_i^2 = \lim_{T \to \infty} E \left[T^{-1} \left(\sum_{t=1}^{T} \hat{\mu}_{it}^* \right)^2 \right]$, Pedroni argues that the associated t-statistic for this estimator may be constructed as

(7) $t_{\hat{\varsigma}_{GD}^*} = N^{-0.5} \sum_{i=1}^{N} t_{\hat{\varsigma}_{D,i}}$

where $t_{\hat{\varsigma}_{D,i}} = (\hat{\varsigma}_{D,i} - \varsigma_0) \left(\hat{\sigma}_i^{-2} \sum_{t=1}^{T} (e_{it} - \bar{e}_i)^2 \right)^{0.5}$. Hypothesis tests involving these t-statistics are asymptotically normal.[7]

[7] *Pedroni (2000) discusses the asymptotic properties of the group mean fully modified OLS (FMOLS) estimator and shows how asymptotic normality is applicable. It should be noted that DOLS estimators always have the same asymptotic properties as fully modified OLS estimators. This is because the only difference between them is in terms of how the nuisance parameters are treated (parametrically versus non-parametrically). Therefore, the asymptotic properties for the between-dimension group mean DOLS estimator are identical to those of the between-dimension group mean FMOLS estimator.*

In the empirical analysis, we focus on between-dimension panel DOLS tests. In doing this, there are several advantages over a within-dimension approach. First, the between-dimension approach allows for greater flexibility in the presence of heterogeneity across the cointegrating vectors where φ_{1i} and φ_{2i} are allowed to vary. Under the within-dimension approach, φ_{1i} and φ_{2i} would be constrained to be the same value for each country under the null hypothesis. Second, the point estimates of the between-dimension estimator can be interpreted as the mean value of the cointegrating vectors. This is helpful in interpreting the results. Third, the between-dimension estimator suffers from lower small-sample size distortions than is the case with the within-dimension estimator.

3. Data

This study uses monthly, end of month data for consumer prices and nominal spot exchange rates with respect to the US dollar over the study period April 1972 to June 2004. Twelve EU countries are included in this study namely, Belgium, Denmark, Finland, France, Germany, Greece, Italy, the Netherlands, Portugal, Spain, Sweden and the UK. All data are obtained from the OECD *Main Economic Indicators*.[8] The sample of countries incorporates a range of stances regarding membership of the EU and the associated exchange rate regime. We might think in terms of a "core" EU group comprising Belgium, France, Germany and the Netherlands who have been members of the EU from the outset and have the strongest record of ERM membership. In addition to this, we might also think of an "outer-core" comprising Denmark, Finland, Greece, Italy, Spain, Portugal, Sweden and the UK. While Italy was a member of the EU from the outset, each of these countries have been characterised by currencies with varied experiences of ERM membership. For example, Italy and the UK were ejected from the ERM in September 1992 while Portugal and Spain took part at the wider ±6% bands of exchange rate

[8] *All estimation is conducted using the software package RATS version 6 by Estima.*

fluctuations having respectively joined the ERM in 1992 and 1989. Denmark, while linking to the DM, has opted out of EMU. Greece, on the other hand, remained outside the ERM during the study period but has now proceeded towards EMU. Of all the sample of twelve countries, Denmark, Sweden and the UK have opted not to participate in single currency membership.

