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CONTENTS

LIST OF TABLES AND FIGURES

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ABSTRACT

The partial effect of nominal exchange rate volatility on exports from each EMU member to the rest of the EMU is estimated on annual data for 1967-97, using modern time-series methods. The long-run relations between exchange rate volatility and exports are mostly negative and in several cases insignificantly different from zero. Thus, these estimates do not provide much support for the hypothesis that the elimination of nominal exchange rate volatility will significantly increase trade within the EMU. However, the EMU will presumably lead to geographical concentration of production and therefore indirectly to increased trade within the EMU and, during a transitional stage, to increased foreign direct investment, both within the EMU and between the EMU and the rest of the world.

1. INTRODUCTION

A major economic goal of the EMU is to increase trade and investment between member states and thereby to strengthen the Single Market. Expectations of reaching the goal are notably higher among policymakers and businessmen than among economists. The former see the costs of currency exchange and exchange rate uncertainty as formidable barriers to international trade and investment, while the latter have been unable in the past to find theoretical and empirical support for such a view.

The main goal of this study is to estimate the effects of nominal exchange rate uncertainty on trade between member countries before the start of the EMU. By doing so, we hope to get an idea of the qualitative and quantitative effects of eliminating nominal exchange rate uncertainty altogether.

Theoretically, the effects of exchange rate volatility on trade are ambiguous. In particular and perhaps contrary to intuition, it is not certain that exchange rate volatility must reduce trade. Empirically, the results are also ambiguous and depend, among other things, on what methodology is employed. Early research based on either time-series or cross-section data and simple OLS estimation produced mixed results. On balance, trade seems to be essentially unaffected, or, if the estimates are significant, only marginally affected by exchange rate volatility. But more recent research based on modern time-series methods delivers results that are less ambiguous; most of the studies estimate negative and substantial effects both in the short and the long run. We will employ such methods in this study. Compared to previous studies, our country sample is different and tailored to the question that we want to answer. Furthermore, we improve on the estimation methodology by using a full system approach rather than that of a single equation.

Foreign direct investment is another important economic variable commonly thought to be affected by exchange rate uncertainty. But, as with trade, the theoretical effects are again ambiguous. When exchange rate volatility is modelled simply as price volatility, it can be shown that ex post flexibility of investment and convexity of the profit function serve to increase profitability and consequently investment in the face of price uncertainty. Another modelling approach focuses on risk aversion. Risk averse investors will decrease their foreign direct investments if exchange rate volatility increases, since the certainty equivalent revenue falls. But if foreign direct investment and local production is a substitute for exports the opposite may occur. Empirical research mostly finds that increased exchange rate volatility has a positive effect on foreign direct investment.

Lacking data for bilateral foreign direct investment between the EMU countries, we will not attempt to undertake any empirical investigation of this issue. Instead, we confine ourselves to a discussion of how changes in market structure following the creation of the EMU may affect investment between EMU countries and between EMU and non-EMU countries.

Section 2 accounts for the ways in which the replacement of national currencies by the euro may affect cross-border transactions costs. Section 3 reviews some of the theoretical and empirical research on the effects of exchange rate volatility on international trade and investment. Section 4 describes the econometric methodology employed in our empirical analysis of trade and exchange rate volatility in the EMU countries. Results are presented in section 5. Section 6 takes a broader perspective and asks whether the EMU implies more for trade and investment than what can be expected from elimination of nominal exchange rate uncertainty alone. Section 7 discusses effects on trade and investment caused by changes in market structure. Finally, section 8 concludes.

2. REDUCTIONS OF TRANSACTION COSTS

In what concrete ways can the replacement of national currencies with a common currency promote more trade? Most obviously, the common currency will eliminate the need for currency exchange in connection with trade within the EMU. An illustrative example of the cost of currency exchange is given in the Delors report (1989): if a bill of some denomination is exchanged successively into all the other currencies in the EU and then back to its original denomination, half of the value is lost. Things are fortunately not as bad for business transactions, where spreads usually are much smaller than for exchanging bills and coins.

Another obvious cost for business transactions is the exchange rate uncertainty when a transaction involves two or more currencies. In international trade, there is usually a considerable time-lag between the date of purchase or sale and the date of payment. Exchange rates may thus appreciate or depreciate in the interim. Usually, the exchange rate risk is seen by business as something best avoided; much, if not most, of it is hedged against in the forward market for currency.

In its evaluation of the EMU, the European Commission (1990) estimated the combined cost to business of banking services in the form of currency exchange and hedging to be equal to 0.2-0.3 per cent of GDP per year on average for the EU countries. (The Swedish EMU Commission [Calmfors *et al*. 1997] arrived at an estimate of 0.2 per cent for Sweden.) The European Commission then added another 0.1 per cent to account for the internal costs to firms for handling foreign currency transactions when trading in the EMU**.** If these numbers are indicative of the social costs for currency exchange and hedging within the EMU, we can conclude that they are small in relative terms and that we should not be surprised if changes in trade from their elimination also turn out to be small.

Having a common currency will, of course, make price differences between countries more transparent than before. It is commonly argued that this should have the potential of increasing trade substantially since firms will become more sensitive to price differences. From an economic point of view, it is hard to know what to make of the argument. Almost all of the total value of cross-border trade is generated by firms. The conversion of prices from one currency to another should not entail more than a negligible cost. However, the argument has a clear intuitive appeal and it may be that 'mental inertia' amounts to much more than the cost of using a calculator.

Finally, the creation of TRANSIT deserves mention. TRANSIT is a computerized system for payments within the European System of Central Banks (ESCB). Since private banks in different countries are connected to their respective central banks, they are through this system also connected to one another across national borders. TRANSIT is therefore also a new, much swifter and less costly system for payments between private banks in different countries and consequently between private agents. TRANSIT is part of the EMU infrastructure, but something similar could, of course, have been created without the EMU—and probably would have, sooner or later.

