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Abstract

The paper discusses why the previous literature has found little evidence of any effect of exchange rate variability on international trade; methodological and statistical issues are discussed. In particular, we compare estimations based on different specifications and using different datasets and show how the results may change depending on the method used. When econometric analysis is conducted accurately, it is found that, indeed, exchange rate variability has had a significant negative impact on trade. The analysis is conducted for manufacturing exports of Italy, France and Germany towards their EMS partners.

Contents

1. Introduction[1]

Although exchange rate variability has often been considered as one of the major shortcomings of the flexible exchange rate regime, since it increases the uncertainty underlying international trade and financial transactions[2] both in the short and long run, the empirical literature has not found decisive evidence of a negative impact of short term volatility on international trade.

One of the first theoretical contributions to the analysis of the effect of short term exchange rate risk on international trade was made by Ethier (1973). He showed, using a model based on a mean-variance specification of the expected utility function of the representative firm engaged in international trade, that the level of trade tends to fall as exchange rate uncertainty rises. Hooper and Kohlhagen (1978) extended this model to take into account monopolistic competition. They relaxed the assumption of infinite elasticity of the exporters' supply function, and showed that exchange rate variability has a negative impact on the volume

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1. Previous versions of this paper were presented at seminars at the University of Chicago and at a joint workshop CEPS-Université Catholique de Louvain. I wish to thank the participants, and in particular J. Frenkel, J. Aizenman and J. Huizinga for their helpful comments. I am the sole responsible for errors and opinions. The data are available from the author on request.

2. See Group of Ten, 1985, p.9: "The Deputies have noted that short term exchange rate variability has been substantial and has not shown any tendency to diminish over time. Although empirical studies conducted by the IMF have been unable to find a significant systematic link between short term exchange rate volatility and the volume of international trade, concern has been expressed that volatility may discourae investment and trade by adding to uncertainty and to financial risks for investors and traders".

of trade but an indeterminate effect on prices.[3]

Several authors attempted to verify the empirical validity of this hypothesis.[4] Overall, the results seem to be quite contradictory, but tend in general to point out that there is no systematic significant relationship between exchange rate variability and trade flows.

Two interpretations can be made of these results. A first is that, indeed, exporters and importers do make extensive use of forward hedging, and that the cost of this hedging is in fact relatively small[5] and therefore not relevant for international trade. However, this conclusion contrasts with other empirical evidence based on surveys or other analysis which suggest that on aggregate firms do not entirely hedge against foreign exchange risk.[6] A second interpretation is that the empirical testing of the relationship between exchange rate variability and international trade is undermined by a series of methodological problems that may lead to imprecise statistical results. In this paper we concentrate on the latter; we examine the major empirical problems encountered in the previous literature, and assess some of the reasons why the hypothetical relationship between exchange rate risk and international trade could not be fully detected.

In the next section we first briefly sketch the

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3. More recent research has concentrated on the effect of longer run exchange rate changes in trade flows, advancing the hypothesis of hysteresis (Baldwin, 1988).

4. Among others, Makin (1976), Hooper and Kohlhagen (1978), Cushman (1983), IMF (1984), Justice (1983), Akhtar and Hilton (1984), Kenen and Rodrick (1986), Gotur (1985), Bailey and others (1986), Kenen (1979), Thursby and Thursby (1985).

5. Alternatively, the cost might be of a fixed nature and not affect the quantity exchanged.

6. Magee and Rao (1980), Grassman (1973, 1976), Jilling (1977).

underlying theoretical framework. In the third we present several estimates of the relationship between exchange rate variability and trade volumes and prices and show how the results may change depending on the data used or on the specification of the econometric testing. The results suggest that the hypothesis of a negative relationship between exchange rate variability and trade volume cannot be rejected; the analysis also attempts to explain why previous research has not reached the same conclusions.