The sample period is divided into four sub-periods defined according to EU inflationary environments that are marked by observed policy changes. In these regimes, 'credibility' is broadly reflected in the prevailing exchange rate regime motivated, in part, by the desire of EU members to enforce some degree of inflation discipline.[9] *Period 1* covers April 1972 to February 1979 which is characterised by floating exchange rates with respect to the US dollar following the breakdown of Bretton Woods and the creation of the European Snake that sought to stabilise bilateral exchange rates between any two European countries through the use of declared fluctuation margins vis-à-vis the US dollar. *Period 2* covers March 1979 to April 1986. This is the initial ERM period up to the month prior to the major realignment of ERM currencies in 1986.[10] *Period 3* covers May 1986 to April 1990. This sub-period includes the events leading up to the abolition of capital controls for most EU members by May 1990 and German unification in July 1990. *Period 4* covers August 1993 to June 2004. This sub-period follows the relaxation of capital controls for most EU members by May 1990,[11] the creation of the single market in 1992, the exchange rate crisis during 1992 and the subsequent widening of the permitted bands of exchange rate fluctuation to ±15%. This period also encompasses the enlargement of the EU in 1995, the preparatory period for the single currency as the majority of EU members adhered to the Maastricht convergence criteria, the introduction of the single currency in January 1999 and the use of the stability pact. In this final sub-period, any autonomous monetary policy on the part of individual members of the Euro area has been replaced by monetary control through the European Central Bank.

On the basis of these four sub-periods, Table 1 reports calculations for the mean annual inflation rates for the full sample of countries. It is clear there has been a steady decline in the mean inflation from 11.094% to 2.480% over the full study period. Dividing the sample into sub-periods is a simple and direct way of analyzing the changing nature of ERPT. A general caveat worthy of consideration, which is not addressed in the general pass-through literature, is the extent to which the unit root and cointegration properties of the series may be sensitive with respect to how the study period is split. While this study offers justification for the particular sub-periods that are employed, one could argue that dividing the sample in a variety of alternative or more arbitrary ways might lead the researcher to conclude that the time series properties of the real and nominal exchange rate move between non-stationarity and stationarity over the entire study period. Potentially, this might result in unstable cointegration with highly variable coefficient estimates.[12]

4. Results

The possibility of panel data cointegration involves the relationship between non-stationary panels of consumer prices and exchange rates so the first stage of the empirical investigation is to employ panel data unit root testing. The

[9] *Chang and Lapan (2003) present results that imply exchange rate variability affects pass through. They consider incentives to commit price or retain price flexibility in a model in which exporting firms face different degrees of exchange rate uncertainty.*

[10] *On the basis of the currency realignments, Caporale and Pittis (1993) justify the use of May 1986 as a significant break point in their study of inflation convergence.*

[11] *The UK abolished capital controls in October 1979, Germany had lifted all restrictions on capital inflows by 1981. Although ERM members from its outset, Belgium, France and Italy, used exchange controls to protect their exchange rates. Other EU members were also permitted to maintain controls throughout this sub-period. See Kenen (1995) for a description of how these controls worked. Belgium, France and Italy removed their remaining controls on March 2 1990, January 1 1990 and May 14 1990 respectively.*

[12] *One way forward here might be to appeal to the procedure advocated by Bacciochi and Fanelli (2005) that tests for cointegration in the presence of I(2) stochastic trends.*

Table 1. EU Annual Inflation Rates, April 1972 to June 2004

	April 1972 to February 1979	March 1979 to April 1986	May 1986 to April 1990	August 1993 to June 2004
Belgium	7.990	6.266	1.853	1.807
Denmark	9.671	8.189	4.196	2.138
Finland	11.615	8.081	4.862	1.358
France	9.276	9.524	2.967	1.553
Germany	4.961	3.894	1.2	1.605
Greece	13.338	19.294	14.8	5.434
Italy	13.365	13.642	5.266	2.929
Netherlands	7.459	4.255	0.414	2.427
Portugal	18.399	19.764	10.287	3.419
Spain	15.272	12.137	6.079	3.194
Sweden	8.806	8.868	5.366	1.391
UK	12.977	8.901	5.222	2.509
Mean	**11.094**	**10.235**	**5.209**	**2.480**

Notes for Table 1. These are annual inflation rates based on national consumer prices indices.

