Exchange Rate Pass-Through to Consumer Prices: A European Perspective

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Der vorliegende Artikel untersucht die Auswirkungen von Wechselkursveränderungen auf die Verbraucherpreise im Euro-Raum durch Schätzung von Vektor-Fehlerkorrekturmodellen für Deutschland, Frankreich, Italien, die Niederlande und Spanien. Unter Verwendung der Gewichte aus dem Harmonisierten Verbraucherpreisindex (HICP) wird ein gewichteter Durchschnitt aus den Ergebnissen der einzelnen Länder für den gesamten Euro-Raum berechnet. Im Ergebnis reagiert der HICP auf eine Abwertung des nominalen effektiven Euro-Wechselkursindex um 10% mit einem Anstieg um 0,4% nach 12 Monaten. Der Gesamteffekt entspricht in etwa 0,75% nach 3 Jahren. Die Länge des Anpassungszeitraumes ist für die einzelnen Länder unterschiedlich.

Keywords: pass-through, inflation, exchange rate, euro area
JEL-Codes: E31, F31, F41

1 Introduction

The continuous depreciation of the euro since its introduction and the recent strong appreciation have posed a challenge for monetary policy in the euro area. In the words of European Central Bank (ECB) president DUISENBERG: "[...] the depreciation of the exchange rate of the euro, until it is reversed, will increase the risks to price stability in the medium term." (DUISENBERG 2000). This view has been supported by the concerted foreign exchange interventions in September 2000 which were aimed at countering the depreciation trend of the common currency. However, up to now the impact of exchange rate changes on consumer prices in the euro area has not been quantified. Scientific studies of European exchange rate pass-through have been scarce as the time horizon since the introduction of the euro is rather short. This paper aims to present an estimate of the effects of a change in the nominal effective exchange rate index for the euro area on the Harmonised Index of Consumer Prices (HICP).

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In an extensive survey of the pass-through literature, MENON (1995) mentions several shortcomings of previous empirical pass-through studies. First, the country coverage has largely focussed on the United States. Our focus on Europe contributes to closing this gap. Second, several empirical studies use traditional Ordinary Least Squares (OLS) analysis and pay little attention to the time series properties of the data. This also holds true for the recent study by RANKI (2000) who focuses on the euro area. We will use up-to-date econometric methods (vector error correction models) to contribute to the existing literature. Third, most of the literature in the past years has concentrated on microeconomic factors (e.g., pass-through in selected product groups). Our study concentrates instead on aggregate data rather than on particular industries or products. We are primarily interested in the overall effect of exchange rate changes on consumer prices, an issue which is relevant for monetary policy.

Following MCCARTHY (2000), we examine the pass-through at different stages along the distribution chain (import prices, producer prices, and consumer prices). Since the time span for the analysis of aggregate euro area data is rather short we study exchange rate pass-through for selected European countries over a time period of 20 years to derive an estimate of the pass-through effects for the whole euro area. We concentrate on Germany, France, Italy, the Netherlands, and Spain as the key countries of the euro area, which together account for 86 percent of the Harmonised Index of Consumer Prices.

The paper is structured as follows: *Section 2* gives a short survey of the relevant literature. Data description and econometric procedures (stationarity and unit-root tests, cointegration tests, estimation of the vector error correction (VEC-)models) are laid out in *Section 3.* Results of the empirical analysis for each country as well as for the euro area as a whole are presented in *Section 4. Section 5* concludes.

2 Survey of the Literature

2.1 Theoretical Considerations

According to GOLDBERG and KNETTER (1997, p. 1248), exchange rate pass-through is defined as "the percentage change in local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries." However, changes in im-

port prices are to some extent also passed on to producer and consumer prices. Thus, in this paper exchange rate pass-through is seen more broadly as the change in consumer prices that can be attributed to a prior change in the nominal exchange rate.

Two channels of exchange rate pass-through are distinguished in the literature: A direct channel and an indirect channel.[1] Both gain in importance with an increase in the openness of an economy.

The *direct* channel of pass-through runs through the price of imports. The change in import prices is also likely to translate into changes in the producer and consumer prices of an economy if producers raise their prices in line with the increase in import prices.

The *indirect* channel of exchange rate pass-through refers to the competitiveness of goods on international markets. A depreciation of the exchange rate makes domestic products relatively cheaper for foreign buyers, and as a consequence exports and aggregate demand will rise and induce an increase in the domestic price level.

2.2 Empirical Literature

MENON (1995) presents an overview of 43 empirical studies of exchange rate pass-through. The main findings are: the majority of studies comes to the conclusion that exchange rate pass-through is incomplete.[2] However, the degree of pass-through seems to be quite different across countries and products. Factors that were found to influence the degree of pass-through are the openness and size of a country. The U.S. is by far the most often studied country. According to MENON (1995), different results for a country stem primarily from the use of different methodology, model spe-

1 See KAHN (1987), MENON (1995), and GOLDBERG and KNETTER (1997) for an exhausting discussion of exchange rate pass-through and LAFLÈCHE (1996) for an illustrative overview of the two channels. TAYLOR (2000) suggests a further channel via expectations. According to this view pass-through is highest when exchange rate changes are perceived to be persistent and prices adjust because of the expectations of the public.

2 GOLDBERG and KNETTER (1997) state that import prices in the U.S. only reflect about 50 percent of exchange rate changes (although the response of prices varies across industries). Probably the most often mentioned explanation for this phenomenon is the strategy of pricing-to-market by exporting firms (see KRUGMAN 1987; DORNBUSCH 1987). Rather than constantly adjusting its prices to exchange rate changes, an exporting firm may choose to hold the price constant and simply reduce or expand the mark-up on prices to avoid long-run losses of its market share (in case of an appreciation of the own currency).

cification and variable selection rather than from different time periods studied.

MENON criticises that most of the empirical studies employ an OLS estimation technique which does not properly take into account the time series properties, e.g. the non-stationarity of the data. He reports that KIM (1991) is the only study that uses a vector autoregressive (VAR) analysis with trend-removed data.

