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

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Non-technical summary

We analyze exchange rate pass-through, i.e. the change in local currency prices resulting from variations in the exchange rate, for consumer prices in the euro area. We first estimate country-specific pass-through coefficients for five large countries of the euro area (Germany, France, Italy, Spain and the Netherlands) using time series data for the past 20 years. Following this we construct a weighted average of these coefficients using the weight of each country in the Harmonized Index of Consumer Prices (HICP).

As Menon (1995) in a comprehensive survey of the relevant literature points out, former empirical studies of exchange rate pass-through focus largely on the US and often neglect the time series properties of the data. To our knowledge, Ranki (2000) is the only source so far that analyzes data for the euro area. Furthermore, many recent studies analyze the pass-through to import prices of different products on the micro level rather than focussing on the effects of aggregate price measures like consumer price indices.

We contribute to the existing literature in several ways. First, our study presents one of the first estimates of the effects of changes in the euro exchange rate on the Harmonized Index of Consumer Prices (HICP) in the euro area. Second, we estimate Vector Error Correction Models to take account of the non-stationarity of most of the used variables and cointegration relationships between them. Third, while a large part of the literature in the past years has focussed on the question "why" there is incomplete pass-through to import prices we present quantified effects on aggregate consumer price indices. Thus, our study is in the spirit of Kim (1998) and McCarthy (2000) who tackle related questions for other markets (Kim (1998) studies the US market) or use different econometric methods (McCarthy (2000) who applies Vector Autoregression) and is of direct relevance for monetary policy makers.

Since aggregated time series data for the euro area are only available from 1999 on we estimate exchange rate pass-through for five large EU-countries separately and then compute an average for the euro area using the relative weight of each country in the HICP. Our country sample includes Germany, France, Italy, Spain and the Netherlands which together represent about 86 percent of the influence on the HICP. Thus, we believe our results are a rather robust estimate of the exchange rate influence.

Our study uses monthly data from 1981 until 2001 and includes as variables nominal national effective exchange rate indices, short-term interest rates, output gaps constructed from industrial production, the oil price and all three stages of the distribution chain: import, producer and consumer prices. After performing unit root tests we cointegration relationships between the variables for each of the countries used in our sample. Thus, we estimate Vector Error Correction Models for the five

countries and generate impulse-response functions in order to quantify the effect of an exchange rate shock on consumer prices.

Regarding the country specific results we find that the Netherlands exhibit the fastest pass-through of exchange rate changes to consumer prices, but the long run effects are highest in Italy and France. Pass-through coefficients, i.e. the share of exchange rate change that is reflected in consumer prices, ranges from 7 (France) to 12 percent (Italy) after one year. After two years, coefficients range from 8 (Spain) to 18 percent (Italy). As expected, the extent of pass-through declines along the distribution chain with the largest effect occurring in import prices.

By computing the variance decompositions for each country we obtain 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 impact on consumer prices. Along with the existing literature we explain the stronger impact in the Netherlands with their import share which is the largest among the five countries in our study.

Aggregating the national results using the relative 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 euro area inflation rate after one year. The total effect converges to 0,8 percentage points after about three years. This amounts to an exchange rate pass-through to consumer prices of 8 percent of the initial exchange rate shock. The result shows that the euro exchange rate does have an effect on consumer price inflation in the euro area and thus needs to be taken into account by the monetary authorities.

Exchange rate pass-through to consumer prices: A European perspective

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Abstract

We study the pass-through of exchange rate changes to consumer prices for the euro area by estimating vector error correction models for Germany, France, Italy, the Netherlands and Spain. Using the weights of the Harmonized Index of Consumer Prices (HICP) we compute a weighted average of the country results for the euro area. We find that in response to a ten percent depreciation of the euro nominal effective exchange rate index, the HICP tends to increase by 0,4 percent after 12 months. The total effect amounts to 0,8 percent and the adjustment of consumer prices is completed after three years.

Keywords: pass-through, inflation, exchange rate, euro area

JEL Classification: E31, F31, F41

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1 Introduction

The continuous depreciation of the euro since its introduction has posed a challenge for monetary policy in the euro area. In the words of 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. These risks have to be taken seriously in the light of the current strong upswing." (Duisenberg, 2000). This view has been supported by the concerted foreign exchange intervention in September 2000 which 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 very scarce given that the time horizon since the introduction of the euro is rather short. This paper aims to present a first estimate of the effects of a change in the nominal effective exchange rate index for the euro area on the Harmonized Index of Consumer Prices (HICP).

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 US (about 35 percent of all pass-through studies until 1995 concentrated on the US and 7 percent on Germany). Our focus on the European experience tries to fill this gap. Second, many empirical studies used traditional OLS analysis and paid little attention to the time series properties of the data. This also holds for the recent study by Ranki (2000) who focuses on the euro area. We will use up-to-date econometric methods (cointegration, vector error correction models) to contribute to the existing literature. Third, we concentrate on aggregate data rather than on particular industries or products. As most of the literature in the past years has dealt with microeconomic factors (e.g. pricing-to-market studies) our study aims to counter this imbalance. We are primarily interested in the overall effect of exchange rate changes on consumer prices, an issue which is most 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 the euro area is rather short we study exchange rate pass-through for selected European countries 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 Harmonized Index of Consumer Prices.