2. The model

The model underlying our empirical analysis is that of Hooper and Kohlhagen (henceforth H-K, 1978). Its main feature is that foreign exchange uncertainty is assumed to be the only source of risk for the economy: unexpected variations of the spot exchange rate affect the unhedged profit stream of firms because of the time lag between the contract and the payment date. The representative importer is assumed to maximize its utility (u), that depends (positively) on the expected value of profits (π) and (negatively) on the standard deviation of these profits:

$$(1) \qquad \text{Max } u = E(\pi) - \gamma\sigma_{\pi}$$

where E(.) is the expectation operator, σ the standard deviation and γ is the measure of relative risk aversion. The importer is assumed to take import prices as given, and to set the domestic price of the imported good taking into account the demand curve faced on the domestic market. The demand schedule depends on income and on the price of competitors on the domestic market. The cost of the imported product depends on the price of the foreign currency and on the way the contract is stipulated:

$$(2) \qquad PM = P_x^* \{ \beta(\alpha R_f + (1-\alpha) R) + (1-\beta) R_f \}; \quad 0<\alpha,\beta<1$$

where P_x^* is the price of imports in foreign currency. The term in brackets represents the average cost of foreign exchange for the importer. β is the proportion of imports denominated in the exporter's currency; α is the proportion of contracts denominated in foreign currency. The importer hedges a proportion $\alpha\beta$ of the contract in the forward market, at a (known) rate R_f. The other portion of the contract $(1-\alpha)\beta$ is instead unhedged, and its value depends on the unknown exchange rate R. Finally, the proportion $(1-\beta)$ is denominated in domestic currency and is therefore automatically hedged by the importer.[7]

The above specification is designed to aggregate the various types of existing contracts. Equation (2) shows that there are two ways of fully hedging against unexpected exchange rate changes. A first would be to buy the entire amount of foreign exchange forward (α=1), so that the overall price is already known at the time of the contract. A second is to denominate the contract in domestic currency (β=0), so that the amount to be paid is invariant to exchange rate changes. By setting α less than one it is implicitly assumed that (on aggregate) firms do not have perfect access or do not fully use forward markets to hedge against foreign exchange risk. This may be due, for instance, to: (i) the imperfection of forward markets; (ii) institutional regulations that limit the amount that can be hedged; (iii) transaction or information costs that reduce the (aggregate) amount of hedging.

In the model it is assumed that the proportion of contracts hedged through the forward market and that denominated in domestic currency are exogenous and not

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7. It is assumed by H-K (1978) that the price in domestic currency is equal to the foreign price times the forward exchange rate.

derived from a utility maximization framework. Further, from the empirical point of view the assumption of constant coefficients α and β might be inappropriate if the period examined includes changes in policy regimes; a move towards more exchange rate flexibility would, for instance, tend to increase the propensity to hedge against foreign exchange risk.

Given the assumption of infinitely elastic export supply, H-K (1978) derive the following export demand function:

$$(3) \qquad X^d = X(\underset{-}{CX}, \underset{+}{Y}, \underset{-}{V})$$

where CX is the export competitiveness, calculated as the ratio between export prices (PX) and the price of competitors on the domestic market (PC), Y is the level of income in the importing country, V is the exchange rate risk. The volume of exports is positively related to income, and negatively related to competitiveness and exchange risk.

If the assumption is made that the exporter has some market power export prices and quantities will be determined simultaneously. The following reduced form can be derived:

$$(4) \qquad \begin{aligned} X &= X(\underset{+}{Y}, \underset{-}{PI}, \underset{+}{PC}, \underset{-}{V}) \\ PX &= P(\underset{+}{Y}, \underset{+}{PI}, \underset{+}{PC}, \underset{\pm}{V}) \end{aligned}$$

where PI are the exporter's production costs. The export price level in equations (4) depends positively on the level of income in the market of destination, on the production costs of the exporter and on the price of competitors. If the (representative) exporter has little market power, the export price will tend to follow closely those of the competitors and the coefficient on PC will be close to unity, while that on PI will be close to zero (Spencer, 1984). If, on the

contrary, the firm has some market power, it will succeed in transferring a large part of its production costs on prices and the coefficient on PI will be close to one. The risk variable enters the reduced form price equation with an indeterminate sign. If the cost of covering against exchange rate risk is borne mostly by the importer, an increase in risk will tend to reduce the demand for exports, thereby reducing the export price level. If the risk is borne mostly by the exporter, a rise in exchange rate variability will tend to reduce the quantity supplied, and therefore push export prices up.

3. Methodological issues in the estimation of the empirical relationship

In this section we test the hypothesis that short term exchange rate variability has an effect on trade volumes and prices by estimating equations (3) and (4) for manufacturing trade. We first discuss the general approach followed in estimating the above relationship, and then compare it with that of previous research and assess the difference in the results obtained.