A further result that has been found in the literature according to MENON is that pass-through relationships have remained largely stable over time.[3] Some studies also find that pass-through has a different size when depreciations and appreciations are considered, i.e., that pass-through effects are asymmetric.

In the last years there has been some empirical work on exchange rate pass-through that tried to improve on the deficiencies of earlier studies as identified by MENON (1995). A comprehensive study is MCCARTHY (2000) who investigates exchange rate pass-through on the aggregate level for selected industrialised economies. He estimates a VAR model for the period 1976–1998 over the whole distribution chain (import, producer, and consumer prices) and finds that pass-through of exchange rate changes to consumer prices is modest in most of the analysed countries. The import share of a country and the persistence of exchange rate changes are found to be positively correlated with the extent of pass-through to consumer prices, while exchange rate volatility is found to be negatively correlated.

KIM (1998) estimates exchange rate pass-through for the U.S. using cointegration analysis and a vector error correction model. His paper relates producer price inflation in the U.S. to the trade weighted effective exchange rate, money supply, aggregate income, and interest rates. He finds that the exchange rate contributes significantly to producer prices which is supported by subsequent Granger causality tests.

KENNY and MCGETTIGAN (1998) also use cointegration analysis and vector error correction models to study exchange rate pass-through for Ireland. Comparing their results with previous studies they find that the degree of pass-through in their study is higher. Their main criticism of

3 This view has been increasingly challenged recently (see TAYLOR 2000; GAGNON and IHRIG 2001). Especially for countries that have adopted inflation targeting there seems to have been a reduction in exchange rate pass-through.

earlier studies is the neglect of the time-series properties of the data, particularly non-stationarity. RANKI (2000) is, to our knowledge, the only study so far that examines exchange rate pass-through for the whole euro area applying the OLS estimation technique. The most surprising result of his study is that the pass-through from the euro/US$ exchange rate into consumer prices is complete and occurs within one month. This finding contradicts with the existing empirical literature as well as with the recent experience within the euro area because the 25 percent depreciation of the euro in the first two years of his existence did not translate into similar increases in consumer prices so far.

Finally, a recent study applies panel estimation methods. GOLDFAJN and WERLANG (2000) investigate pass-through effects to consumer prices for a sample of 71 countries. They identify for the period of 1980–1998 that the pass-through effects on consumer prices increase over time and reach a maximum after 12 months. Over- and undervaluation of the real exchange rate, the initial inflation rate, GDP deviation from an estimated trend, and the degree of openness are found to influence the pass-through coefficient. In general, the pass-through is found to be substantially lower in developed economies than in emerging market countries.

Our survey of the empirical literature showed that there is a number of issues that leave space for further research. Our focus on the European experience tries to reduce the lack of sufficient country coverage. Second, many empirical studies used traditional OLS analysis and paid little attention to time series properties of the data. This also holds for the only euro area study so far (RANKI 2000). Following the work of KIM (1998) and KENNY and MCGETTIGAN (1998), we use cointegration analysis and vector error correction models to take account of the non-stationarity of several variables. Third, we focus on aggregate data rather than on certain industries or products. We are primarily interested in the overall effect of exchange rate changes on consumer prices, an issue which is most relevant for monetary policy. Thus, we agree with KENNY and MCGETTIGAN (1998, p. 1148) who state: "The partial nature of disaggregated studies means that findings of incomplete PT [pass-through], while very interesting in themselves, should not be adduced as evidence that this result carries over to the broader macroeconomy."

3 Data and Econometric Procedures

3.1 Data Description

For each country (France, Germany, Italy, the Netherlands, and Spain) we use seven variables: three price indices (import prices, producer prices, consumer prices)[4], the output gap (constructed using industrial production), a short-term interest rate, an oil price, and the effective nominal exchange rate. All data have a monthly frequency and are taken from the IMF International Financial Statistics except for the effective exchange rate indices which are obtained from the Bank of England (see *Appendix I* for details about the data). The time span covered is January 1982 until December 2000. While many of the variables used are available for a much longer period producer prices for France and Italy are only available from 1980 and 1982 onwards, respectively. Thus, we take 1982 as the starting date for our analysis. Exchange rate data are effective nominal exchange rates of the national currencies. The oil price is the petroleum spot price of UK Brent. We use the national call money rates to approximate central bank behaviour. The output gap is computed as the difference between actual industrial production and potential output (calculated using a Hodrick-Prescott filter). Import prices, producer prices, and consumer price indices reflect the pass-through at several stages of the production chain.

We follow McCarthy (2000) and include all stages of the distribution chain in our estimation. This gives us the opportunity to analyse how exchange rate fluctuations pass through the production process from the import of products to the consumer level. In our model the oil price serves as a proxy for supply shocks, and the output gap models demand shocks. Short term interest rates are used to incorporate central bank policy in our system. By including a separate central bank reaction function we follow the result of Parsley and Popper (1998) who find that taking into account monetary policy significantly improves the estimation results of exchange rate pass-through. Since central banks that target consumer price inflation will try to insulate prices from exchange rate movements, neglecting their behaviour should distort the true consequences of ex-

4 We use the national consumer price indices. While the use of harmonized indices would be more desirable for the calculation of a European pass-through effect, data for harmonized consumer price indices are unfortunately only available from 1990 onwards (for some countries only from 1995 on). As it is particularly important to use long-term time series for cointegration analysis we decided to use the national price indices. We use consumer prices rather than an index of core inflation since this is the figure most closely watched by the public and is also the focus of the ECB.

change rate variations. That way, the observed relationship between prices and exchange rates would take into account the central bank behaviour rather than the direct influence of exchange rates on prices.[5]