3. THEORY AND PREVIOUS EMPIRICAL RESULTS

3.1 Exports and imports

Exchange rate uncertainty arises in connection with international trade because of the time-lag between contract and payment in a foreign currency. Firms are usually assumed to be risk averse and exchange rate uncertainty will therefore affect exports and imports.

It is not certain, however, that exchange rate uncertainty serves to reduce the levels of exports and imports. That depends on the degree of risk aversion on the part of exporters and importers in partial equilibrium, as shown by de Grauwe (1988). He considers an exporter who is faced with the choice of allocating a fixed amount of resources either to riskless domestic sales or risky exports. The utility of exports as a function of export revenue is concave due to risk aversion. A mean-preserving spread in exchange rate volatility is shown to reduce total utility of exports, but may actually increase marginal utility if the exporter is sufficiently risk averse. For example, if the coefficient of relative risk aversion is assumed to be constant, then expected marginal utility is a convex or concave function of the expected exchange rate depending on whether the coefficient of relative risk aversion is greater or smaller than unity. A mean-preserving spread in the exchange rate volatility will increase expected marginal utility if the function is convex and decrease it if the function is concave. A risk averse exporter will then shift resources on the margin from domestic sales to exports to compensate for the risk of a more negative outcome than before (an income effect), while a less risk averse exporter will do the opposite (a substitution effect).[1]

Much of the early empirical research is based on the theoretical model by Hooper and Kohlhagen (1978), which is specified in such a way that an increase in exchange rate uncertainty unambiguously leads to a reduction in exports and imports.

Empirical research on exchange rate uncertainty and trade before Koray and Lastrapes (1989) used either cross-section or time-series data, but took no account of the integration properties of the time-series data. The results from this research are mixed. Hooper and Kohlhagen (1978) found no significant effect of exchange rate uncertainty on trade, while Gotur (1985), who used data for five industrialized countries, found a mix of significant and insignificant negative effects. Cushman (1983, 1986, 1988) found mostly negative and significant effects, as did Kenen and Rodrik (1986) for eleven industrialized countries during the period of floating exchange rates. De Grauwe (1988) compared the period of fixed exchange rates in the 1960s with the period of floating rates in the 1970s and 1980s using a cross section of countries and arrived at significantly negative effects during the period of floating. Likewise, Bini-Smaghi (1991) established a significant negative effect for the EMS countries in the 1980s. Rose (2000) used a panel dataset of bilateral trade between 186 countries and territories, spanning 1970-90 and holding about 34,000 observations, to estimate, among other things, the effect of exchange rate volatility on trade. He found a highly significant negative effect. Also, the experiment of eliminating the mean exchange rate volatility of 5 per cent resulted in a 13 per cent increase in trade.

In the last ten years, it has become common to employ modern time-series methods to take account of the trend properties of the data. The results of these studies are also more clear cut; most suggest a significant negative relation between exchange rate uncertainty and trade. The first studies are those by Koray and Lastrapes (1989) and Lastrapes and Koray (1990). They arrived at a relatively strong and negative long-run (cointegration) relation between exchange rate uncertainty and bilateral imports for five industrialized countries, and a smaller and weaker, but still negative, short-run relation. Chowdhury (1993) studied trade between the G-7 countries and found the relations to be negative and significant both in the short and long run. In a simulation study, Gagnon (1993) found the maximum effect of exchange rate uncertainty on trade to be negative but quite small. In a series of studies, Arize (1995, 1996, 1997) and Arize and Shwiff (1998) arrived at significant and negative short- and long-run relations. Arize (1996) compared ERM and non-ERM countries and found little difference between the two groups. Arize (1997) is a replication of Chowdhury (1993). The only studies that we have found with results that do no suggest negative effects are those by Daly (1998) and McKenzie (1998). Daly studied bilateral trade between Japan and seven other industrialized countries, and found significantly positive relations in seven import and five export flows out of fourteen. McKenzie obtained mixed results for exports and imports by sectors of the Australian economy.

It is not clear what the theoretical and empirical research implies for the effects of the EMU on trade between the member countries. Nominal exchange rate volatility is completely eliminated within the EMU, but some amount of real exchange rate volatility will remain, due to differences in local rates of changes of prices and productivity. Moreover, it is not certain that total nominal exchange rate volatility will decrease, since much trade is still conducted in outside currencies. Also, one must allow for general equilibrium effects. It is possible that the elimination of some nominal volatility will give rise to greater real volatility, e.g. in greater fluctuations in variables such as real aggregate income and relative prices. But the presumption should be that the EMU will lead to convergence in macro variables, as argued by Frankel and Rose (1997), and to less macro uncertainty.

In this paper, we will adopt the approach of the more recent empirical research and focus on the partial relation between nominal exchange rate volatility and trade in the long and short run. We will mainly be concerned with nominal volatility since the EMU eliminates nominal but not real exchange rate volatility. Previous studies have looked at nominal or real exchange rate volatility or both. When the effects of both are investigated, the results are basically the same. This is presumably due to the fact that nominal and real volatility are essentially equal in the short run.

3.2 Foreign direct investment

Theoretical modelling of and empirical research about exchange rate uncertainty and foreign direct investment is scant. The existing theoretical literature follows one of two different approaches. The first approach focuses on production flexibility and is an extension of the research on price volatility and domestic investment. Effects of exchange rate volatility will in this approach generally depend on sunk costs in capacity, competitive structure and the convexity of the profit function in prices. With ex post flexibility in resource allocation and convexity, profits will increase with increased flexibility and convexity for a given level of exchange rate volatility. A recent example of this approach is Darby *et al.* (1999), who establish parametric conditions under which exchange rate volatility will reduce or increase foreign direct investment in an extension of the Dixit-Pindyck (1994) model of investment.