The paper is structured as follows: section 2 gives a short survey of the relevant literature. Data and econometric procedure 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) 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." (p. 1248). However, changes in import 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 become more important with an increase in the openness of an economy.

The *direct* channel of pass-through runs via the external sector of a country, i.e. through the price of imports. Let E be the exchange rate in terms of domestic currency per unit of foreign currency and P^* the foreign-currency price of the imported good, then $E*P^*$ represents the domestic-currency price of the imported good. If P^* remains fixed and E depreciates (rises) then the domestic-currency price of the imported good will rise in proportion. The result is called the pass-through from the exchange rate to import prices. However, pass-through is only complete (= 100 percent) if (a) markups of prices over costs are constant and (b) marginal costs are constant (see Goldberg and Knetter, 1997: 1248). 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. Since nominal wage contracts are fixed in the short run, real wages will decrease and output will increase. However, when real wages will be bid up to their original level over time, production costs increase, the overall price level increases and output falls. Thus, in the end the exchange rate depreciation leaves a permanent increase in the price level with only a temporary increase in output (see Kahn, 1987).

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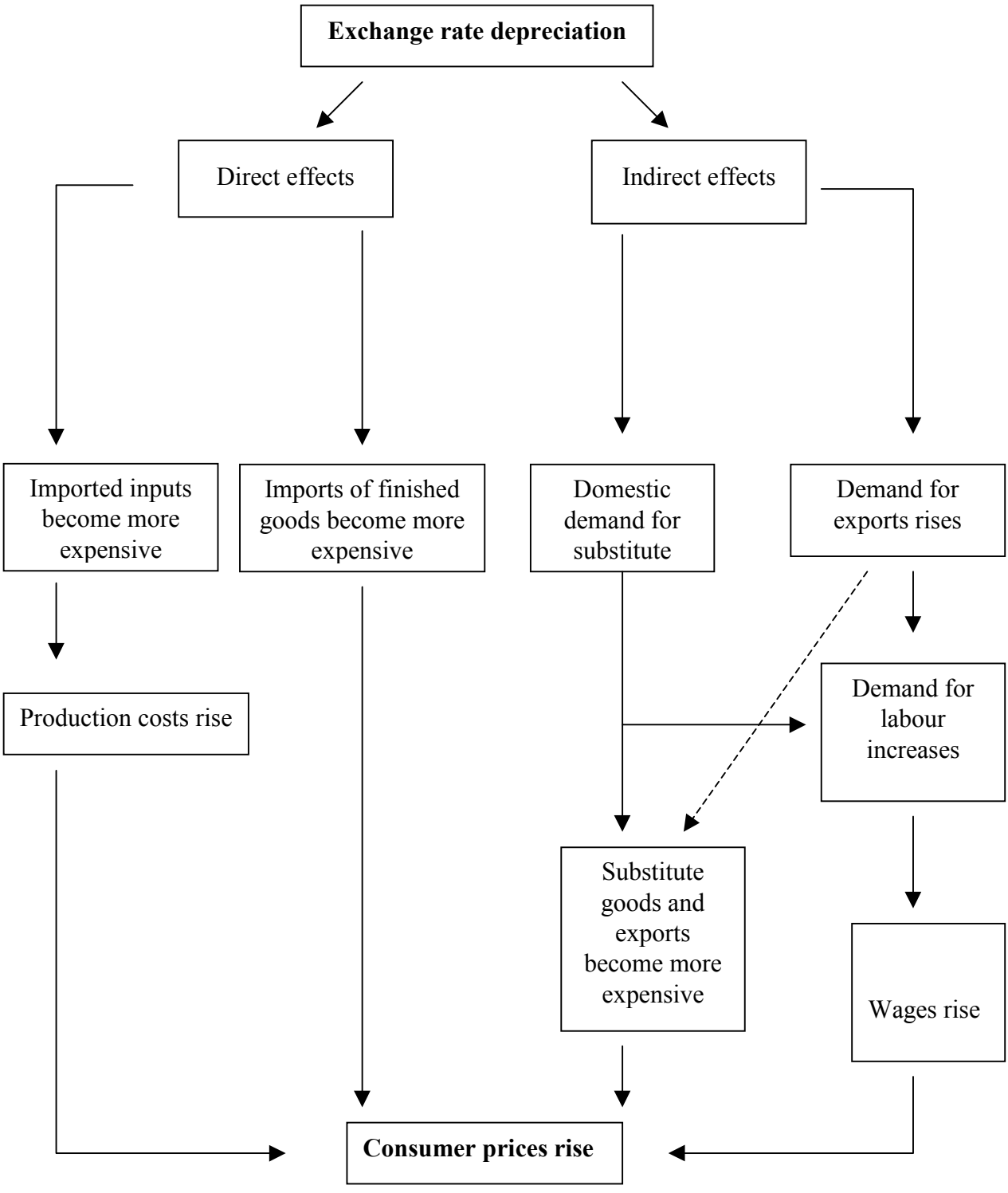
[1] See Kahn (1987), Menon (1995) and Goldberg and Knetter (1997) for an exhausting discussion of exchange rate pass-through. Taylor (2000) suggests a further channel via expectations. According to this view pass-through is highest when exchange rate changes are perceived to be persistent and prices adjust because of the expectations of the public.