3.1. The general estimation procedure

The data utilized in this research have been obtained from the Volimex (EEC) data bank which allow to reconstruct a country's trade volumes and prices towards a specific area.[8] We have considered the manufacturing exports of three countries, Germany, France and Italy, towards the

8. The data is available on a yearly basis, but its frequency has been extended to a quarterly basis using overall trade prices and volumes as reference series (see appendix). See Bini Smaghi and Vona (1989).

group of EMS countries.[9] We chose to restrict the analysis to intra-EMS trade for two main reasons: (i) for these countries the data are more accurate; (ii) in the period considered (1976-84) no major changes have occurred in commercial policy within that area, since the countries considered have been EEC members from the start; (iii) these countries are highly integrated and there is less scope for firms to undertake differenciated trade practices.

As is usually done for export equations we have used relative prices as the independent variable measuring competitiveness.[10] The income variable is the aggregate demand of the EMS area (excluding the exporting country). The exchange rate risk variable is the standard deviation of weekly rates of changes of the intra-EMS effective exchange rate, for the quarter considered.[11] The estimation procedure is ordinary least squares, for the period 1976-84.

Line A on the first three tables gives, for the three countries considered, the results of the estimation of equation (3), which represents the export volume of manufactures towards the EMS area in the hypothesis of

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9. Only the countries participating to the exchange rate agreement are considered, i.e. Germany, France, Italy, Belgium, the Netherlands and Denmark. Ireland has been omitted for the unavailability of certain data.

10. It is constructed as the ratio of the manufacturing export price towards the EMS area and the price of the competitors on that area. The latter is a weighted average of three variables: (i) the manufacturing wholesale price of domestic producers of the area; (ii) the manufacturing export prices of the other EMS countries towards that area, and; (iii) the manufacturing export price of the rest of the world towards that same area (see Appendix). The weighting scheme used to calculate this variable takes therefore into account the competition not only with domestic producers but also with exporters of other countries.

11. De Grauwe (1987) and De Grauwe and Verfaille (1988) concentrated on the effects of long run real exchange rate variability; the results indicate that the latter had a negative impact on intra-EMS trade volumes.

infinite supply elasticity. Different dynamic specifications have been undertaken.

For France and Italy the estimation in logarithmic first differences proved to be the most satisfactory from the statistical point of view. The F-test proposed by Hendry and Mizon (1978) does not reject the hypothesis of a unit root in the French and Italian export volumes.[12] For Germany instead the specification in (logarithmic) levels presents the best results in terms of t statistics of all independent variables and no autocorrelation of the residuals (see below). The different specification of Germany's exports, with respect to the other two countries, is confirmed by other studies, such as Spencer (1984). For France and Italy relative prices have an effect on volumes within two quarters, while in Germany the lag is longer; this different dynamic reaction is also confirmed by previous research (Spencer, 1984). The income variable has an instantaneous effect on export volumes in Italy and Germany, and with a one quarter lag in France. Two dummies have been used for each country to capture effects due to dock strikes or outliers in the data; they are highly significant and improve the goodness of fit of the equations.[13] Compared to previous research that used similar data for the same sample period (Bini Smaghi and Vona, 1989), the sum of square residual and the standard error are lower for all three countries.

The coefficient on exchange rate variability has a negative sign for all the three countries considered, and is significant at the 2.5 per cent level for Italy and France and at the 5 per cent level for Germany. The R^2 is lower for Italy and France than for Germany because it refers to the

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12. The value of the test for Italy is F(4.26)=.29 and for France F(4.27)=.56, with a significance level of 4.

13. These dummies cover the second quarter of 1979 and the third of 1981 for Italy, the first quarter of 1977 and the second of 1981 for France, and the third quarter of 1976 and the fourth of 1980 for Germany.

first difference of the export volumes.

Table 4 presents the results for equations (4). The exporter's production costs are proxied by the producer prices on the domestic market. Overall, the results for the volume equations are not significantly different from those of equation (3). For Italy equations (4) perform slightly worse than equation (3), as the coefficient for the domestic producer price is not significant; that for exchange rate variability is similar in the two regressions but less significant in the first. For France the coefficient on exchange rate variability is slightly larger in equations (4) and more significant. For Germany equations (4) perform marginally better than equation (3), which may suggest that exporters of that country have some market power on the EMS area.