3.2 Stationarity and Unit-Root Tests

In order to determine whether the variables exhibit non-stationary behaviour we perform unit-root tests.[6] We start using the KPSS test (see KWIATKOWSKI, PHILLIPS, SCHMIDT and SHIN 1992) which tests the null hypothesis "stationarity" against a unit root alternative. The test uses the regression of the time series to be analysed (Y_t) against a constant ("stationarity") or a constant and a time trend ("trend stationarity"):

$$(1) \quad Y_t = \alpha + \beta \cdot t + \varepsilon_t$$

Then the stationarity of the residuals of these regressions (ε_t) is tested. An essential part of the test statistic is the consistent estimation of the variance of the residual time series. Usually a Bartlett kernel is used to estimate a heteroskedasticity and autocorrelation consistent variance. The KPSS test statistic therefore depends on the choice of the lag length of the Bartlett kernel that is needed to correct for autocorrelation in the residual term. HOBIJN ET AL. (1998) analysed different approaches to choose the lag length and concluded that the automatic lag selection procedure developed by NEWEY and WEST (1994) improves the performance of the test compared with the original KPSS test. Therefore, we also used this generalised KPSS procedure to test for stationarity.[7]

The results can be summarised as follows. Almost all variables are clearly non-stationary as the KPSS test in most cases rejects the null hypothesis "stationarity" in both test versions at usual significance levels. Only for the German producer price index, the import price index in France and

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5 Both PARSLEY and POPPER (1998) and MCCARTHY (2000) include a monetary aggregate in their system of variables. We are including interest rates (INT) instead and thereby follow the evidence of BERNANKE and MIHOV (1997) who showed that monetary targets were not significant in the Bundesbank reaction function. Furthermore, most central banks in the world by now target short-term interest rates. GERLACH and SVENSSON (2000) provide further evidence for the euro area that the relationship between money-growth and future inflation is weak at best.

6 We did not include the output gap (GAP) in these tests as this variable is stationary by construction due to the application of the Hodrick-Prescott Filter.

7 The tables including the test statistics for the two regressions ((a): constant and (b): constant and trend), the chosen lag length, and the conclusion concerning the degree of integration are available from the authors upon request.

Spain the null hypothesis of stationarity could not be rejected. Therefore, we also conducted an Augmented Dickey-Fuller (ADF) unit-root test which has the null hypothesis "non-stationarity" for these three variables. Equation (2) shows the general form of the test regression for the ADF test including a constant and a linear time trend:

$$(2)\quad dY_t = \alpha + \beta \cdot t + \sum_{i=1}^{p} \lambda \cdot dY_{t-i} + \delta \cdot Y_{t-1} + \varepsilon_t$$

The lag length (p) of the first differences in the ADF equation (dY_{t-i}) has been chosen according to the AIC2-rule developed by PANTULA ET AL. (1994). The lag length is determined as the lag at the minimum of the Akaike Information Criterion (AIC) plus two lags.[8] The results of the ADF tests confirm that the French import prices and the German producer prices are stationary. Concerning Spanish import prices the ADF tests cannot reject the null hypothesis. As the KPSS test also could not reject the null hypothesis of "stationarity" it remains unclear whether this time series should be considered as stationary or non-stationary.

The result for Germany that PPI is a I(0)-variable whereas CPI is I(1) does not make much sense in economic terms.[9] As a consequence PPI and CPI would be unrelated in the long run. The same is true for France (import prices are unrelated to PPI and CPI in the long run) and for Spain, if the Spanish import prices are taken as I(0). To overcome this problem of unbalanced long-term equations we treat all three variables (German PPI, import prices in France and Spain) tentatively as I(1)-variables in the cointegration tests and VEC-models.

3.3 Cointegration Tests and Estimation of VEC-Models

In the main part of our analysis we construct a vector error correction model for each of the five countries. Then we carry out impulse-response analyses to estimate the pass-through effect of changes in the effective exchange rate to prices at the import, producer and consumer level.

In a first step we take account of the non-stationarity of the majority of the variables and apply cointegration tests for each country. If these tests indicate the presence of one or more cointegrating equations (CE) we

8 The tables including the results are available from the authors upon request.
9 We are grateful to an anonymous referee for helpful comments on this problem.

estimate in a second step VEC-models that incorporate the long run relationships among the variables.

To determine whether the seven variables in our system are cointegrated we use the Johansen procedure. We include all variables in the test, i.e., all non-stationary and all stationary variables.[10] As a consequence, the cointegration rank increases by the number of stationary variables.[11] The correct number of cointegrating equations to be included in the VEC-model is therefore equal to the number of CE found by the Johansen test minus the number of stationary variables. As mentioned in *Subsection 3.2* we treat the German producer prices as well as the French and Spanish import prices tentatively as I(1)-variables due to economic reasoning, although the stationarity and unit-root tests show that these variables might be I(0).

Table 1 Summary of the VEC-Models Used for Impulse-Response Analysis

	No. of Lags[1]	No. of CE[2]	Type of Model
France	3	2	Constant in CE and VAR
Germany	3	2	Constant in CE and VAR
Italy	2	1	Constant + linear trend in CE, constant in VAR
Netherlands	1	2	Constant in CE and VAR
Spain	1	2	Constant in CE and VAR

Source: Calculations by the authors.

Note: [1] The optimal number of lags in the VEC-models was determined using the AIC criterion.

[2] The number of cointegrating equations is equal to the number of CE found by the Johansen test minus the number of stationary variables (= 1; output gap).

Table 1 summarises the results of the Johansen tests. The detailed analysis for each country is shown in the *Tables 7a–7e* in *Appendix 2*. We have found 2 or 3 cointegrating vectors for each country. The correct number of CE

10 According to e.g. HANSEN and JUSELIUS (1995), the selection of variables to be included in cointegration tests should be based on economic reasoning, i.e., stationary variables should be included if reasonable. However, at least two variables need to be non-stationary in order to find cointegration.

11 See e.g. HANSEN and JUSELIUS (1995).

after subtracting the number of stationary variables is one for Italy and two for the other four countries.

Table 1 also shows that the specification of the VEC-models is similar across the five countries. With the only exception of Italy we included only a constant in the cointegrating equations and in the short-term part of the VEC-model. Only in the model for Italy was the linear time trend in the cointegrating equation significant.

Before estimating the VEC-models the long-run equations have to be identified as in case of more than one CE the Johansen procedure leads only to an identification of the cointegrating space spanned by the CE, but not of the single cointegrating equations. Thus, we include economically sound restrictions with the aim to identify the CE and to reach parsimonious models. In addition, we assume that the oil price is weakly exogenous, i.e., it does not react to deviations from long-run equilibria and, thus, the alpha coefficients of the oil price are restricted to zero.