The second approach focuses on risk aversion. Exchange rate risk arises because of the time-lag between investment and profits in foreign currency. If it is assumed that exchange rate volatility reduces the certainty equivalent of prices in foreign currency, then an increase in volatility will reduce future profits and therefore foreign direct investment. However, if foreign direct investment is a substitute for domestic production and exports, then the opposite may result. The exposure to exchange rate risk is higher when foreign markets are supplied by exports than by local production, since in the former case both costs and revenues are in foreign currency. In the general case, when production at home and abroad can be sold on both the home and foreign markets, increased exchange rate volatility can reduce as well as increase foreign direct investment (Cushman 1986).

The empirical research mostly finds that increased exchange rate uncertainty has a positive effect on foreign direct investment. Positive effects are found by Cushman (1985, 1986) on pooled US bilateral outflow data for 1963-78 and on inflow data for 1963-86, and by Goldberg and Kolstad (1995) on bilateral investment flows between the US on the one hand and the UK, Canada and Japan on the other for 1978-91. Darby *et al.* (1999) likewise found positive effects for aggregate foreign direct investment for the UK, France, Germany, Italy, and the UK in recent decades, using a Dixit-Pindyck (1994) type of model. On the other hand, Bailey and Tavlas (1991) were unable to find any significant impact of exchange rate volatility on US foreign direct investment.

4. ECONOMETRIC METHODOLOGY

In this section we give a brief presentation of the econometric methodology used for our investigation of the relation between nominal exchange rate volatility and exports between EMU countries. To this end, let for each of the ten countries under consideration x_t be a four-dimensional vector of time-series comprising the logarithm of real exports (X_t), the logarithm of real foreign aggregate income (Y_t), the logarithm of the relative price of exports (P_t), and a measure of exchange rate volatility (V_t).[2] The volatility variable V_t is constructed as the logarithm of a moving sample 'standard deviation' of the change in the logarithm of the exchange rate:

$$V_t = \ln\left\{\left[(1/m)\sum_{i=1}^{m}(Q_{t+i-1} - Q_{t+i-2})^2\right]^{1/2}\right\}, \tag{1}$$

where Q_t is the logarithm of the exchange rate and $m = 4$. This measure of exchange rate uncertainty is similar to those used in much of the literature. For example, Chowdhury (1993) and Arize (1996) use e^{V_t} and thus consider a simple (non-linear) transformation of the measure in (1).[3] While formula (1) in our applications is generally found to deliver empirical results that are somewhat more stable than those obtained using other related and previously used measures, our empirical analysis also includes rather extensive robustness checks with respect to using different measures of exchange rate volatility. These involve using non-linear transformations of (1); using alternative definitions of the exchange rate data Q_t in (1); and using different lag-lead structures for determining the moving average in (1).

Our preferred measure of exchange rate uncertainty enables us, for each of the ten EMU countries under consideration, to estimate the percentage response of real exports to a one per cent change in the 'standard deviation' of the change of the effective nominal exchange rate. As will be shown below, our econometric methodology allows us to investigate such effects both in the short and in the long run. The issue of primary importance is to what extent these estimates (which are, of course, based on the historical information in our sample) can be used to make predictions about the future effects on trade within the EMU. While we believe that such estimates are useful initial guides in helping one to understand better the partial effects of exchange rate uncertainty, we also believe that there are severe limitations as concerns their direct applicability to the analysis of expected quantitative effects of the EMU. We will return to these limitations later, but there is one obvious 'technical' limitation to our preferred measure of exchange rate uncertainty (based on formula (1) and nominal exchange rates) that deserves comment already at this stage. This limitation arises because, for this specification, exchange rate uncertainty becomes undetermined in the 'EMU case' of $\Delta Q_t = 0$. This is, of course, one of the reasons why we choose to consider a rather extensive set of alternative measures of exchange rate risk.

Given the choice of variables to enter x_t, the following VAR process with k lags is fitted to each of the ten countries at hand:

$$\Delta x_t = \mu D_t + \Pi x_{t-1} + \sum\nolimits_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \varepsilon_t, \qquad (2)$$

where Δ is the difference operator; μ, Π, and Γ_i parameter matrices; D_t a vector of deterministic variables (which for example may include constants, trends, and different dummy variables); and ε_t a normally distributed vector of IID innovations. The analysis assumes that each variable in x_t is at most integrated of order one (denoted I(1)): that is, the variables in x_t are either I(1) or stochastically stationary (I(0)).[4] If some of the variables in x_t are I(0) or if there exist linear combinations among the I(1) variables in x_t that are I(0), then x_t is said to be cointegrated. In this case, the matrix Π is of reduced rank so that we can write

$$\Pi = \alpha\beta', \qquad (3)$$

where α and β are $4 \times r$ matrices (of full column rank) with $0 < r < 4$. The parameter r gives the number of I(0) stationary linear combinations that characterize the system (that is, the number of cointegrating vectors), and these are given by $\beta' x_t$.[5] It should be noted that if $\beta' x_t \neq 0$, then the system is in disequilibrium and the (relevant) elements of α determine how fast the system converges (error corrects) back to its long-run equilibrium (in which $\beta' x_t = 0$). The elements of α are therefore sometimes called speed of adjustment (or loading) coefficients and the model is termed vector error correction model (VECM).

Most of the empirical cointegration analysis in this paper is based on Johansen's (1988, 1991) so-called maximum likelihood (ML) procedure. Monte Carlo evidence reported by for example Gonzalo (1994) and Hargreaves (1994) suggests that the Johansen procedure is superior to many other alternative techniques available for analysing cointegration. Furthermore, in contrast to many other techniques, the Johansen procedure offers direct test statistics for the number of cointegrating vectors (that is, for the value of r).