This is confirmed by the results of the export price equation.[14] German exporters set their export prices in a way very similar to that of their domestic prices, and take little into account the prices of their competitors on the EMS area. Conversely, Italian exporters follow much more closely their competitors in setting their export prices: the coefficient on the PC variable is .63 against .30 for German exporters. French exporters are in an intermediate position.[15] Exchange rate variability has a significant effect on the export prices of the three countries at the 10 per cent confidence level. For Italy and France the effect is positive, thereby suggesting that a large part of the unhedged exchange rate risk is borne by the exporters of these two countries. For Germany, instead, the coefficient is negative, which suggests that most of the unhedged foreign exchange risk is not borne by German exporters but rather by the importers of the other countries of the EMS area. This

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14. The income variable has been omitted in all three cases because not significant.

15. These results are in line with those of Bini Smaghi and Vona (1989).

seems to be consistent with the currency denominations of exports in these countries: while in Germany 80 per cent of exports are denominated in DM, only 62 per cent of French exports are priced in French francs and 32 per cent of Italian exports in lire.

3.2. Main differences with previous research

a) The data

As previously noted, the dataset used in this paper is relatively new, constructed by the EEC Statistical office on the basis of UN trade data. It provides sectoral and geographical disaggregated data for export values and volumes, which are particularly useful for the purpose of our analysis. Previous research often ignored the problem of sectoral disaggregation: H-K (1978), Cushman (1983), Makin (1976), IMF (1984) and Bailey and others (1986) examined the effects of exchange rate variability on overall trade flows. This amounts to constraining the income, price, and exchange rate risk elasticities to be equal across sectors. It has been shown that the aggregation bias may be quite relevant, given the different nature of the markets in which trade occurs, in particular between primary commodities and manufactures (Goldstein and Khan, 1985). The geographical disaggregation is also important, since the behavioural relationships specified in equations (3) and (4) may differ across the various markets of destination, as the results of H-K (1978), Cushman (1983), IMF (1984) and Vona and Bini Smaghi (1988) showed. In particular, trade with less developed countries or centrally planned economies is often made at non-market conditions, with strong state intervention, that affect the contract characteristics and the exporter's exposure to the foreign exchange risk.

However, geographically disaggregated trade volume

data are in general not available; only overall volumes and prices can be obtained for each country. This problem was in part solved in the literature through the construction of "proxy" series of bilateral unit values on the basis of the product composition of bilateral trade. These series would then be used to deflate bilateral trade values. H-K (1978) and Cushman (1983) assumed that exporters charge the same price in all markets and thereby reconstruct the theoretical bilateral export prices from country i to country j by weighting overall export prices by the specific bilateral sectoral trade shares.

These procedures do however introduce a measurement error. Kemp (1962) showed that if there is an error in the measurement of the export prices, used both as a deflator for the export values and as an explanatory variable, the coefficient on the latter, as estimated from equation (3), will be biased towards -1. In equations (4) the bias in the price and volume equations would be of opposite direction.[16] The estimates of all the other coefficients would also be biased. If, for instance, the price series used to deflate the export value is more correlated to exchange rate variability than the true unknown series, the exchange rate variability coefficient will be biased upwards (in absolute terms) in the export price equation and downwards in the export volume equation.

To assess the magnitude of this bias we used for the calculation of export volumes the overall unit values of exports instead of that specific to the country's exports towards the EMS area. For each of the countries considered the new series of unit values were also used in the construction of the competitiveness variable. Line B in tables 1-3 presents the results of the estimation of equation (3) with the new export price data. All other variables

16. This is due to the fact that the price level is used in the volume equation as a deflator for the trade value.

remained unchanged. Comparing line B with line A, it can be seen that the results are generally worse and the price elasticities tend to become less significant. The value of the coefficient for exchange rate variability increases but becomes insignificant at the 5 per cent level for Germany and for France. The use of this data in the empirical analysis would therefore have led to a rejection of the hypothesis that exchange rate variability has a negative effect on trade.