Table 2a VEC-Models: Cointegration Relationship No. 1

	Germany	France	Spain	Netherlands	Italy
Oil Price	–0.023	0	–0.166	–0.033	–0.050
Effective FX	0.207	0.239	0	0	0
Import Price	–0.266	–0.243	–0.280	–0.072	–0.156
Output Gap	0	0	–0.036	0	–0.007
Interest Rate	–0.024	–0.052	–0.029	–0.054	0
PPI	1.00	1.00	1.00	1.00	1.00
CPI	–0.402	0	0	–0.751	–0.629
Constant	–0.453	–2.21	–2.575	–0.639	–0.881
Trend	–	–	–	–	0.0002
P-Value	45.7%	84.3%	77.7%	79.9%	49.1%

Source: Calculations by the authors.

Notes: "0" = restricted to zero; "P-Value" = implicit Type I-error if the hypothesis that all zero restrictions of the model (restrictions on cointegration relationships 1 and 2 and restriction on the alpha coefficients of the oil price) is rejected.

Table 2b VEC-Models: Cointegration Relationship No. 2

	Germany	France	Spain	Netherlands
Oil Price	0	0	–0.316	0
Effective FX	28.256	1.752	0.233	–1.298
Import Price	–23.360	0	0	1.00
Output Gap	2.101	0.301	–0.065	–0.095
Interest Rate	1.00	0	0	0
PPI	0	–1.636	0	0
CPI	0	1.00	1.00	0
Constant	236.123	11.045	–2.316	–10.659

Source: Calculations by the authors.
Notes: "0" = restricted to zero.

Tables 2a and *2b* show the results of the identification process concerning the long-run equations. "P-Value" (see *Table 2a*) indicates that for all countries the zero restrictions on the long-run equations and the alpha coefficients cannot be rejected at usual significance levels. The first CE indicates for all countries a long-run equation for the producer price index which is used to normalise the long-run equation:

$$(3)\quad PPI_t + \sum_i \beta_i X_{i,t} = 0$$

In all these equations the beta coefficient of the import prices has a negative sign which means that import prices have a positive impact on PPI and are an important driver of PPI in the long run. With the exception of France the oil price also determines the long-run path of PPI, but it has a smaller impact than the import prices. With the exception of Italy the interest rate enters also with a negative sign in equation (3). This means that, for example, a higher PPI usually coincides with higher interest rates. In the second CE different variables are used for normalisation: in case of France and Spain the CPI, for Germany the interest rate, and for the Netherlands the import price index is used.

Looking at both cointegrating equations we find for Germany (CE1, CE2), France (CE1) and the Netherlands (CE2) an interesting relationship between import prices and the effective exchange rate: the coefficients of these two variables are relatively similar but with opposite signs.

This means that changes in the currency are (partly) balanced by opposite changes in the import prices, and vice versa.

Table 3 Adjustment to Deviations from the Long-Run Equilibrium: Significance of the Alpha Coefficients for the Cointegrating Equations

	Germany		France		Spain		Netherlands		Italy
	CE1	CE2	CE1	CE2	CE1	CE2	CE1	CE2	CE1
Oil Price	0	0	0	0	0	0	0	0	0
Effective FX	–	***	**	***	–	–	–	–	–
Import Price	–	–	**	***	–	–	–	–	–
Output Gap	–	***	–	***	–	–	–	***	***
Interest Rate	***	–	–	**	–	–	***	–	***
PPI	***	–	***	–	***	***	***	–	***
CPI	–	–	***	–	–	–	***	–	***

Source: Calculations by the authors.

Notes: "0" = restricted to zero; "–" = not significant; *, **, *** = significant at the 10%, 5%, or 1% significance level, respectively.

The adjustments to deviations from the long-run equilibria are measured by the so-called alpha coefficients. *Table 3* shows that in all countries the PPI adjusts to a disequilibrium in the first CE, but also often the CPI and the interest rates react. With the exception of France the effective exchange rates and the import prices are weakly exogenous to CE1. With the exception of Italy the output gaps are weakly exogenous as well. In contrast, disequilibria in the second CE are most often reduced by adjustments in the output gaps and – in case of Germany and France – also by changes in the effective exchange rates.

As the treatment of German PPI as well as French and Spanish import prices as I(1)-variables leads to meaningful economic results, these VEC-models are used to conduct impulse-response analyses in the next section.

4 Results

4.1 Impulse-Response Functions

In order to determine impulse-response functions the variables need to be given a plausible ordering. This is to some extent subjective and is done with a fair amount of plausibility. We used the following ordering for the impulse-response analysis:

$$\text{OIL} \rightarrow \text{EX} \rightarrow \text{IMP} \rightarrow \text{GAP} \rightarrow \text{INT} \rightarrow \text{PPI} \rightarrow \text{CPI}$$

We have the following model of pass-through in mind. Both oil price and exchange rate changes influence import prices. Since the oil price is likely to have an influence on the exchange rate but not vice versa we start our causal structure with the oil price. Import prices directly influence economic activity, i.e., the output gap. The central bank takes into account both developments in import prices (as a predictor of future inflation) and the output gap in its monetary policy rule. Thus, short-term interest rates are set next. Our final two variables are producer prices which directly influence consumer prices. We also checked for alternative orderings, particularly different orderings among IMP, GAP and INT, and found that this did not change the results in a significant way.

In the study of McCarthy (2000) interest rates rank last, as he assumes a reactive behaviour of the central bank. However, we argue that the position of the interest rate might also be prior to the producer prices. Given the long and variable lags of monetary policy, central banks usually react to expected inflation rather than realised inflation (forward-looking behaviour).[12] In this respect it would make sense to position the interest rate variable prior to the producer price index and thus let prices react to central bank policy, i.e., central banks set interest rates after observing leading indicators for inflation like oil prices, exchange rate changes or import prices. However, we did not find significant changes due to a different ordering of the interest rate.