Table 2. Panel Data Unit Root Tests

	April 1972 to February 1979	March 1979 to April 1986	May 1986 to April 1990	August 1993 to June 2004
r	–1.178	–0.494	–0.426	–0.591
e	–0.238	–0.150	–0.973	–0.564
Δr	–24.705***	–23.979***	–13.843***	–28.943***
Δe	–15.736***	–12.462***	–10.691***	–24.462***

Notes for Table 2. These are panel data unit root tests advocated by Im *et al.* (2003). These statistics tend to a standard normal distribution as N, $T \to \infty$. *** denotes rejection of the null of joint non-stationarity at the 1% significance level with a critical value of –2.33.

motivation behind the employment of panel data unit roots is the employment of more observations and exploitation of the cross-country variations of the data in estimation thereby yielding higher test power than standard unit root tests based on individual time series. For this purpose, Im *et al.* (2003) tests are employed. Essentially, this test statistic is an average ADF statistic based on demeaned data where each individual series is expressed in relation to the cross-sectional mean. Moreover, for a given series q_{it}, the Im *et al.* (2003) panel data unit root tests are based on the following regression for each panel member

$$(8) \quad \Delta\tilde{q}_{it} = \tilde{\alpha}_i + \tilde{\phi}_i\tilde{q}_{i,t-1} + \tilde{\xi}_{it}$$

where $\tilde{q}_{it} = q_{it} - N^{-1}\sum_{i=1}^{N} q_{it}$. The test statistic is computed as the average t statistic attached to $\tilde{\phi}_i$ where demeaning the data deals with common shocks to the individual series. The null hypothesis is expressed as $H_0 : \tilde{\phi}_i = 0 \ \forall \ i$ and the alternative hypothesis is written as $H_1 : \tilde{\phi}_i < 0$, $i = 1, 2, \ldots, N_1$, $\tilde{\phi}_i = 0$, $i = N_1 + 1, N_1 + 2, \ldots, N$. The formulation of the alternative hypothesis allows for ϕ_i to differ across the series and is less restrictive than earlier panel data unit root tests that define $H_1 : \phi_i = \phi < 0 \ \forall \ i$.[13] This is a one-tail test where the null hypothesis is joint non-stationarity of all the individual series in the panel while the alternative hypothesis is that at least one of the individual series is stationary. The test statistic tends to a standard normal distribution as N, $T \to \infty$.

Table 2 reports the findings from the panel data unit root tests. At the 5% significance level, the null hypothesis of joint non-stationarity is accepted for the panels comprising all real ex-

[13] *See, for example, Levin and Lin (1993).*

Table 3. Panel Data Cointegration Tests

	April 1972 to February 1979	March 1979 to April 1986	May 1986 to April 1990	August 1993 to June 2004
Group PP	–3.295***	–1.746**	–1.787**	–1.332*
Group ADF	–3.168***	–2.672***	–2.444**	–2.956***

Notes for Table 3. These are the Pedroni tests for panel cointegration (discussed in Pedroni 1999, (2004) between each EU:US dollar real exchange rate , the nominal EU:US dollar exchange rate and the US consumer price index. All variables are in natural logarithm form. These estimates include common time dummies. Individual lag lengths are based on the Akaike information criterion. These statistics tend to a standard normal distribution as N, $T \to \infty$. ***, ** and * denote rejection of the null of non-cointegration at the 1, 5 and 10% significance levels critical values of –2.33, –1.64 and –1.28 respectively.

Table 4. DOLS Group Mean Estimation

	April 1972 to February 1979	March 1979 to April 1986	May 1986 to April 1990	August 1993 to June 2004
$\hat{\varphi}^*_{1,GD}$	–0.374	–0.318	–0.606	–0.781
	(–27.807)	(–15.209)	(–48.231)	(–4.918)
	[47.346]	[–35.626]	[9.639]	[6.604]
$\hat{\varphi}^*_{2,GD}$	–0.201	–0.470	–0.459	–0.137
	(–1.754)	(–3.142)	(–3.041)	(–11.494)
	[15.231]	[12.499]	[23.039]	[37.804]

Notes for Table 4. This table reports dynamic OLS (DOLS) panel data group mean estimates $\hat{\varphi}^*_{1,GD}$ and $\hat{\varphi}^*_{2,GD}$ using the Pedroni panel data cointegration methodology. These estimates include common time dummies. Each slope estimate is accompanied by two t-statistics. The t-statistics in parentheses are based on the null of a zero slope. The t-statistics in square brackets are based on the null of a unity slope. All t-statistics tend to a standard normal distribution as N, $T \to \infty$.