b) Dynamic specification

Previous empirical studies have generally not examined extensively the specification of the trade volume equations. This has led to the result that not only the exchange rate risk variable but also other variables such as competitiveness and income had little significance in explaining trade flows. H-K (1978), Cushman (1983) and IMF (1984) considered only the contemporaneous effect of the independent variables with respect to export volumes, and do not consider any lagged effects. H-K (1978) found that the home price variable has a significant effect on exports only in one out of seven regressions for Germany, and in four out of nine for the US. The foreign price variable was never significant for Germany, and only one out of nine times for the US. In Cushman (1983) the coefficients for income and relative prices were both of the correct sign only in six out of sixteen regressions. It is worth noting that exchange rate variability was found to have a significant effect on export volumes in four of these six cases. IMF (1984) examined 42 bilateral trade flows. Only in 9 cases the income and price elasticities had both a correct and significant sign. The price elasticity had the correct sign and was significant only in 29 cases, and had the wrong sign in 4 cases. From these results one could easily conclude that relative prices

and income do not have any systematic effect on trade flows!

Further, since most regressions denote first order autocorrelation of the residuals, correction of the latter is often conducted without any test on the appropriateness of the procedure. H-K (1978) corrected 13 out of 16 regressions for the first order autocorrelation, the coefficient being often close to unity. Cushman (1983) applied the correction in 6 out of 16 equations, while the IMF (1984) applied no correction and presented DW statistics which are included between 1.70 and 2.30 only in 4 out of 42 cases. Hendry and Mizon (1978) showed that the indiscriminate use of this correction introduces a bias in the results if it is undertaken in substitution for a more sophisticated dynamic specification. As indicated above, the tests conducted for the Italian and French equations suggest that regressing on first differences is more appropriate than on levels.

To verify the effects of the above mentioned problems, we present in tables 1-3, line C, the results for the regression in levels and with no lag structure for the independent variables and correction for first order autocorrelation of the residuals. This specification performs worse than that of line A for all three countries. In the case of Germany and France the coefficient on export competitiveness becomes insignificant. For France and Italy the value of the coefficient on exchange rate variability remains unchanged, and declines for Germany, and in all cases it becomes insignificant at the 5 per cent level (2.5 per cent for Italy), leading to a rejection of the hypothesis that exchange rate variability has an impact on trade. This suggests that the dynamic specification of the equation tested should be closely investigated when addressing the relationship between exchange rate variability and trade.

c) Competitor's price variable

Previous research used as proxy for the competitor's price variable the level of wholesale prices,[17] or the GNP deflator,[18] in the importing country. This implies that the competition with other exporters is omitted from the analysis, thereby introducing a bias in the estimation, since the latter might be correlated with exchange rate variability. In our case the coefficient on exchange rate variability would be biased upward if intra-EMS competitiveness was positively correlated with overall exchange rate variability: a rise in exchange rate variability produces a fall in the price of competitor's and therefore a further fall in the demand for one country's exports.

Line D of tables 1-3 shows the results of the estimation of the export volume equation in which the CX variable has been modified so as to consider only competition with domestic producers, and therefore exclude that with the other countries' exporters. Comparing it with the baseline regression of line A, it can be seen that this specification tend to perform worse for Germany and France. Both the CX and the exchange rate risk variables become insignificant at the 2.5 per cent level. Again, these results point out that the omission of an important variable can lead to an erroneous conclusion on the effect of exchange rate variability on trade. For Italy the results are instead practically unaffected.

17. See for instance H-K (1978).

18. See Cushman (1983).

d) Exchange rate risk variable

Various measures of exchange rate risk have been considered in the literature. Risk should regard nominal rather than real exchange rate risk, contrarily to what has been done by Cushman (1983), Kenen and Rodrick (1986), Justice (1983) and IMF (1984): real exchange rate variability depends not only on the variance of the nominal exchange rate but also on that of relative prices, which presents a different type of risk for private agents.

Several authors[19] used proxies for the difference between the expected and the realized exchange rate, as advocated by Ethier (1973). This variable presents however the disadvantage of being highly correlated with the change of the spot exchange rate, since the forward and the spot rates tend to move closely together (Mussa, 1979); it is therefore also correlated with the change in the real exchange rate and may therefore capture the effects of changes in competitiveness rather than those of exchange rate risk. Furthermore, if the exchange rate behaviour is characterized by wide and frequent unexpected fluctuations around a constant level, the above variable would have all observations close to zero, in spite of the high variability inherent in foreign exchange markets.