12 See Clarida et al. (1999).

Table 4 Effects of National Consumer Price Indices to a 1%-Exchange Rate Shock

	After 6 months	After 12 months	After 18 months	After 24 months
France	0.02	0.08	0.13	0.17
Germany	0.05	0.06	0.07	0.07
Italy	0.05	0.08	0.11	0.12
Netherlands	0.13	0.13	0.13	0.13
Spain	0.09	0.08	0.07	0.07

Source: Calculations by the authors.

Notes: The effects are measured as percentage changes in the national consumer price indices in response to a shock in the national effective exchange rate indices.

Table 4 displays the responses of national consumer prices to a one-percent shock in the national nominal exchange rate indices (an increase corresponds to a depreciation) after 6, 12, 18 and 24 months.[13] As expected, consumer prices increase in response to the depreciation. However, extent and speed of pass-through differ across countries which is a well-known phenomenon (see MENON 1995). The fastest effect can be observed in the Netherlands with a consumer price index increase of 0.13 percent after 6 months. In the long run, pass-through is highest in France, the Netherlands and Italy with a response of 0.17, 0.13 and 0.12 percent after two years, respectively. In Spain only 0.07 percent of the initial exchange rate change is reflected in consumer prices.

Whereas impulse-response functions trace the effects of a shock to one endogenous variable on to the other variables in the VEC-model, variance decomposition separates the variation in an endogenous variable into the component shocks to the endogenous variables of the VEC-model. Thus, the variance decomposition provides information about the relative importance of each random innovation in affecting the variables in the system. *Table 5* displays the variance decomposition of import, producer and consumer prices to a shock in the effective exchange rate after 6, 12 and 24 months.

13 The response patterns of consumer prices to a one-standard deviation shock in effective exchange rates are shown in *Appendix 3* for each country separately.

Table 5 Variance Decomposition: How Much Does the Exchange Rate Explain?

	Import prices			Producer prices			Consumer prices		
Months after initial shock	*6*	*12*	*24*	*6*	*12*	*24*	*6*	*12*	*24*
France	27.2	23.9	21.9	0.9	3.0	5.8	0.2	2.1	6.7
Germany	30.9	32.6	28.3	21.7	21.1	15.9	0.7	0.9	0.9
Italy	20.0	21.0	20.3	7.6	8.4	7.3	1.0	1.4	1.4
Netherlands	33.0	35.4	36.4	24.0	24.6	21.0	9.2	7.5	4.7
Spain	13.4	15.4	16.3	14.2	18.1	17.7	4.9	4.0	2.2

Source: Calculations by the authors.

Notes: Displayed are the percentages of the price variable variances that result in response to a one-standard deviation shock in the national effective exchange rate indices.

It can be seen that changes in effective exchange rates explain a fairly large part of the variation of import prices while this effect declines along the distribution chain. Again, results differ across countries. Roughly one third of the import price variance in France, Germany and the Netherlands is explained by exchange rate movements. In Italy and Spain the effect amounts to only about 20 and 15 percent, respectively.

The influence of the effective exchange rate on producer prices is significantly smaller in all countries, except for Spain. In France and Italy exchange rate shocks account for less than 10 percent of the variance of the producer prices. In Germany and Spain about 16 and 18 percent, respectively, of the producer price variance is due to the exchange rate in the long run (= 24 months). It is remarkable that the effect on the Spanish producer prices is as large as that on the import prices. The largest effect on producer prices is found for the Netherlands where more than 20 percent of the producer price variance can be explained by exchange rate shocks.

Consumer price variance, on the other hand, is affected by exchange rate fluctuations only to a small degree. France and the Netherlands exhibit the largest effect with 6.7 and 4.7 percent after two years. In the other countries only between one and three percent of the variance of consumer prices can be explained by exchange rate shocks in the long run. In the Netherlands and Spain there is a strong impact of the exchange rate

on consumer price fluctuations in the short term which then diminishes in the long-run.

The results shown in *Table 5* provide a relative ranking of the magnitude of the exchange rate effect across countries for the explanation of price changes. The largest fraction of import price changes explained by exchange rate changes is found in Germany, the Netherlands and France. The effect on producer prices is relatively large in the Netherlands, Spain and Germany, and the Netherlands and France exhibit the strongest long-run impact on consumer prices.

McCarthy (2000) finds that a country's import share is positively correlated with exchange rate pass-through to consumer prices. *Table 6* shows the import shares of the five countries in our sample. The figures shed light on the different pass-through effects. The results of *Table 5* also showed that the impact of exchange rate changes on consumer prices occurs relatively fast in the Netherlands, whereas in France, Italy and Germany the adjustment takes much longer. This could be partially due to the very large import share of the Netherlands. Combined with the relatively large share of imports from non-European countries (e.g. the United States) this helps to explain the stronger impact of exchange rate fluctuations on the variance of the Dutch prices.

Table 6 The Structure of Imports

	Import share	Imports from non-European countries
France	21.56%	33.01%
Germany	25.68%	37.27%
Italy	20.60%	33.59%
Netherlands	48.13%	39.54%
Spain	22.85%	32.85%

Sources: IMF (for GDP; see *Appendix 1*); OECD (2000); calculations by the authors.

Notes: Import shares are calculated as the average of imports as a percentage of GDP over 1989-98. Imports denominated in US$ are used as proxy for the percentage of imports from countries outside of OECD-Europe in 1994.

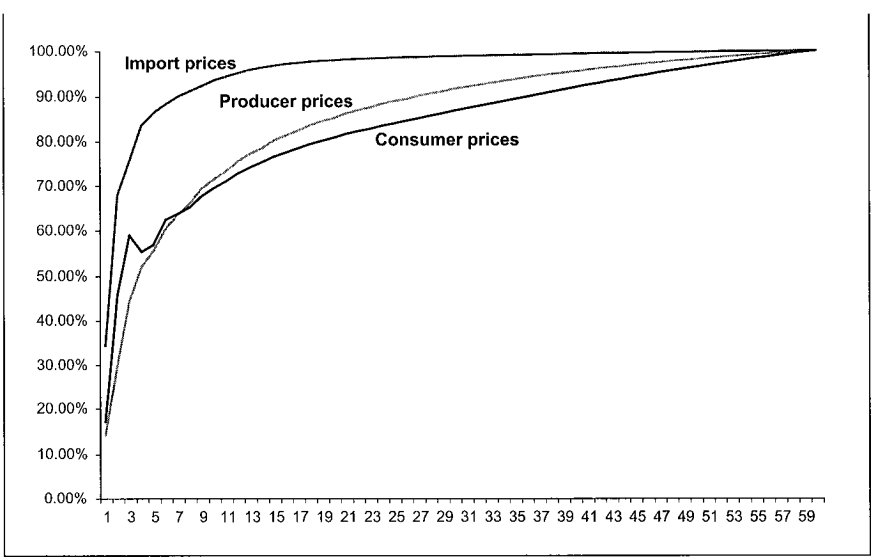
Figure 1 compares the average speed of price adjustment for import, producer and consumer prices after an exchange rate shock. The speed of ad-