Following the previous literature we expect, for each of the ten countries, to find $r = 1$ and a long-run export relationship that satisfies the following qualitative features:

$$X_t = \underset{(+)}{\delta_1} Y_t + \underset{(-)}{\delta_2} P_t + \delta_3 V_t . \tag{4}$$

The export relationship in (4) can be derived from a bilateral trade model, as presented in Goldstein and Khan (1985). This model solves simultaneously for exports and imports as functions of nominal incomes in the two regions, the price of all domestically produced goods and the exchange rate. It is assumed that goods are imperfect substitutes, whether produced domestically or abroad, and that competition is perfect. To be able to solve for exports without solving the full system, we have to make the strong assumption that export supply is infinitely elastic so that the exporting country can supply any amount at constant cost (price). In other words, domestic conditions play a direct role only through the relative price of exports. Empirically, we could, of course, relax this assumption and include other domestic aggregates, such as real domestic income, in our analysis, but a lack of degrees of freedom due to short time-series prevents us from pursuing this route further

Since we wish to analyse the importance of exchange rate uncertainty for trade flows, it is the parameter δ_3 that is of primary interest. If the hypothesis $\delta_3 = 0$ cannot be rejected, then exchange rate uncertainty is suggested to be of little importance for trade in the long run.

Even if it is not possible to reject the hypothesis that exchange rate volatility does not matter for trade in the long run, exchange rate risk may still be of great importance for trade by having short-run effects. Model (2) potentially allows for such short-run effects through the parameters in the Γ_i matrices. Tests for the presence of such effects can thus easily be performed by formulating hypotheses of exclusion restrictions on the lags of Δx_t.

Methodologically we improve on several previous analyses in this area by not *a priori* restricting attention to a conditional model for real exports (as do for example Chowdhury 1993 and Arize 1996). Rather, we acknowledge that the analysis of a conditional single equation model for exports requires certain parametric restrictions on the full system in (2). These restrictions are testable but depend on the purpose of the analysis (Engle, Hendry and Richard 1983). For statistical inference to be fully efficient within the conditional model for exports, it is generally sufficient to assume that the conditioning variables are *weakly exogenous* (with respect to the parameters of interest), which here implies that they are not directly affected by the cointegrating relationships that characterize the system. Thus, these restrictions may be tested as zero restrictions on certain of the elements in the loading matrix α. If the purpose of the analysis is instead to derive multi-step forecasts of real exports from the conditional model, then weak exogeneity is not sufficient. In this case the relevant concept is that of *strong exogeneity*, which in addition to weak exogeneity requires the absence of Granger causality from endogenous variables (here real exports) onto the conditioning variables. These additional restrictions are thus equivalent to zero restrictions on some of the elements of the Γ_i matrices in (2).

A third concept is that of *super exogeneity*. The conditioning variables are called super exogenous if and only if they are weakly exogenous and, in addition, the parameters of the conditional models are invariant to interventions affecting the parameters in the models of the conditioning variables (the so-called marginal models). Under these circumstances one may validly undertake policy analysis on the parameters of the conditional models when 'structural' changes occur in the models of the conditioning variables.

While the Johansen method, as mentioned previously, has proven itself to be superior to many other alternative methods for analysing cointegration, it requires the estimation of a rather large number of parameters, in our case $4^2 k$ parameters excluding the deterministic components in μ. This may constitute a problem in particular when the available number of observations is small. Because this is unfortunately the case in our applications, we have chosen also to consider an alternative method for analysing cointegration, namely the so-called canonical cointegration regression (CCR) approach of Park (1992). This method does not require estimation of the full dynamic system in (2), but uses instead directly relationship (4) with transformed stationary data. In general, the approach represents the same cointegrating relationships as the original models but constructs them in such a way that the usual least squares procedure yields asymptotically efficient estimators and χ^2 inference for hypothesis tests (for further details see Park 1992).

5. EMPIRICAL RESULTS[6]

This section reports the results of empirically analysing the importance of exchange rate volatility for short- and long-run trade among the EMU countries Austria (AUS), Belgium (BEL), Finland (FIN), France (FRA), Germany (GER), Ireland (IRE), Italy (ITA), Netherlands (NET), Portugal (POR), and Spain (SPA) along the lines suggested in the previous section. The data for each country consist of annual observations on the aforementioned four variables (real exports, real foreign aggregate income, the relative price of exports, and a measure of exchange rate volatility) covering the time period 1967 to 1997.[7] While we note that this choice of sample period implies the use of data from both flexible and fixed (or semi fixed) exchange rate periods, restricting the analysis to a period with exclusively fully flexible rates would require considering an unacceptably short sample period.[8] A battery of diagnostic tests is used to check whether the statistical properties of the estimated models are acceptable. The results and discussions below are based on the exchange rate volatility measure being calculated according to formula (1) with m set equal to 4 and Q_t constructed using nominal exchange rates. The final part of this section provides a robustness analysis with respect to, among other things, using alternative measures of exchange rate uncertainty.

The first step in the empirical analysis consists of determining an appropriate lag length (k) for the VAR model of each country. The approach adopted in this paper is a sequential top-down reduction procedure, undertaken through a series of F tests.[9] The results, given a maximum lag order of three lags, are shown in Table 1. As can be seen, $k = 1$ is rejected in only three cases at the 5 per cent test level (FIN, ITA, and POR) and is accepted in all cases at the 1 per cent test level. From the analysis in Table 2, it however appears that this choice of lag length does not provide empirical models with fully satisfactory error term properties. In particular, with $k = 1$ the assumption of normality is rejected in several cases. Furthermore, the assumption of uncorrelated residuals over time is rejected in two cases at the 1 per cent test level (IRE and POR). Unfortunately, these problems could neither be mitigated by increasing the lag length further nor by including other deterministic components such as linear or quadratic trends. In what follows we therefore retain the specification with $k = 1$ for each country, but it is emphasized that these statistical problems imply that the subsequent results have to be interpreted cautiously.