Figure 1 displays the mechanism of direct and indirect exchange rate pass-through effects.

In reality, however, exchange rate pass-through is far from complete. 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. 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. Such a behavior is called pricing-to-market.[2] It means that exporting firms accept temporal losses in their revenues in order to avoid long-run losses (in case of an appreciation of the own currency) of their market share. Many studies find that the extent of pricing-to-market is positively correlated with market concentration: pricing–to–market tends to be more present in competitive industries.

[2] See Krugman (1987) and Dornbusch (1987).

Figure 1: Pass-through from an exchange rate depreciation to consumer prices



Source: Laflèche (1996).

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. 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 US 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 specification and variable selection rather than from different time periods studied.

Menon criticizes 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 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 (i.e. following Menon, 1995) there has been some empirical work on exchange rate pass-through that tried to improve the deficiencies of earlier studies that were identified by Menon (1995). A comprehensive study is McCarthy (2000), who investigates exchange rate pass-through on the aggregate level for selected industrialized 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 analyzed 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 US using cointegration analysis and a vector error correction model (VECM). His paper relates producer price inflation in the US 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.

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[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.

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 earlier studies is the neglect of the time-series properties of the data, particularly non-stationarity.

Ranki (2000), to our knowledge, is the only study so far that examines exchange rate pass-through for the whole euro area applying 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 (the 25 percent depreciation of the euro in the first two years of his existence did not translate in 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 up to now showed that there are indeed a number of issues that leave space for further research. Our focus on the European experience tries to contribute to the lack of sufficient country coverage. Second, many empirical studies use 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), and in contrast to McCarthy (2000) who applies VAR analysis, 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) 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.” (p. 1148).

3 Data and Econometric Procedure

3.1 Data description

For each country (France, Germany, Italy, the Netherlands and Spain) we use seven variables in the estimations: three price indices (import prices, producer prices, consumer prices[4]), output gap (constructed using industrial production), short-term interest rate, oil price and the effective nominal exchange rate. All data have 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 1 for details). The time span covered is January 1982 until January 2001. 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 on. Thus, we use this date as a starting date for our whole sample. Exchange rate data are effective nominal exchange rates of the national currencies which use the trade weights of each country. After the introduction of the euro the share of trade with other EMU-countries is weighted with the fixed conversion values. The oil price is the petroleum spot price of UK Brent. We use the national call money rates to approximate central bank behavior. The output gap is computed as the difference between actual industrial production and potential output (constructed with 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 (import, producer and consumer prices) in our estimation. This gives us the opportunity to analyze 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 behavior should distort the true consequences of exchange rate variations. That way, the observed relationship

[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 on (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.

between prices and exchange rates would take into account the central bank behavior rather than the direct influence of exchange rates on prices.[5]

3.2 Unit root tests

In order to determine whether the variables exhibit non-stationary behavior we perform unit root tests.[6] We start using the KPSS-Test which tests the null hypothesis "stationarity" against a unit root alternative.[7] The test uses the regression of the time series to be analyzed (Y_t) against a constant ("stationarity") or a constant and a time trend ("trend stationarity"):

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

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) analyzed 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 generalized KPSS procedure to test for stationarity. The test statistics for the two regressions ((a) only with a constant and (b) with constant and trend), the chosen lag length, and the conclusion concerning the degree of integration are given in Tables 1 – 5 in Appendix 3.

The results can be summarized 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 and the import price index in France and Spain the null hypothesis of stationarity could not be rejected. Therefore, we also conducted an ADF unit root test which has the null hypothesis "non-stationarity" for these three variables.

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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.

[7] See Kwiatkowski et al. (1992).

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 AIC criterion plus two lags. The results are given in Table 6 in Appendix 2. 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 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 consequences for the model for Spain will be discussed in the next section.

3.3 Cointegration tests

In the main part of our analysis we construct a vector error correction (VEC)-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 the 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 estimate in the 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.[8] As a consequence the cointegration rank increases by the number of stationary variables.[9] The correct number of cointegrating equations (CE) 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.

[8] 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 perform a cointegration test.

[9] See e.g. Hansen and Juselius (1995).

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	1	Constant in CE and VAR
Germany	3	1	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

Notes: [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. For Spain the number of cointegrating equations could be 1 or 2 according to the number of stationary variables (see Appendix 3, Table 11). We have chosen 2 CE because both were significant in the VEC-model.

Table 1 summarizes the results of the Johansen tests. The detailed analysis for each country is shown in the Tables 7 – 11 in Appendix 4. We have found 2 or 3 cointegrating vectors for each country. The correct number of CE after subtracting the number of stationary variables is one for France, Germany and Italy and two for the Netherlands and Spain. In case of Spain, as shown in the previous section, the unit root tests were inconclusive about the number of stationary variables. As a consequence, the number of CE could be 1 or 2 for the VEC-model of Spain. We decided to include 2 cointegrating equations in this VEC-model as both were significant in the estimation.

Table 1 also shows that the specification of the VEC-models is very 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 the linear time trend in the cointegrating equation was significant. These VEC-models are the basis for our impulse-response analysis for which the results are described in the following 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:

$$OIL \rightarrow EX \rightarrow IMP \rightarrow GAP \rightarrow INT \rightarrow PPI \rightarrow CPI$$

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 between 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 behavior 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 realized inflation (forward-looking behavior).[10] 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.