change rates and then all nominal exchange rates across the four sub-periods. The acceptance of the null with respect to the real exchange rate panel suggests that PPP does not hold among EU countries and so complete pass through is absent. Table 2 also reports panel data unit tests on the first differences in the exchange rates and the first differences in the nominal exchange rates. In these cases, we can see that the non-stationary null is rejected at the 5% significance level or better.

It is possible that r, e and p^f form a cointegrating relationship characterised by less than complete pass through. Table 3 reports that both the *Group PP* and *Group ADF* panel cointegration tests are able to reject non-cointegration for the EU panel across all the sub-periods. Rejection of non-cointegration implies a long-run relationship between the real exchange rate, nominal exchange rate and foreign consumer prices. The degree of long-run pass through is indicated by the estimates $\hat{\varphi}^*_{1,GD}$ and $\hat{\varphi}^*_{2,GD}$.[14] Table 4 reports the DOLS group mean estimates of the cointegrating relationships for the cases where cointegration is confirmed. If one considers ERPT to begin with, each $\hat{\varphi}^*_{1,GD}$ is significantly different from both zero and unity at the 1% significance level,. This is indicative of incomplete long-run ERPT across the study period.

In the case of the April 1972 to February 1979 sub-period, the group estimate yields a slope value of $\hat{\varphi}^*_{1,GD} = –0.374$ for DOLS estimation. This may be contrasted with the final sub-period of September 1992 to June 2004 which is characterised by a long-run ERPT estimate of

[14] *This result may also be seen in the context of the purchasing power parity (PPP) literature that fails to find cointegration between domestic and foreign prices and the nominal exchange rate. It should be noted that the finding of cointegration in his study is based on the estimation of equation (1) with unrestricted slopes. In addition to this, the enhanced test power offered by these panel data tests makes rejection of the non-cointegration null more likely.*

$\hat{\varphi}^*_{1,GD}$ = –0.781. With an estimated coefficient that is closer to minus unity, this is indicative of much lower ERPT and is accompanied by an average inflation rate of 2.480%. As pointed out by Taylor (2000), in the case of the United Kingdom, neither the 20% depreciation in 1992 nor the 20% appreciation in 1996 caused retail price inflation to deviate noticeably from the 2.5% trend. The same is true for the 1992 depreciation in Sweden. With regard to the estimate for the final sub-period, the null H_0: $\hat{\varphi}^*_{1,GD}$ = –0.374 drawn from the first sub-period result is strongly rejected with a Wald statistic of 24.590. This confirms that the final sub-period is characterised by the lowest slope estimate. The low ERPT coefficient for this period occurs against a background of several factors that enhanced the credibility of a low inflation regime. The Maastricht treaty stipulated strict guidelines on permitting fluctuations in domestic inflation rates,[15] the creation of the single market in 1992 removed restrictions on international trade within the European Union, the introduction of the single currency in January 1999 removed the scope for autonomous monetary policy, and the introduction of the European Central Bank with inflation targeting all served to enhance the credibility behind the low inflationary environment. This decrease in pass through with respect to the US dollar signifies an increase in EU monetary integration at the expense of a decline in monetary interdependence with the US.

The analysis so far points to a decline in ERPT when comparing the first and final subperiods. The finding of cointegration for subperiods based on the early ERM years enables us to reflect on how ERPT changed in the 1980s. Whereas the April 1972 to February 1979 sub-period features $\hat{\varphi}^*_{1,GD}$ = –0.374, this estimate increases to $\hat{\varphi}^*_{1,GD}$ = –0.318 in the following ERM sub-period of March 1979 to April 1986. Indeed, there is a significant difference in the two slope estimates. In the case of the result for the first sub-period, the null H_0: $\hat{\varphi}^*_{1,GD}$ = –0.318 is rejected with a Wald statistic of 4.178.