An important economic implication of the single lag specification is that Γ=0 for each i. That is, there are no lagged first differenced terms on the right hand side of the equations of the VAR models and export volumes are thus unaffected by exchange rate uncertainty in the short run (in this particular sense). This result is a bit surprising given the previous literature. Using somewhat different data, Chowdhury (1993) finds strongly significant negative short-run effects on French, German, and Italian real exports from increased uncertainty in real effective exchange rates over the period 1976-90. The effects documented by Arize (1996) over the period 1973-92 are less clear cut but even in this study some significant short-run effects appear to be present in countries such as Belgium and Spain (and maybe also the Netherlands). While it obviously is impossible exactly to know why our results differ from those of previous studies, one potential explanation is the difference in the frequency of the data. Both Chowdhury and Arize derive their results using quarterly data, but our study is, as mentioned previously, based on data observed at the annual frequency. If the unpredictable exchange rate variations that are of importance occur at higher frequencies than covered by annual lags, then using annual data may not give an adequate picture of the importance of exchange rate volatility in the short run.[10] However, if this assertion were correct, then one would also expect our results to be less in conflict with those of previous studies when it comes to the analysis of fluctuations at lower frequencies.[11] This brings us to the analysis of the cointegration properties of our data.

The first step in the analysis of cointegration involves examining the rank of the estimate of the matrix Π in (2); that is, testing for the number of cointegrating vectors that are present. The two rank tests (the so-called trace and maximum eigenvalue tests) of the Johansen ML procedure deliver the results given in Table 3. The relevant null and alternative hypotheses are as displayed in rows 2 and 3 of the table. The recommended testing strategy involves starting with the trace test and test the null of $r = 0$. If this hypothesis is not rejected, then we conclude that the system is not cointegrated, provided the conclusion is consistent with the outcome for the maximum eigenvalue test. If on the other hand $r = 0$ is rejected, then we move on to the next null hypothesis $r \leq 1$. We then apply the same decision rules for this null as for the null $r = 0$ and continue the process until some null is accepted or we arrive at the final alternative hypothesis. Considering the details of Table 3, it is seen that the evidence in most countries, at conventional levels of significance, indeed favours the single cointegrating vector model. However, there are some borderline cases (IRE, ITA, NET, and POR) and also two cases for which the hypothesis $r = 1$ is firmly rejected (AUS and SPA). Because the precise choice of r appears not to significantly matter for the particular empirical questions that we are addressing here, we choose in the remaining analysis to stick to the $r = 1$ specification for each country.[12]

The next question is then whether the estimated cointegrating vectors tally with the properties suggested by equation (4)? In order to be able to answer this question, we need to look at the particular element estimates in the (unique) cointegration matrices subject to the (non testable) normalization that the element on X_t in each country is equal to unity. The results are presented in Table 4. In each case, foreign aggregate income is positively related to exports. The long-run export elasticity with respect to foreign income ranges from a low of 0.86 in Belgium to a high of 19.55 in Germany. Compared to previous studies these income elasticities generally stand out as relatively high and, in some cases, even as ridiculously high (GER and POR). The relative price of exports has the expected negative sign in all countries except France and Portugal. While the elasticities for this variable generally are more in line with previous findings (that is, closer in absolute value to previous estimates), some point estimates again are clearly unreasonable (in particular again those for GER and POR). As concerns the measure of exchange rate uncertainty, parameter estimates are negative in seven out of the ten cases. The last column of Table 4 reports the results from undertaking the LR tests of excluding V_t from the cointegrating relationships. Disregarding the unreasonable vector estimates in case of Germany and Portugal, the results point to very weak significance, rejecting the null hypothesis at the 5 per cent test level only in two cases (BEL and SPA).[13] Indeed, in no other case is the test even rejecting at the 15 per cent level of significance (the lowest *p* value obtains for the test on Finnish data and is 0.15).

An issue of considerable interest is how important *economically* these estimated exchange rate uncertainty effects are. To shed some light on this issue, Figure 1 uses the estimated long-run elasticities provided in Table 4 to compute, for each country and year of the sample period (1969-94), $\hat{\delta}_3 100\Delta V_t$ and plot it against the actual growth rate of exports ($100\Delta X_t$). When interpreting these graphs it is important to bear in mind that several of the underlying point estimates are not significantly different from zero at conventional test levels (see Table 4). We believe nevertheless that it is informative to know that even relatively small estimated elasticities (in absolute value) may have rather substantial economic implications in terms of changes in $\hat{\delta}_3 100\Delta V_t$. For example, between the years 1978 and 1979 Dutch real exports grew by approximately 9 per cent and at the same time the quantity $\hat{\delta}_3 100\Delta V_t$ changed by almost –5 percentage points, despite an estimate of δ_3 no larger than 0.04 in absolute value. Similarly, in the case of Austrian data, an elasticity estimate of only 0.1 per cent in absolute value implied in 1992 that $\hat{\delta}_3 100\Delta V_t$ changed by roughly –12 percentage points in a situation where exports themselves only grew by approximately 4 per cent.

As concerns the speed of adjustment (or loading) coefficients (that is, the elements of the matrix α, cf. equation (3)), which measure the importance of a disequilibrium situation for the short term evolution of the variables, our estimates are, broadly speaking, not too different from those obtained in previous studies. For example, in the export equations, the significant parameters are in the range –0.2 to 0.9, which is rather close to the ranges reported by Chowdhury (–0.3 to –0.6) and Arize (–0.2 to –1.0). These estimated adjustment effects imply that the previous year's equilibrium value relative to the actual value of exports in that year (on average) has a weight of between 20 (BEL) and 90 (FRA) per cent in explaining the current year's export level. However, as before, the empirical significance levels of our estimates are generally higher than in previous studies (and thus the effects more uncertain).

To summarize so far, our results suggest that exchange rate volatility has had a rather limited importance for the progression of trade within the EMU since the late 1960s, both in the short and the long run. These results are not in line with previous findings that suggest significant negative effects on exports from exchange rate uncertainty in the short as well as the long run. However, due to several statistical problems our results need to be interpreted with care.