Table 2: 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	0,01	0,07	0,12	0,16
Germany	0,07	0,08	0,09	0,10
Italy	0,06	0,12	0,16	0,18
Netherlands	0,12	0,11	0,11	0,11
Spain	0,09	0,08	0,08	0,08

Note: 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 2 displays the responses of national consumer prices to a 1-percent shock in the national nominal exchange rate indices (an increase corresponds to a depreciation) after 6, 12, 18 and 24 months.[11] 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

[10] See Clarida et. al. (1999).

[11] The response patterns of consumer prices to a one-standard deviation shock in effective exchange rates are shown in appendix 2 for each country separately.

effect can be observed in the Netherlands with an consumer price index increase of 0,12 percent after 6 months. In the long run, pass-through is highest in France and Italy with a response of 0,16 and 0,18 percent after two years, respectively. In Spain only 8 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 VECM, variance decomposition separates the variation in an endogenous variable into the component shocks to the VECM. Thus, the variance decomposition provides information about the relative importance of each random innovation in affecting the variables in the system. Table 3 displays the variance decomposition of import, producer and consumer prices to a shock in the effective exchange rate after 6, 12 and 24 months.

Table 3: 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	31,7	30,3	28,9	1,4	4,1	6,9	0,2	1,4	5,2
Germany	35,3	39,5	38,2	16,8	17,7	14,8	1,6	1,9	2,3
Italy	20,3	21,8	22,1	8,1	9,5	9,4	1,2	2,2	2,9
Netherlands	33,2	36,1	37,9	21,9	22,6	20,6	8,3	6,7	4,5
Spain	13,9	15,3	15,4	13,5	16,3	14,9	4,4	3,7	2,3

Note: 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 the producer prices is significantly smaller in all countries. 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 15 percent 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 almost as large as on the import prices. The largest effect on producer prices is found for the Netherlands, where about 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 Netherlands exhibit the largest effect with 5,2 and 4,5 percent after two years. In the other countries in the long run only between 2 percent and 3 percent of the variance of consumer prices can be explained by exchange rate shocks. 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 3 give 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 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 4 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 3 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. This and the relatively large share of imports from non-European countries (e.g. the United States) helps to explain the stronger impact of exchange rate fluctuations on the variance of the Dutch prices.

Table 4: 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%

Note: 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. Sources: IFS, OECD.

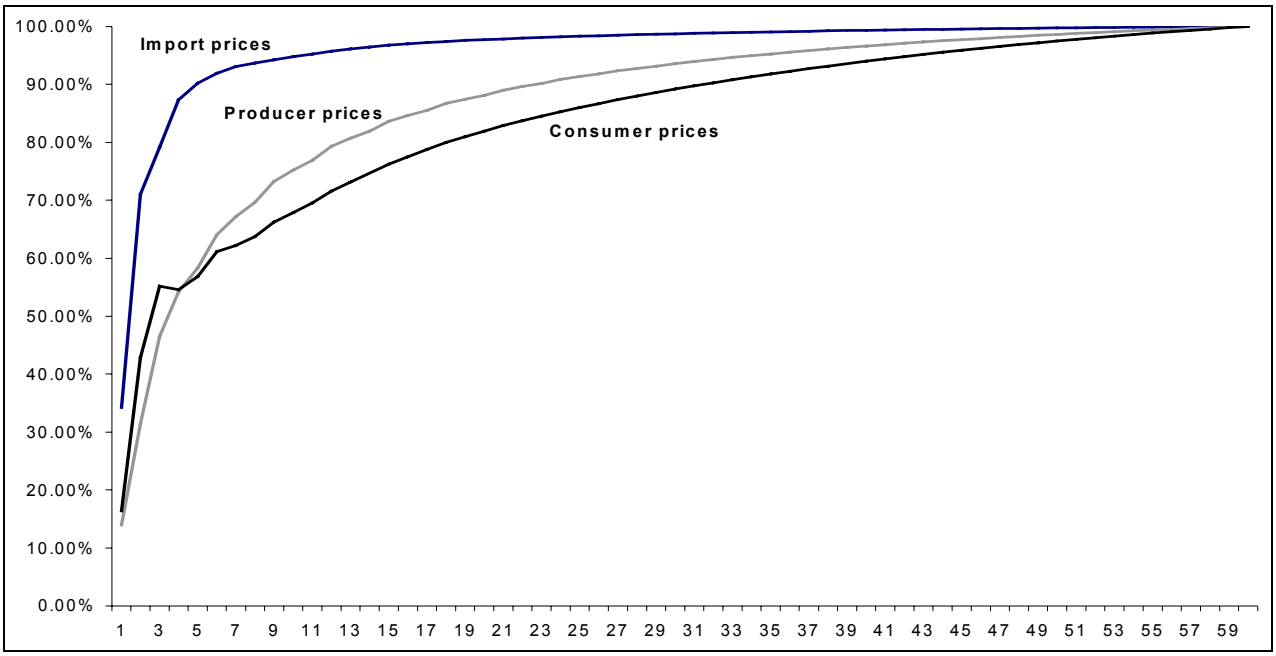
Figure 2 compares the average speed of price adjustment for import, producer and consumer prices after an exchange rate shock. The speed of adjustment is defined as the ratio of the price response after t periods relative to the long run response in percent. As expected the adjustment period is very 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 90 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 about 97 percent of the adjustment in import prices has materialized.