justment is defined as the ratio of the price response after t periods relative to the long run response in percent. As expected the adjustment period is very short for import prices and much longer for consumer prices. The speed of adjustment of producer prices is between the results for import and consumer prices.

To sum up, the speed of adjustment of consumer prices is rather slow compared with the aforementioned import prices. On average, after 36 months about 84 percent of the total pass-through effect is reflected in the consumer price inflation rate whereas after twelve months only about 70 percent of the total adjustment has occurred. In comparison, after one year already 95 percent of the adjustment in import prices has materialised.

This difference can be explained by the indirect channel of exchange rate pass-through: an exchange rate depreciation leads to increased competitiveness of the export sector and thus promotes economic activity. In the longer term this will put upward pressure on prices, and thus the inflation rate tends to increase. Whereas the direct effect through higher import prices works much faster, these indirect effects can be expected to influence the consumer price inflation rate only slowly.

Figure 1 Speed of Price Adjustment After an Initial Exchange Rate Shock



Source: Own graph.

Note: The x-axis displays the months after the initial exchange rate shock.

4.2 Aggregation across EMU countries

The most interesting question in the European context is the overall effect of a change in the effective exchange rate of the euro on aggregated consumer prices in the euro area (HICP). This is also the relevant issue from the viewpoint of monetary policy. Since the ECB is only concerned with aggregated inflation rather than with price developments in the different member countries a measure of exchange rate pass-through for the HICP is needed.

In the preceding section we already obtained results for the exchange rate pass-through in five of the twelve EMU countries over a longer time period. In order to aggregate results, the weights of the national inflation rates in the calculation of the HICP have to be taken into account.[14] The five countries in our sample represent 86 percent of the weight of the HICP. Thus, we should be able to calculate a fairly precise measure for the exchange rate pass-through.

With 30.9 percent, the German inflation rate has the largest influence on the European HICP, followed by France with 20.5 percent. The smallest country in our sample, the Netherlands, has an influence of 5.3 percent. Even though one might argue that countries like Ireland or Portugal, which as small open economies can be expected to have a high pass-through coefficient, are missing in our sample, they do not have a significant influence on the calculation of the HICP.

In order to compute the effect of a change in the euro effective exchange rate on the aggregate consumer price index we need to take into account the specific definition of this exchange rate index. In the analysis described in the previous sections we used the national effective exchange rate indices. These indices incorporate the fixed bilateral intra-European exchange rates, whereas the nominal effective exchange rate index for the euro area only encompasses countries outside the euro area. As a consequence, the euro effective exchange rate index does only cover about half of the import and export trades of the EMU countries.

For example, the national effective exchange rate index for Germany is calculated as

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14 We are aware that the national consumer price indices are not harmonized and may differ to some extent in their construction. However, long time series for the harmonized indices are not available.

$$(4)\quad E_N = \alpha_1 \frac{DM}{US\$} + \alpha_2 \frac{DM}{JPY} + \ldots + \beta_1 \frac{DM}{FRF} + \beta_2 \frac{DM}{ITL} + \ldots \text{ and } \sum_{i,j} \alpha_i + \beta_j = 1$$

with α_i as the trade weights concerning non-EMU-countries and β_i as the trade weights concerning EMU-countries. After the introduction of the euro the exchange rates for the EMU-countries were replaced by the fixed conversion rates. However, the effective exchange rate index for the euro area, which is calculated by the ECB, only incorporates exchange rates with non-EMU countries. Thus, all β_j in equation (4) are set to zero.[15] Therefore, the national effective exchange rate indices and the euro effective exchange rate (E) are related according to formula (5), where the weights are defined as in (1):

$$(5)\quad \Delta E_N = \Delta E \cdot (1 - \sum_j \beta_j) = \Delta E \cdot \sum_i \alpha_i$$

Thus, to calculate the pass-through effects of a change in the euro effective exchange rate index, we need to multiply it with the share of trade between euro area countries and non-euro area countries. According to OECD (2000), 49.6 percent of the total trade of the euro-zone was conducted with countries outside the euro-zone in 1999. Applying the weighting scheme for the euro effective exchange rate index of the ECB results in a weight of 49.5 percent.[16]

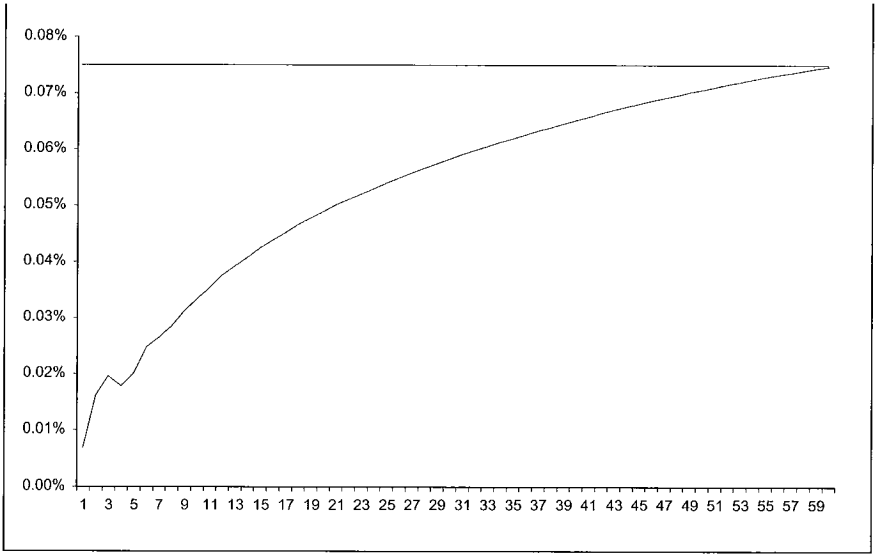
Figure 2 shows the cumulative effect on the European price level as measured by the HICP.[17] One year after the nominal effective euro index has depreciated by one percent, consumer prices in the euro area increase by 0.04 percent. The total effect converges at about 0.075 percent. Accordingly, a ten percent shock amounts to a pass-through to consumer prices of 0.75 percent. Three years after the initial shock about 84 percent of the total adjustment has occurred. These results are comparable with the findings of KAHN (1987) for the U.S. He finds that after a ten percent depreciation of the US$ consumer prices in the U.S. tend to increase by about 0.84 percent after nine quarters. As the euro area is often compared with the U.S. in terms of openness and size of the economy, the similarity of our results with KAHN's findings is reassuring.