As a means of increasing the reliability of our results we undertake robustness checks with respect to several aspects.

First, Table 5 investigates the effects of using an alternative method for analysing the cointegration properties of the data. Here we employ the so-called canonical cointegration regression (CCR) approach suggested by Park (1992). Compared to the trivial static least squares approach (also reported in the table within parentheses) this approach has the distinct advantages of both correcting for the missing 'short-run dynamics' of the data (through non parametric corrections) and delivering estimates from which it is possible to construct test statistics that permit standard inference procedures. From the details of the table (which is structured in the same way as Table 4), it can be seen that many of the previously gained results are unchanged in qualitative terms: the income elasticities are always positively signed and quite large; the elasticities for the competitiveness variable are mostly negatively signed; and, the elasticities on the volatility variable are mostly negatively signed but often quite imprecisely estimated. As it happens, the peculiarities for the German and Portuguese data vanish when using this method of estimation.

Second, Table 6 checks the effects of conditioning the analysis on alternative measures of exchange rate uncertainty, using a differing weighting scheme for constructing 'effective' variables, and controlling for the Single European Act in 1986.[14] As can readily be seen, the results are again largely unchanged as concerns the qualitative key features. The exchange rate volatility measure mostly enters the analysis negatively signed and is documented not to be statistically significant (at conventional levels of significance) in several cases.[15]

Finally, Table 7 repeats (essentially) the analysis of Table 6 using quarterly instead of annual data. Due to lack of data this analysis only comprises seven of the ten previously considered countries (AUS, BEL, and IRE are excluded) and is undertaken on a shorter sample period (see the last row of Table 7). We emphasize that while the information in this table again gives a qualitative picture that is not too different from that previously obtained (that is, using the results in Tables 4 to 6), the *p* values for testing the null of $\delta_3 = 0$ now in general appear to be somewhat lower. However, given that the estimates are still insignificant in approximately half of the cases and that the statistical properties of these quarterly models (not shown to save space) are even worse than those of the corresponding annual models (cf. Table 2), it is difficult to know what to make out of this apparent slight increase of significance.[16]

6. THE EFFECT OF NATIONAL BORDERS

Several studies have been undertaken in recent years on the effect of national borders on international trade. The first study is by McCallum (1995), who found that trade between Canadian provinces is 22 times larger than trade between contiguous provinces and states across the Canada-US border, after controlling for distance and GDP. In a related study, Engel and Rogers (1996) found that price differences for various goods between city pairs in Canada and the US varied much more between cities situated in different countries than between city pairs situated in the same country, after controlling for distance. The border was found to be able to explain about 30 per cent of the standard deviation of the differences in prices (in logs), while distance was found to explain only about 20 per cent. Measured differently, the border between Canada and the US was estimated to add 1,750 miles between city pairs! A substantial part of the border effect, but still less than half, was found to be attributed to sticky nominal prices, i.e. to the fact that export and import prices adjust very little to changes in the nominal Canada-US exchange rate.

Both of these studies demonstrate that the national border has a strong effect on trade and price setting. The question is how relevant the findings are for the establishment of the EMU. Canada and the US are more similar in terms of language, culture, and institutions than the average country pair in the EMU. This would presumably tend to make the border effect stronger in Europe. Furthermore, we do not know how much of the border effect that is due to the existence of separate currencies.

In the aforementioned study by Rose (2000), he found that countries with a common currency traded 3.5 times more with each other than countries with separate currencies, controlling for a large number of factors, such as distance, GDP, language, colonial past, contiguous borders and, in particular, exchange rate volatility. Hence, it appears that a currency union is much more than the elimination of exchange rate uncertainty.

Unfortunately, it will take many years before we have sufficient data to undertake ex post tests of whether and to what extent the creation of the EMU has contributed to intra-EMU trade.

7. MARKET STRUCTURE, TRADE AND INVESTMENT

Until now, we have focussed on the direct effects of eliminating exchange rate uncertainty on trade within the EMU. As we have seen, these effects seem to be rather insignificant. But the EMU may have additional indirect effects on trade and these may be more important than the direct effects, although perhaps even more difficult to estimate. What we have in mind in particular are effects that eventually affect trade (and foreign direct investment) via structural changes in markets for goods and services.

The secondary market for government debt issue—treasury bills and bonds—serves well as a prototypical example of indirect effects through changes in market structure. Prior to the Single Market, regulations of various kinds, such as controls of foreign direct investment and capital flows, protected national secondary markets for treasury bills and government bonds. With the advent of the Single Market, these regulations have been abolished. But, as long as exchange rates were not completely fixed before the establishment of the EMU, national agents had a competitive advantage based on their knowledge of the domestic economy and in particular of policy making by the central bank and the government, since exchange rate risk (inflation risk) then still was the major component of risk attached to government debt issue.

After the introduction of the euro, or the irrevocable fixing of exchange rates, this changed radically. For investors inside the EMU, there is no longer any exchange rate risk. There are still credit, liquidity, settlement, legal, and event risks, but these are presumably not as important and can be handled equally well by foreign investors. Also, all investors in EMU countries have about equal opportunities to forecast future interest rates, since short term interest rates are set by the ECB and depend on economic conditions in the EMU as a whole, as do long-term interest rates.

All of this—increased market access, elimination of exchange rate risk, lower transactions costs, and refocus on credit risk—will lead to portfolio shifts out of national markets and into the integrated EMU market and to increased competition among financial intermediaries. This will in turn probably cause a reduction in the total number of financial intermediaries in the EMU market as a whole, increased trading volumes for those that remain in the market, and lower price-cost margins.[17]

In the secondary market for government debt, as well as in other markets, the effects will not only be limited to lower prices, larger volumes, and fewer and larger firms. Geographical concentration is most likely also going to increase. Financial centres benefit from geographical concentration because of many different kinds of positive externalities. For example, concentration promotes the development of a large pool of specialized labour and lower transaction costs in dealings with other financial intermediaries and customers. The concentration of the financial industry in the US is an indication that the present trend towards concentration in Europe is going to continue and bring about substantial changes in the location of the industry.