This difference can be explained with 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 consumer price inflation rate only slowly.

Figure 2: Speed of price adjustment after an initial exchange rate shock



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

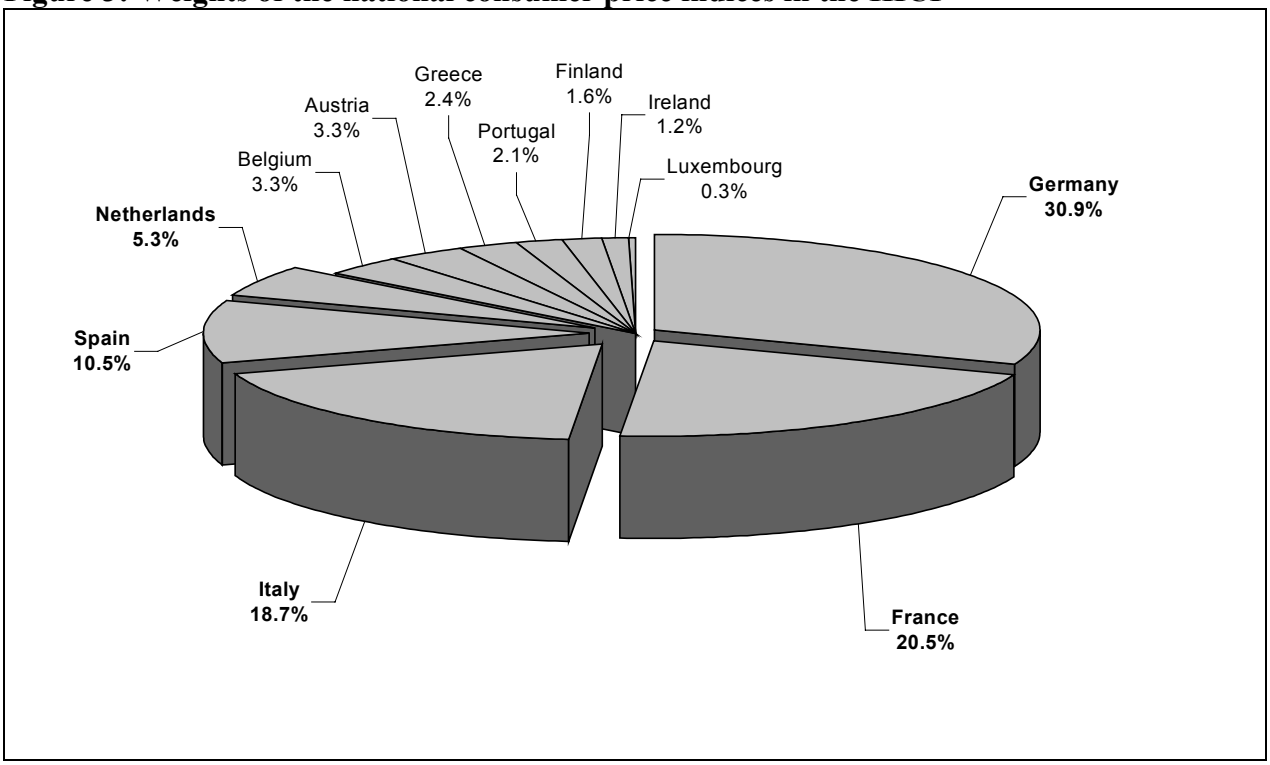
4.2 Aggregation over EMU countries

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 European Central Bank is only concerned with the aggregated inflation rate rather than the 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 weight of the national inflation rates in the calculation of the Harmonized Index of Consumer Prices has to be taken into account.[12] The five countries in our sample represent 86 per cent of the weight of the HICP (see Figure 3).[13] 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.

Figure 3: Weights of the national consumer price indices in the HICP



Source: Eurostat new release 28 February 2001.

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

[12] 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.

[13] The weights for the individual country are determined as if all five countries would represent 100 percent. Thus, the adjusted shares with which our country-specific results are weighted are: Germany 36%, France 24%, Italy 22%, Netherlands 6%, Spain 12%.

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

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

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 are replaced with 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 (1) are set to zero.[14] Therefore, the national effective exchange rate indices and the euro effective exchange rate (E) are related according to formula (2), where the weights are defined as in (1):