15 See BULDORINI ET AL. (2002), p. 12.

16 The weighting scheme for the euro effective exchange rate index of the ECB calls for exports to be double-weighted to take account for "third-market effects" which reflect the competition that euro area exporters face in foreign markets from domestic producers as well as from exporters of third countries (see BULDORINI ET AL. 2002).

17 The weights for each country are rescaled so that all five countries together represent 100 percent. The adjusted shares used as weights are: Germany 36 percent, France 24 percent, Italy 22 percent, Netherlands 6 percent, and Spain 12 percent.

Figure 2 Effects of a 1% Depreciation of the Euro Effective Exchange Rate on the HICP



Source: Own graph.

Note: The x-axis displays the months after the initial exchange rate shock.

A look at the past development of inflation in the euro area might be quite illustrative at this point. According to our results, the roughly ten percent depreciation of the effective exchange rate of the euro during the year 1999 seems to be partly responsible for the increase in the inflation rate that followed in 2000. From January to December 2000 the HICP rose from 1.9 percent to 2.6 percent. Our estimates suggest that about 0.4 percentage points of this increase can be attributed to exchange rate pass-through. The remaining part of the increase (about 0.3 percent) might be explained by other factors like e.g. the oil price increase and domestic causes.

5 Conclusion

In this paper we analysed the effects of exchange rate fluctuations of the euro on the Harmonised Index of Consumer Prices. As the time period since the introduction of the euro is rather short we studied exchange rate pass-through for the core countries Germany, Italy, France, the Netherlands and Spain over the last twenty years using a vector error correction model. We find that the Netherlands exhibit the fastest pass-through of exchange rate changes to consumer prices, and long run effects are highest in Italy and France. Pass-through coefficients, i.e., the share of the exchange rate change that is reflected in consumer prices, ranges from 8 percent (France) to 13 percent (Netherlands) after one year. After two years, coefficients range from 7 percent (Germany, Spain) to 17 percent (France).

Aggregating the national results using the weights of each country's inflation rate in the HICP we find that, on average, a ten percent depreciation of the effective euro exchange rate leads to an increase of 0.4 percentage points in the inflation rate after one year. The total effect converges to 0.75 percentage points after about three years. This amounts to an exchange rate pass-through to consumer prices of 7.5 percent of the initial exchange rate shock for the euro area.

Our result is relevant for policymakers, especially in central banks. While the exchange rate enters the strategy of the ECB via the second pillar, there is no target for the exchange rate—it only becomes relevant for monetary policy in that it influences the inflation rate. However, up until now there is no consensus about the importance of different factors that influence inflation, and thus ECB watchers are often unclear about the weights the ECB puts on factors like the exchange rate. Nevertheless, operations like the concerted foreign exchange interventions in September 2000 and the subsequent interest rate increase suggest that the monetary authority does indeed care about the euro exchange rate. Our findings suggest that the exchange rate indeed has an influence on consumer prices that should be taken into account if price stability is threatened.

The major problem for an analysis of euro area exchange rate pass-through is the lack of sufficiently long time series. By estimating pass-through for the member countries separately our approach is a first step in quantifying the total effect. However, in interpreting the results, several caveats have to be kept in mind.

First, we do not know whether the estimated relationships are stable over time or if exchange rate pass-through has changed significantly during the period studied.[18] In particular, we do not know if the introduction of the euro has changed the relationships between the exchange rate and the prices. The observation period for the two years since the beginning of the European Monetary Union is still too short to test for structural breaks in the short-term or long-term relationships of our models.

Second, we do not address the question of possible asymmetry of pass-through in appreciation and depreciation periods which has been mentioned in the literature (see COUGHLIN and POLLARD 2000). To test for this effect, researchers have usually divided the time series into appreciation and depreciation periods. Our VEC-model approach incorporates long-term relationships between variables and, thus, separate estimates for sub-periods are not feasible since the period considered would then be too short to perform cointegration tests. Moreover, a recent study for European Union countries by GIL-PAREJA (2000) found little evidence of asymmetry.

18 The literature on this subject presents mixed results: MENON (1995) reports that only a few studies find structural breaks in the pass-through relationship. However, a recent study by GAGNON and IHRIG (2001) suggests that pass-through has changed in many countries in the 1990s, specifically in countries that adopted inflation targeting. They did not find a structural break, however, for Germany – the only country of our sample they included in their study.

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Appendix 1: Data Description

All data have monthly frequency. Where available, data are taken from the IMF International Financial Statistics (IFS) (obtained through Thomson Financial Datastream). All time series are included in logarithms, except for interest rates and output gap. The interest rates are used as (log(1+r)).

Oil price (OIL): World petroleum spot price in US$ per Barrel, UK Brent (Source: IFS code 11276AAZZF...).

Nominal Effective Exchange Rate Indices (EFF): Source: Bank of England (Datastream codes BDDMEF.., FRFRANCE, ITLIREFF, NLGUILDE, ESPESEFF), period average.[19]

Short term interest rates (INT): Money market rates (Source: IFS line 60B..ZF...), period average. From the start of EMU (1999:1) onwards for all countries the German call money rate.