The consequences for trade in financial services and foreign direct investment are quite clear. Geographical concentration should lead to increased trade in financial services inside the EMU. National financial intermediaries may move part of their operations to the financial centre(s) in order to benefit from agglomeration economies. In other words, foreign direct investment will increase during the period of restructuring of the European financial industry. Similar effects will presumably be seen when it comes to trade between the EMU and non-European countries, primarily the US. The creation of a large integrated financial market in the EMU will continue to attract US financial intermediaries to invest in Europe, since volumes can be increased for more or less the same fixed costs. Likewise, the increase in size and competitiveness of European firms will induce these firms to start or expand operations in the US and elsewhere. This will in turn lead to an expansion of international intra-firm transactions, i.e. to increased international trade in financial services.

The financial industry and in particular the secondary market for government debt is a special example, and the introduction of the euro will probably have a stronger effect in this market than in most other markets.

But the same mechanisms can be expected to be at work in other service industries and in manufacturing as well, presumably leading to a period of restructuring across national borders and a permanent increase in trade within the EMU and between the EMU and outside countries.

8. SUMMARY

We have made an empirical investigation of the relation between exports and nominal exchange rate uncertainty for ten EMU countries during the period 1967-97. In contrast to early empirical research on trade and exchange rate uncertainty, we have employed modern time-series methods, and in contrast to more recent research, we have used a full system approach rather than that of estimating a single equation.

When estimating the relation between exports and a measure of exchange rate uncertainty on annual data, we found that the Johansen method gives seven negative and three positive parameter estimates in the long-run cointegrating vectors. Two of the negative and two of the positive estimates are significant. The CCR method, on the other hand, yields eight negative and two positive parameter estimates. Five of the negative and one of the positive estimates are significant. The results are quite robust to changes in the exchange rate uncertainty variable and to controlling for the Single European Act. When re-estimating on quarterly data for fewer countries and a shorter time period, we found similar results but very poor statistical properties of the model. Although the estimated elasticities are small in most cases, we demonstrate that recorded exchange rate volatility can give rise to partial effects on exports that are of the same magnitude as the recorded changes in exports themselves.

Our analysis suggests that the relation between exports and exchange rate volatility has been negative in the EMU countries in general, but also that we cannot expect intra-EMU trade to increase significantly when nominal exchange rate volatility is eliminated. At the same time, at least four qualifications to this conclusion must be made. First, our analysis only concerns partial effects, holding the effects of income and relative prices constant. The introduction of a common currency will give rise to general equilibrium effects, including effects through changes in income and relative prices. Second, the introduction of the EMU can give rise to changes in parameter values. Third, recent research on the 'border effect' and on the effects of a common currency in addition to those of exchange rate volatility indicates that the EMU may have substantial effects on trade for reasons that remain to be uncovered. Fourth, the statistical properties of our models are somewhat poor.

There are yet other ways in which trade—and investment—may be increased as a result of EMU. Price differences between countries will become more transparent and financial costs in relation to trade may fall. This should increase competition in product and service markets and lead to restructuring, with fewer firms and geographical concentration. We have taken the financial market as a prototypical example. Greater concentration means more trade within the EMU and, at least during a transition stage, more cross-border investment within the EMU and between EMU and the rest of the world.

NOTES

[1] Ambiguous effects on exports of increased exchange rate uncertainty are found in other models as well. One example is Sercu and Vanhulle (1992), where exporting is seen as an option, and other options are mothballing, exit and foreign direct investment. Another example is Cushman (1986), who allows for exchange rate uncertainty with a third country.

[2] Due to lack of data, Luxembourg is excluded from the empirical analysis.

[3] For further discussions of the properties of the volatility variable, see Chowdhury (1993) and Arize (1996) and the references in those papers.

[4] The possibility of I(2)-ness is thus excluded. This does however not appear to be an overly restrictive assumption, given the chosen endogenous variables.

[5] Above, it was noted that some of the endogenous variables may themselves be I(0). This is the case of so-called trivial cointegration in which some columns of the cointegrating matrix may contain only one single element which is not equal to zero.

[6] The estimations in this paper are undertaken using PcFIML version 9.0 and GAUSS version 3.2.38.

[7] The sources for the data are as follows: the data on exports are from UN, *International Trade Statistics Yearbook*, various issues. Exchange rates, indices of export and consumer prices, and GDP are from IMF, *International Financial Statistics*, September 1999. Details of the transformations are available from the authors upon request.

[8] Because observations both at the beginning and at the end of the sample period are lost due to the construction of the volatility measure, the effective sample period is quite short even when including all available observations.

[9] This procedure has been recommended by for example Ng and Perron (1993) for determining the lag length in univariate tests of integration.

[10] A robustness analysis with respect to using quarterly rather than annual data is given in Table 7. Although this analysis is limited, it does not give much support to the hypothesis that the differences between the results can be explained by the frequency of the data.

[11] If the frequency problem more generally implies that our constrained volatility variable as such is inadequate, then of course our results may also be misleading as concerns the importance of exchange rate volatility in the long run.

[12] For the analysis to be meaningful in the case when $r > 1$, one however needs to impose (non testable) identifying assumptions on the cointegration vectors.

[13] This conclusion also holds true for (conditional) *t* tests of the null hypothesis δ_3.

[14] The signing of the Single European Act was a strong signal that the Single Market would be successively implemented and completed by 1993. It led to a substantial increase in intra-EU foreign direct investment (cross-border mergers) in the late 1980s and possibly to an increase in trade. Ideally, one would like to control for the reduction and elimination of various administrative trade barriers individually, but that is not possible given the lack of data on such trade barriers.