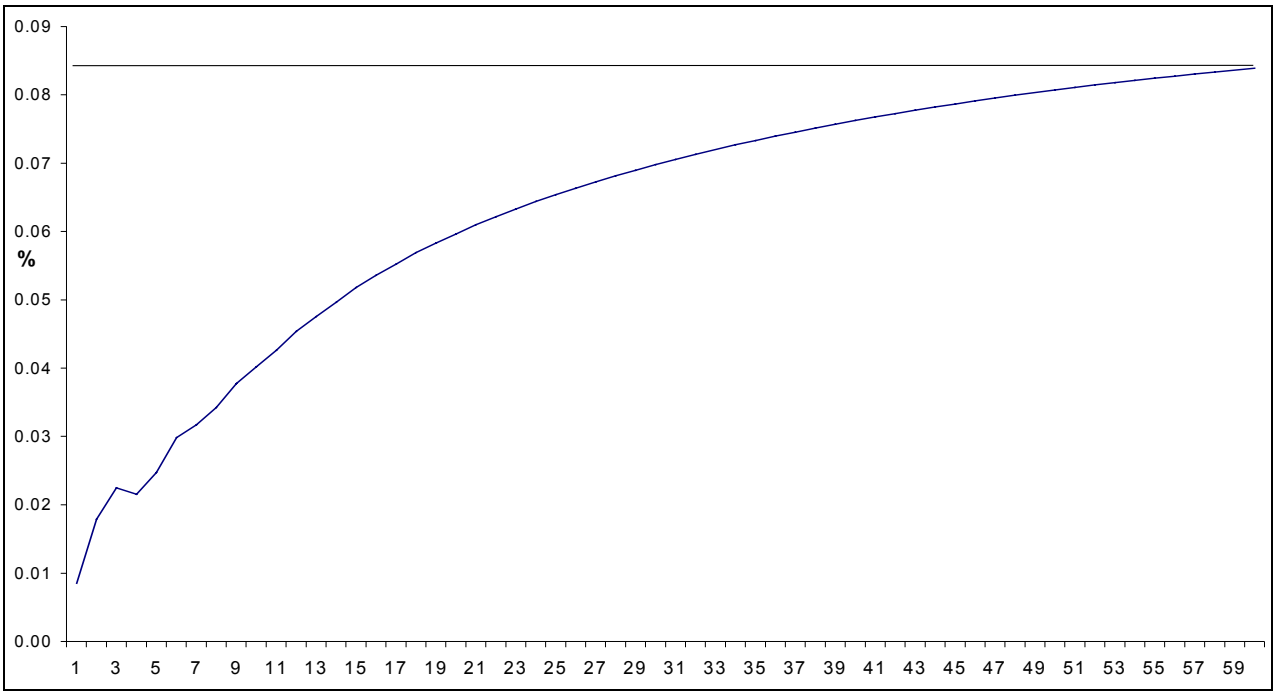
$$(2)\quad \Delta E_N = \Delta E \cdot (1 - \sum_j \beta_j) = \Delta E \cdot \sum_i \alpha_i$$

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 in 1999 was conducted with countries outside the euro-zone. Applying the weighting scheme for the euro effective exchange rate index of the ECB results in a weight of 49,5 percent.[15]

[14] See ECB Monthly Bulletin, methodological notes (available at www.ecb.de).

[15] 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 from third countries (see ECB Monthly Bulletin, methodological notes, available at www.ecb.de).

Figure 4: Effects of a 1% depreciation of the euro effective exchange rate on the HICP



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

Figure 4 shows the cumulative effect on the European price level as measured by the HICP. One year after the nominal effective euro index depreciates by one percent, consumer prices in the euro area increase by 0,04 percent. The total effect converges at about 0,08 percent. Accordingly, a ten percent shock amounts to a pass-through to consumer prices of 0,8 percent. Three years after the initial shock about 88 percent of the total adjustment have occurred. These results are comparable with the findings of Kahn (1987) for the US. He finds that after a ten percent depreciation of the US$ consumer prices in the US tend to increase by about 0,84 percent after nine quarters. As the euro area is often compared with the US in terms of openness and largeness of the economy, the similarity of our results with Kahns' findings are reassuring this opinion.

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 with other factors like e.g. the oil price increase.

Conclusion

In this paper we analyzed the effects of exchange rate fluctuations of the euro on the Harmonized 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, but the long run effects are highest in Italy and France. Pass-through coefficients, i.e. the share of exchange rate change that is reflected in consumer prices, ranges from 7 (France) to 12 percent (Italy) after one year. After two years, coefficients range from 8 (Spain) to 18 percent (Italy).

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 0,8 percentage points after about three years. This amounts to an exchange rate pass-through to consumer prices of 8 percent of the initial exchange rate shock.

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, by 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 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 really stable over time or if exchange rate pass-through has changed significantly during the period studied.[16] In particular we do not know if the introduction of the euro has changed

[16] The literature on this subject presents mixed results: Menon (1995) reports that only quite 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

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 treated 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 VECM 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.

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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 used in logs.

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.[17]

Short term interest rates (INT): Money market rates (Source: IFS line 60B..ZF...), period average. From the start of EMU (1999:1) on 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 / Spain* – Import price index (Source: IFS line 76.X.ZF...), 1995=100.

France – Raw materials import prices, nadj. (Source: Datastream code FRIMIRAWF).

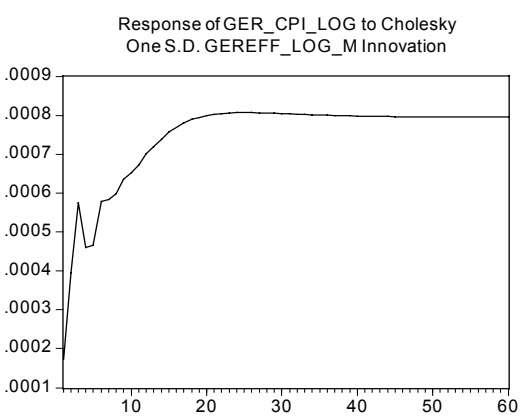
Producer prices (PPI): *Germany / Netherlands / Spain / Italy* – PPI (Source: IFS line 63...ZF...)