Output gap (GAP): Computed as the difference between industrial production (Source: IFS line 66...CZF..., seasonally adjusted) and potential production (computed with a Hodrick-Prescott-Filter, smoothing parameter: 14.400).

Import prices (IMP): Germany, Italy, Netherlands, and Spain – Import price index (Source: IFS line 76.X.ZF...), 1995=100; France – Raw materials import prices (Source: Datastream code FRIMIRAWF).

Producer prices (PPI): Germany, Netherlands, Spain, and Italy – PPI (Source: IFS line 63...ZF...); France – PPI Intermediate goods for industry (Source: Datastream code FRPPIINTF).

Consumer prices (CPI): CPI index (Source: IFS line 64...ZF...).

GDP: Gross Domestic Product (Source: IFS line 99b).

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19 Concerning these series the Bank of England states: "Data for the eleven currencies continue to be published using the existing trade weights and by converting the current euro exchange rate to that of the legacy currency exchange rate using the fixed conversion values as defined on the 31st December 1998. It should be noted that for these legacy currencies the effective exchange rate indices should be referred to as national competitiveness indicators. These rates will tend to be more stable than before 1999 because a large proportion of each countries trade will be with other euro area countries – thus no exchange rate movements." (see Internet: http://www.bankofengland.co.uk/mfsd/rates/effective.htm (downloaded July 30, 2003); and BANK OF ENGLAND 1999).

Appendix 2: Results of the Cointegration Tests (Johansen Test)

Table 7a Results of the Cointegration Tests for France (Johansen Test)

Jan. 1982 – Nov. 2000 Lag length (AIC) = 3	Trace Test (Constant in CE and VAR)[1]	Number of CE in VECM[2]
None	186.91**	
At most 1	116.83**	
At most 2	73.98*	3 – 1 = 2
At most 3	42.72	
At most 4	20.33	

Table 7b Results of the Cointegration Tests for Germany (Johansen Test)

Jan. 1982 – Dec. 2000 Lag length (AIC) = 3	Trace Test (Constant in CE and VAR)[1]	Number of CE in VECM[2]
None	154.97**	
At most 1	107.18**	
At most 2	70.26*	3 – 1 = 2
At most 3	35.37	
At most 4	15.98	

Table 7c Results of the Cointegration Tests for Italy (Johansen Test)

Jan. 1982 – July 2000 Lag length (AIC) = 2	Trace Test (Constant + linear trend in CE, constant in VAR)[1]	Number of CE in VECM[2]
None	204.86**	
At most 1	143.07**	2 – 1 = 1
At most 2	85.55	
At most 3	56.54	
At most 4	32.63	

Table 7d Results of the Cointegration Tests for the Netherlands (Johansen Test)

Jan. 1982 – Dec. 2000 Lag length (AIC) = 1	Trace Test (Constant in CE and VAR)[1]	Number of CE in VECM[2]
None	215.32**	
At most 1	139.00**	
At most 2	72.79*	3 – 1 = 2
At most 3	35.87	
At most 4	14.66	

Table 7e Results of the Cointegration Tests for Spain (Johansen Test)

Jan. 1982 – Dec. 2000 Lag length (AIC) = 1	Trace Test (Constant in CE and VAR)[1]	Number of CE in VECM[2]
None	185.98**	
At most 1	105.52**	
At most 2	72.34*	3 – 1 = 2
At most 3	44.07	
At most 4	20.99	

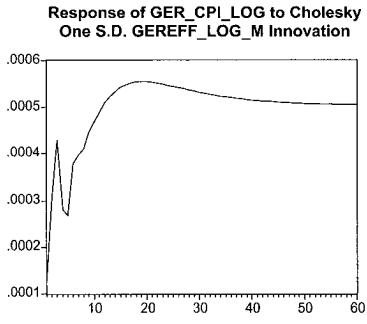
Source: Calculations by the authors.

Notes: Testing procedure = Johansen test; seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate; **, * = significant at the 1% or 5% significance level, respectively.

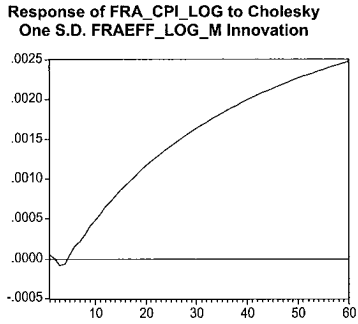
[1] A linear trend in the cointegrating equations (CE) was not significant, except in Italy.

[2] Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 1; output gap). French and Spanish import prices and German PPI are treated as I(1)-variables (see *Section 3*).

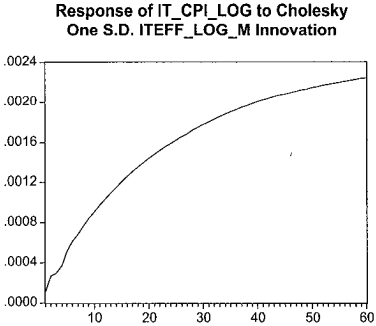
Appendix 3: Response of National Consumer Prices to a One-Standard Deviation Exchange Rate Shock



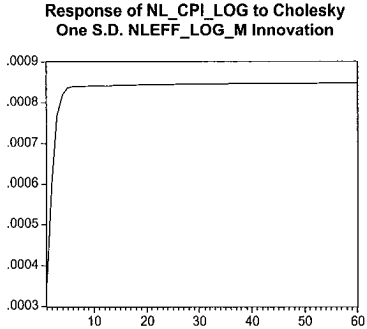
Germany



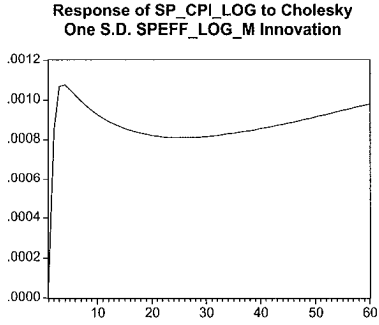
France



Italy



The Netherlands



Spain

Source: Own graphs.