Another measure that has been utilized by Akhtar and Hilton (A-H, 1984) and Gotur (1985) is the standard deviation of the level of the exchange rate. This measure is satisfactory only in the hypothesis that the nominal exchange rate fluctuates around a constant level, in the absence of any permanent changes. The experience of the recent years suggests instead that exchange rates tend to display a random-walk behaviour, with "long waves" that have most often

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19. Justice (1983) used the average difference between the forward and spot rate, H-K (1978) the average absolute value of the above difference and Cushman (1983) the first difference of the spot exchange rate.

not been anticipated by the market.

A more satisfactory measure of risk is the standard deviation of the rate of change of the exchange rate. It has the advantage of capturing higher frequency movements of the exchange rate. Most studies have used a moving average transformation to smooth out the series. This involves taking into account lags as long as four quarters (Cushman, 1983; Justice, 1983; IMF, 1984) to eight quarters (Bailey and others, 1986; H-K, 1984; Gotur, 1985; Kenen and Rodrick, 1986). The smoothing of the variability series is not justified for two reasons. The first is that exchange rate variability does not display such long autocorrelation; looking so far back in the past distorts the measurement of the variable. In line E of tables 1-3 we present the results of the estimations when using a four quarters moving average of the variable, as done in the previous literature. For Germany and Italy, the results show that with this specification the exchange rate risk variable looses significance (at the 2.5 per cent level for Germany and at the 10 per cent level for Italy), therefore leading (erroneously) the researcher to conclude that exchange rate variability has no effect on trade. For France the coefficient remains significant (at the 2.5 per cent level), and its size increases by more than three times.

Further, there are grounds to doubt whether this is the correct way to proxy for exchange rate risk. In general, as Pagan and Ullah (1986) assess, "this proxy suffers from a variant of the error in variable problem, given that risk should be measured as a function of the conditional moments of a distribution, and produce an underestimation of the effect of risk on decision. Fundamentally, the concept of risk must be defined in relation to some information set; if perfect prediction could be made, risk would be absent". The authors thereby advocate the use of an instrumental variable technique to construct proxies for risk variables.

We have used this approach to try to decompose

exchange rate variability in an expected and unexpected component, by regressing the contemporaneous values on lagged values of that variable and on the level of the exchange rate. Then, in a second stage, we considered separately the two components of exchange rate variability as explanatory variables in the export volume equations. Line F in tables 1-3 shows the results of the regression. For France and Italy the coefficient on the unexpected component is about twice as large as that on the expected component, and is significant (at the 2.5 per cent level), while the latter has t-statistics lower than one. For Germany the two coefficients have a significant effect on the trade volume. These results suggest that more care should be taken when considering the various measures of the proxy for exchange rate risk.

4. Conclusions

In this paper we discuss why previous research may have been unable to detect any significant correlations between exchange rate variability and trade. We argue that an important shortcoming has been the lack of appropriate data and the insufficient attention dedicated to the specifications of the econometric testing. We show that when the appropriate estimation methodology is used the results are likely to be reversed, and the skepticism generally shared on the significance of the relationship between exchange rate variability and trade vanishes. The results found in this paper suggest that new energies should be invested in the analysis of this relationship and that greater effort should be devoted to the econometric specification of the hypothesis testing.

Table 1

Italy: Results of regressions equation (3)

Specification	Cst	V	CX Σ(0-2)	Y	D_1	D_2	R^2	DW	SSR	SE	$\hat{\rho}$
A	.008 (.15)	-.014 (-2.74)	-.917 (-3.84)	3.29 (4.90)	-.155 (-5.56)	.071 (2.60)	.646	2.21	.020	.026	
B	-.005 (-.85)	-.019 (-2.99)	-.465 (-1.65)	4.17 (5.26)	-.183 (-5.54)	.087 (2.69)	.597	2.20	.027	.030	
C	-4.36 (-1.63)	-.015 (-1.77)	-.629 (-2.30)	2.57 (4.63)	-.062 (-2.15)	.050 (1.53)	.908	1.57	.036	.034	.72 (6.23)
D	.001 (.25)	-.014 (-2.72)	-.914 (-3.95)	3.26 (4.90)	-.156 (-5.63)	.072 (2.63)	.653	2.23	.019	.025	
E	.002 (.32)	-.026 (-1.17)	-.944 (-3.40)	3.06 (4.21)	-.151 (-4.95)	.059 (1.98)	.577	1.88	.024	.028	
F	-.001 (-.21)	$-.018^{U}$ $-.008^{E}$ (-2.68) (-.86)	-.994 (-3.88)	3.27 (4.84)	-.161 (-5.78)	.069 (2.52)	.643	2.18	.019	.026	