France – PPI Intermediate goods for industry nadj. (Source: Datastream code FRPPIINTF).

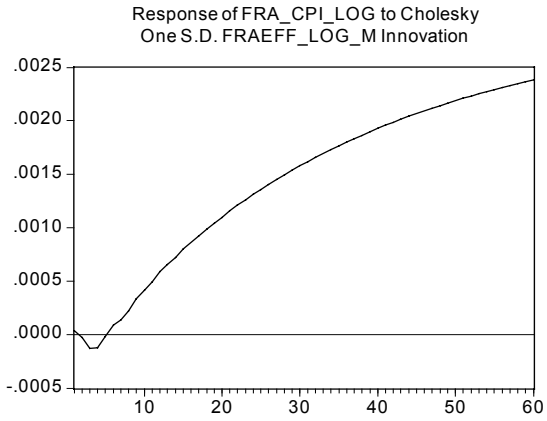
Consumer prices (CPI): CPI index (Source: IFS line 64...ZF...).

[17] 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 www.bankofengland.co.uk).

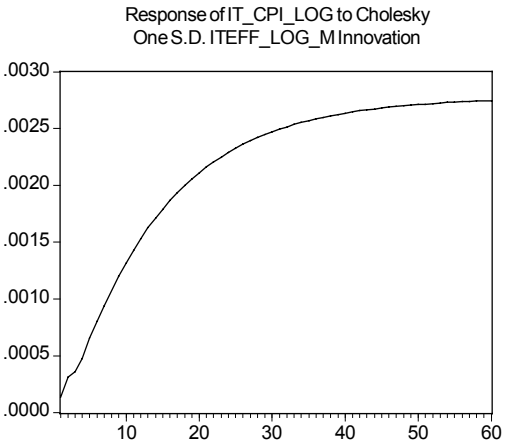
2. Response of national consumer prices to a one-standard deviation exchange rate shock



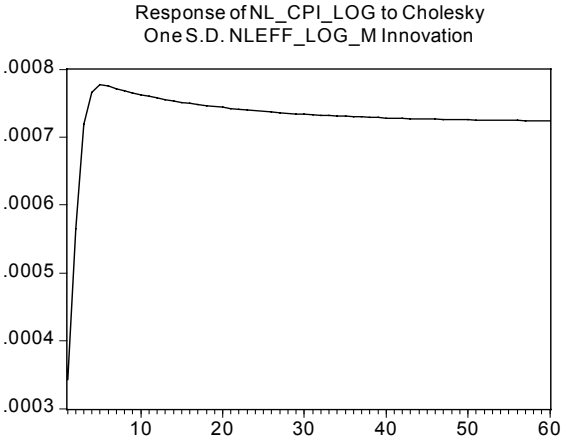
Germany



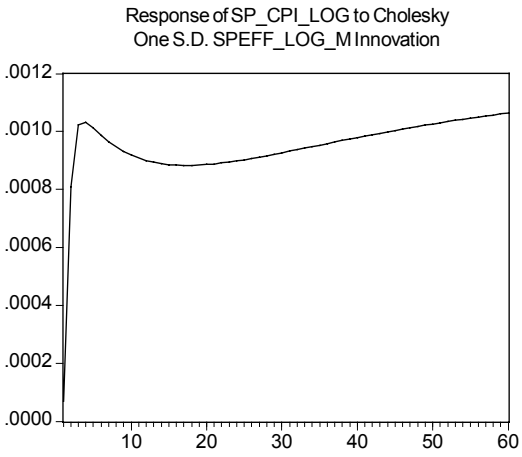
France



Italy



The Netherlands



Spain

3. Results of the Unit Root Tests

Table 5: Results of the KPSS-Test for Germany (All Variables in Logs)

Jan. 1982 – Dec. 2000	Optimal Lag[1]	KPSS with Constant[2]	KPSS with Trend[2]	Degree of Integration[8]
CPI[3]	11	1.99****	0.22****	I(1)
PPI[4]	11	1.84****	0.11	I(0)?
Import Prices[5]	11	0.92****	0.20***	I(1)
Exchange Rate[6]	11	1.53****	0.38****	I(1)
Interest Rate[7]	11	0.58***	0.25****	I(1)
Oil Price[9]	11	0.55**	0.19***	I(1)

Table 6: Results of the KPSS-Test for France (All Variables in Logs)

Jan. 1982 – Dec. 2000	Optimal Lag[1]	KPSS with Constant[2]	KPSS with Trend[2]	Degree of Integration[8]
CPI[3]	11	1.89****	0.43****	I(1)
PPI[4]	11	1.17****	0.35****	I(1)
Import Prices[5]	11	0.10	0.07	I(0)?
Exchange Rate[6]	11	1.08****	0.19***	I(1)
Interest Rate[7]	11	1.54****	0.23****	I(1)

Table 7: Results of the KPSS-Test for Italy (All Variables in Logs)