Despite the reduction in average inflation from 11.094 to 10.235% reported in Table 1, the extent of ERPT actually increased during the years that followed the inception of the ERM. At first sight, this result runs contrary to the theoretical arguments advanced by Taylor (2000) and others. However, two considerations should be borne in mind. First, although inflation was lower on average than during the Snake era of the 1970s, one might question the extent of lower inflation credibility during the March 1979 to April 1986 sub-period. As pointed out by Bailliu and Fujii (2004), there is the possibility that the changes in the monetary policy regime experienced in the 1980s had less credibility than those changes that took place during other periods. Moreover, Artis and Taylor (1988) find evidence of stabilised nominal exchange rates during the early years of the ERM. However, although the permitted fluctuations in exchange rates were set at ± 2.25% around a central parity, there were several realignments within the ERM.[16]

While the main focus of this paper is in terms of investigating ERPT, the estimates reported in Table 4 also throw light on pass through with respect to foreign prices. These findings here are important insofar as they suggest that the recent theories relating pass through to low inflation only have support with respect to exchange rate pass through. Moreover, $\hat{\varphi}^*_{2,GD}$ measures the deviation from complete pass through stemming from p^f_t. Across all sub-periods, the null hypothesis H_0: $\hat{\varphi}^*_{2,GD}$ = –1 is rejected at the 1% significance level thereby indicating a significant degree of direct pass through from foreign prices to domestic consumer prices. However, at the 5% significance level, the null H_0: $\hat{\varphi}^*_{2,GD}$ = 0 is accepted in the cases of the April 1972 to February 1979 and August 1993 to June 2004 sub-periods indicating the pres-

[15] *The Maastricht Treaty stated that individual members' inflation rates should be no more than 1.5% higher than the average of the three lowest rates in the European Monetary System.*

[16] *See Artis (1990) for an account of the early realignments in the ERM. Italy was allowed a larger band fluctuation of ± 6% until the narrower band was followed during 1990–92. The UK was a formal member of the ERM for only 1990–92 though shadowed the German Mark during 1987–88. During its brief membership, the UK also adhered to the ± 6% limit.*

ence of complete pass through. The results indicate that for the first and final sub-periods, the main source of pass through is foreign prices rather than the exchange rate.

5. *Summary and conclusion*

Using monthly data for a sample of twelve European Union countries over a study period spanning 1972–2004, the employment of panel data cointegrating techniques reveals that the extent of exchange rate pass through from the US dollar to European Union consumer prices has declined. In general terms, this decline has occurred against a background of reduced inflation as progression was made towards the introduction of the single currency in the 1990s. In the latter part of the study period, changes in exchange rates have had surprisingly small effects on consumer prices even in small open economies where imported products are a large fraction of final consumption and intermediate inputs to production. In this sense, the degree of pass through varies positively with the degree of inflation. However, the early 1980s saw the extent of pass through actually increase despite the inception of the Exchange Rate Mechanism and reduction in average inflation rates as compared to a decade earlier. This suggests that it is not just lower inflation *per se* that will reduce the extent of pass through, but rather lower inflation combined with a credible regime of inflationary or monetary control. The fairly modest reduction in pass through experienced in the late 1980s may be seen against a background of ongoing limited credibility behind the lower inflation rates that were experienced. In the 1990s, the convergence criteria of the Maastricht Treaty, progression towards and introduction of the single currency and use of inflation targeting along with an independent European Central Bank served to enhance the credibility behind the low average rate of inflation in the European Union. Accordingly, the extent of pass through in this latter sub-period is very low.

The findings from this study point us towards the following avenues for future research. First, the importance of credibility behind low inflation rates suggests that theoretical models of firm pricing behaviour might be more closely and explicitly related to the general macroeconomic environment in which firms operate. Second, the addition of ten new European Union members in May 2004 offers the opportunity to examine exchange rate pass through for a wider sample as these countries operate within their new targets for exchange rate and monetary control.

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