A = Regression on (logarithmic) first differences of equation (3); range 1976I-1984IV; B = Same as A but the overall export price index is used, instead of the export price index towards EMS countries, to deflate the export values towards the EMS countries and to calculate the CX variable; C = Regression on levels, with no lagged values of independent variable; D = Same as A but the CX variable includes competition only with domestic producers; E = Same as A but the exchange rate variability is a 4 quarters moving average; F = Same as A but distinction between expected (E) and unexpected (U) components of exchange rate variability; D_1 and D_2 are Dummies for 1979II and 1981III;; Σ(0-2) is the sum of the coefficient of the 1st order polinomial with lags (0-2), except for regression C.

Table 2

France: Results of regressions
equation (3)

Regression	Cst	V	CX $\Sigma(0-2)$	Y $\Sigma(0-2)$	D_1	D_2	R^2	DW	SSR	SE	$\hat{\rho}$
A	-.002 (-.62)	-.009 (-2.18)	-.512 (-2.15)	2.46 (5.20)	-.057 (-2.76)	.050 (2.22)	.531	2.47	.012	.020	
B	-.003 (-.59)	-.009 (-1.72)	-.003 (-.008)	2.56 (4.25)	-.039 (-1.48)	.037 (1.30)	.320	2.67	.019	.026	
C	.128 (.09)	-.009 (-1.77)	-.367 (-1.79)	1.35 (4.38)	-.022 (-1.19)	.017 (.945)	.880	1.96	.013	.022	.68 (5.49)
D	-.003 (-1.60)	-.010 (-1.79)	-.055 (.183)	2.57 (4.26)	-.039 (-1.50)	.040 (1.33)	.421	2.63	.019	.026	
E	-.002 (-.59)	-.029 (-2.41)	-.629 (-2.82)	2.26 (4.87)	-.057 (-2.83)	.047 (2.18)	.545	2.47	.011	.020	
F	-.002 (-.55)	$-.011^{U}$ $-.005^{E}$ (-2.18) (-.58)	-.538 (-2.20)	2.42 (5.03)	-.056 (-2.71)	.053 (2.28)	.521	2.52	.011	.020	

A = Regression on (logarithmic) first differences of equation (3) range 1976II-1984IV; B = Same as A but the overall unit value of exports is used, instead of the unit value of exports towards EMS countries, to obtain the export volume towards EMS countries, and to calculate the CX variable; C = Regression on levels with no lagged values of the independent variables; D = Same as A but the CX variable include competition only with domestic producers; E = Same as A but the exchange rate risk variable is a 4-quarters moving average; F = Same as A but distinction between the expected (E) and unexpected (U) component of exchange rate variability; D_1 and D_2 are Dummies for 1977I and 1981II; $\Sigma(0-2)$ is the sum of the coefficients of the first order polinomial with lags (0-2) except for regression C.

Table 3

Germany: Results of regressions equation (3)

Regression	Cst	V	CX $\Sigma(2\text{-}4)$	Y	D_1	D_2	R^2	DW	SSR	SE	$\hat{\rho}$
A	-2.08 (-2.22)	-.015 (-1.96)	-.522 (-2.38)	1.98 (16.92)	.056 (2.34)	-.067 (2.89)	.923	1.66	.015	.022	
B	-3.69 (-2.48)	-.020 (-1.63)	-.600 (-1.72)	2.40 (12.91)	.069 (1.82)	-.069 (1.95)	.876	1.82	.037	.035	
C	-3.04 (-2.40)	-.008 (-1.19)	-.367 (-1.79)	1.83 (11.34)	.045 (2.05)	-.034 (-1.57)	.918	1.74	.015	.023	.38 (2.49)
D	-2.74 (-3.36)	-.014 (-1.69)	-.391 (-1.93)	1.99 (15.61)	.060 (2.41)	-.066 (-2.75)	.919	1.57	.015	.023	
E	-2.33 (-2.65)	-.022 (-1.84)	-.418 (-2.09)	1.93 (15.41)	.060 (2.46)	-.067 (-2.87)	.933	1.74	.015	.022	
F	-.157 (-1.52)	$-.016^{U}$ $-.026^{E}$ (-2.06) (-2.09)	-.600 (-2.61)	1.94 (16.20)	.061 (2.50)	-.073 (-3.07)	.923	1.80	.014	.022	