Jan. 1982 – Dec. 2000	Optimal Lag[1]	KPSS with Constant[2]	KPSS with Trend[2]	Degree of Integration[8]
CPI[3]	11	1.95****	0.43****	I(1)
PPI[4]	11	1.96****	0.25****	I(1)
Import Prices[5]	11	1.58****	0.21***	I(1)
Exchange Rate[6]	11	1.79****	0.15**	I(1)
Interest Rate[7]	11	1.55****	0.25****	I(1)

Table 8: Results of the KPSS-Test for Spain (All Variables in Logs)

Jan. 1982 – Dec. 2000	Optimal Lag[1]	KPSS with Constant[2]	KPSS with Trend[2]	Degree of Integration[8]
CPI[3]	11	1.95****	0.46****	I(1)
PPI[4]	11	1.86****	0.25****	I(1)
Import Prices[5]	11	0.27	0.26****	I(0)?
Exchange Rate[6]	11	1.51****	0.20***	I(1)
Interest Rate[7]	11	1.47****	0.34****	I(1)

Table 9: Results of the KPSS-Test for the Netherlands (All Variables in Logs)

Jan. 1982 – Dec. 2000	Optimal Lag[1]	KPSS with Constant[2]	KPSS with Trend[2]	Degree of Integration[8]
CPI[3]	11	1.98****	0.33****	I(1)
PPI[4]	11	1.91****	0.13*	I(1)
Import Prices[5]	11	1.01****	0.27****	I(1)
Exchange Rate[6]	11	1.43****	0.36****	I(1)
Interest Rate[7]	11	0.88****	0.27****	I(1)

Notes: [1]Optimal lag according to the automatic lag selection procedure developed in Newey/West (1994). [2] KPSS-test statistic and significance level: * = 10%, ** = 5%, *** = 2.5%, **** = 1%. The KPSS-tests with trend and constant and with only a constant have different critical values. See Kwiatkowski et al. (1992). [3] Consumer Price Index. [4] Producer Price Index. [5] Import Price Index. [6] Effective exchange rate. [7] Call money rate. [8] I(1) = alternative hypothesis, I(0) = null hypothesis. The result is I(1) if both tests reject the null hypothesis. [9] The oil price is measured in US$ and is the same for all country models.

Table 10: Additional ADF-Tests for the I(0)?-Variables of the Tables 1 – 5. (All Variables in Logs)

Jan. 1982 – Dec. 2000	Lags[1]	ADF (Trend + Constant)[2]	ADF (Constant)[2]	ADF (None)[2]	Degree of Integration[3]
PPI Germany[4]	5	-3.171*	-1.197	1.567	I(0)
Import Prices France[5]	3	-2.710	-2.705*	0.314	I(0)
Import Prices Spain[5]	3	-2.034	-2.052	1.131	I(1)?

Notes: [1] The lag length is specified using the AIC2-rule of Pantula et al. (1994) = minimum of AIC plus two lags. [2] ADF test statistic and significance level: * = 10%, ** = 5%, *** = 1%. [3] I(0) = alternative hypothesis, I(1) = null hypothesis. The result is I(0) if at least one test rejects the null hypothesis. [4] Producer Price Index. [5] Import Price Index. [6] Call money rate, Period: Jan. 1982 – Dec. 1996.

4. Results of the Cointegration Tests (Johansen Test)

Table 11: Results of the Cointegration Tests for France (Johansen Test)

Jan. 1982 – Nov. 2000	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 – 2 = 1
At most 3	42.72	
At most 4	20.33	

Notes: Seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate, Lag length = 3 (determined by AIC). Significance levels: ** = 1%, * = 5%. [1] A linear trend in the cointegrating equations (CE) was not significant, [2] Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 2).

Table 12: Results of the Cointegration Tests for Germany (Johansen Test)

Jan. 1982 – Dec. 2000	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 – 2 = 1
At most 3	35.37	
At most 4	15.98	

Notes: Seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate, Lag length = 3 (determined by AIC). Significance levels: ** = 1%, * = 5%. [1] A linear trend in the cointegrating equations (CE) was not significant. [2] Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 2).

Table 13: Results of the Cointegration Tests for Italy (Johansen Test)

Jan. 1982 – July 2000	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	

Notes: Seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate, Lag length = 2 (determined by AIC). Significance levels: ** = 1%, * = 5%. [1.] A linear trend was significant in the cointegrating equations. [2.] Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 1).

Table 14: Results of the Cointegration Tests for The Netherlands (Johansen Test)

Jan. 1982 – Dec. 2000	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	

Notes: Seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate, Lag length = 1 (determined by AIC). Significance levels: ** = 1%, * = 5%. [1.] A linear trend in the cointegrating equations (CE) was not significant. [2.] Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 1).

Table 15: Results of the Cointegration Tests for Spain (Johansen Test)

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

Notes: Seven variables (all in logs, except gap): consumer price index, producer price index, import price index, interest rate, gap, oil price, effective exchange rate, Lag length = 1 (determined by AIC). Significance levels: ** = 1%, * = 5%. [1.] A linear trend in the cointegrating equations (CE) was not significant. [2]. Number of cointegrating equations according to column 2 minus the number of stationary variables in the system (= 1 or 2).