A = Regression in (logarithmic) levels of equation (3) range 1976I-1984IV; B = Same as A but the overall unit value of exports is used, instead of the unit value of exports towards EMS countries, to obtain the export volume towards EMS countries, and to calculate the CX variable; C = Regression with no lagged values of the independent variable; D = Same as A but the CX variable includes competition only with domestic producers; E = Same as A but the exchange rate risk variable is a 4-quarters moving average; F = Same as A but distinction between the expected (E) and unexpected (U) component of exchange rate variability; D_1 and D_2 are Dummies for 1976III and 1980IV; $\Sigma(2\text{-}4)$ is the sum of the coefficients of the first order polynomial with lags (2-4), except for regression C.

Table 4

Results of regressions
equation (4)

1. export volumes

Country	Cst	V	PC	PI	Y	D_1	D_2	R^2	DW	SSR	SE
Germany	-4.74 (-7.15)	-.017 (-2.20)	.715 Σ(2-4) (2.92)	-.616 Σ(2-4) (-2.70)	1.93 (12.22)	.56 (2.41)	-.062 (-2.77)	.928	1.80	.013	.021
France	.009 (.65)	-.010 (-2.47)	.340 Σ(0-2) (1.60)	-.805 Σ(0-2) (-1.69)	2.49Σ(0-2) (5.46)	-.073 (-3.48)	.054 (2.44)	.546	2.74	.011	.020
Italy	.004 (.27)	-.014 (-2.26)	.766 Σ(0-2) (2.45)	-.815 Σ(0-2) (-1.67)	3.23 (4.11)	-.141 (-4.29)	.054 (1.76)	.547	2.11	.024	.029

For the definition of variables, and in particular of the dummy variables for each country, see footnotes to tables 1-3.

2. export prices

Country	Cst	V	PC	PI	R^2	DW	SSR	SE
Germany	.003 (.58)	-.005 (-2.01)	.305 Σ(1-5) (4.70)	.695 Σ (0-2) (10.68)	.996	1.45	.002	.008
France	.002 (1.09)	.005 (1.81)	.458 (4.48)	.542 Σ (0-3) (5.30)	.200	2.61	.006	.013
Italy	.033 (6.01)	.007 (1.89)	.630 Σ (0-5) (8.96)	.370 Σ(0-4) (5.27)	.998	1.93	.006	.013

For Germany and Italy the regression for export prices is on levels while for France on first differences. Σ(i-j) is the sum of the coefficient of the polynomial, with lag (i-j). For France the variable PC in the price equation is lagged one period. The sum of the coefficients on the PC and PI variables is constrained to be equal to unity. The range is 1976II - 1984IV.

APPENDIX

1. The CX and Y variables are calculated for each country i as follows:

(A.1) $$CX_i = PX_i/PC_i$$

$$PC_i = PX_j^{\alpha} + PI_h^{\beta} \quad j,h \neq i$$

where $\Sigma\alpha_j + \Sigma\beta_h = 1$
$j \in$ world; $h \in$ EMS

(A.2) $$Y_i = \sum_{h \neq i} \gamma_h DI_h$$

where DI_h is the aggregate demand of country h (in EMS area).

2. The EEC (Volimex) export volume and prices are available at a yearly frequency. The frequency has been extended to quarterly data by an interpolation procedure based on an OLS regression (on yearly data) using a reference series that is available also on a quarterly basis. The coefficients of the regression are then used to obtain the estimates of the base series at higher frequency, from the original data. The residuals are disaggregated using the Chowlin method, and added to the fitted values to get the new quarterly serie. The reference series are manufacturing exports volumes and prices, taken from OECD (Series A, Monthly Statistics of foreign Trade). The source of the other series used in the regression analysis, such as domestic demand and producer prices is also the OECD (Main Economic Indicators).

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