[15] All models used in Table 6 are based on $k = 1$. In many cases the residual diagnostics given in Table 2 are representative also for the statistical properties of these alternative models. In particular, the introduction of the Single Act dummy variable does not improve the models from a statistical point of view.

[16] The models in Table 7 are again based on $k = 1$. In some cases, the poor residual diagnostics could be mitigated by increasing the lag length, but the qualitative results of Table 7 were unaffected by this change. Details of the residual diagnostics for these models are available from the authors upon request.

[17] For a discussion of the wider and more detailed effects of the EMU on the structure of financial markets, see Prati and Schinasi (1997).

TABLE 1
F TESTS FOR LAG ORDER

Country	3 Lags vs. 2 lags	2 Lags vs. 1 lag
AUS	1.41 [0.19]	0.82 [0.65]
BEL	0.85 [0.63]	0.96 [0.52]
FIN	1.55 [0.13]	2.20 [0.04]
FRA	0.80 [0.68]	0.97 [0.51]
GER	1.36 [0.21]	0.49 [0.93]
IRE	2.42 [0.01]	1.72 [0.11]
ITA	1.50 [0.15]	2.31 [0.03]
NET	1.05 [0.43]	0.66 [0.80]
POR	1.75 [0.08]	2.56 [0.02]
SPA	1.53 [0.14]	0.73 [0.74]

Notes: Each VAR model is augmented by a vector of constants. The numbers within square brackets in each entry are p values. The maximum lag length considered is 3. The common sample period is 1971-94. The F tests are distributed as $F(16, 37)$ and $F(16, 25)$ in case of 3 vs. 2 lags and 2 vs. 1 lag respectively.

TABLE 2
DIAGNOSTIC TESTS ON UNRESTRICTED VAR SYSTEMS

Country	Lag length	Vector AR	Vector normality	Vector hetero. I	Vector hetero. II
AUS	1	1.35 [0.19]	19.38 [0.01]	0.96 [0.57]	166.6 [0.06]
BEL	1	1.00 [0.49]	24.90 [0.00]	0.72 [0.87]	176.4 [0.02]
FIN	1	1.38 [0.17]	19.59 [0.01]	0.43 [1.00]	132.4 [0.66]
FRA	1	0.98 [0.52]	26.30 [0.00]	0.73 [0.86]	149.6 [0.27]
GER	1	0.90 [0.62]	10.91 [0.21]	0.66 [0.92]	140.0 [0.49]
IRE	1	3.11 [0.00]	18.49 [0.02]	1.08 [0.43]	175.3 [0.02]
ITA	1	1.38 [0.17]	17.76 [0.02]	0.61 [0.95]	153.4 [0.21]
NET	1	1.07 [0.42]	24.45 [0.00]	0.56 [0.97]	160.4 [0.11]
POR	1	3.40 [0.00]	4.44 [0.82]	0.58 [0.97]	149.5 [0.28]
SPA	1	1.60 [0.08]	10.02 [0.26]	0.63 [0.94]	146.5 [0.34]

Notes: The numbers within square brackets in each entry are *p* values. The sample period is 1969-94. The vector AR test is a multivariate LM test against autocorrelation of order 2. This test uses an $F(32, 38)$ distribution under the null of no autocorrelation. The vector normality test is a multivariate normality test suggested by Doornik and Hansen (1994). This test has an asymptotic $\chi^2(8)$ distribution under the null of multivariate normality. The vector heteroscedasticity tests (I: without cross products; II: with cross products) are multivariate versions of the White test against heteroscedasticity. These tests use $F(80, 27)$ and $\chi^2(140)$ distributions respectively under the null of no heteroscedasticity. For further details of the tests, see Doornik and Hendry (1997).

TABLE 3
TESTS FOR THE NUMBER OF COINTEGRATING RELATIONSHIPS

	Maximum eigenvalue test				Trace test			
H_0 :	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
H_1 :	$r \geq 1$	$r \geq 2$	$r \geq 3$	$r = 4$	$r = 1$	$r = 2$	$r = 3$	$r = 4$
AUS	16.70 [14.13]	13.02 [11.02]	3.95 [3.34]	1.12 [0.95]	34.79 [29.44]	18.09 [15.31]	5.07 [4.29]	1.12 [0.95]
BEL	32.19** [27.23*]	17.48 [14.79]	10.69 [9.05]	0.10 [0.08]	60.45** [51.15*]	28.26 [23.91]	10.79 [9.13]	0.10 [0.08]
FIN	29.90* [25.30]	16.90 [14.30]	13.52 [11.44]	0.70 [0.60]	61.03** [51.64*]	31.12* [26.34]	14.22 [12.03]	0.70 [0.60]
FRA	38.56** [32.63**]	10.89 [9.22]	5.06 [4.28]	4.01* [3.40]	58.52** [49.52*]	19.96 [16.89]	9.07 [7.68]	4.01* [3.40]
GER	38.23** [32.35**]	20.59 [17.42]	3.61 [3.06]	0.83 [0.71]	63.26** [53.53*]	25.04 [21.18]	4.45 [3.76]	0.83 [0.71]
IRE	23.99 [20.30]	16.88 [14.28]	5.67 [4.79]	1.23 [1.04]	47.76* [40.41]	23.77 [20.11]	6.89 [5.83]	1.23 [1.04]
ITA	31.75* [26.87]	13.07 [11.06]	3.82 [3.23]	0.33 [0.28]	48.97* [41.44]	17.22 [14.57]	4.15 [3.51]	0.33 [0.28]
NET	24.92 [21.09]	20.89 [17.67]	8.32 [7.04]	0.01 [0.01]	54.14** [45.81]	29.22 [24.72]	8.33 [7.05]	0.01 [0.01]
POR	35.85** [30.34*]	24.89* [21.06*]	5.22 [4.42]	0.32 [0.27]	66.28** [56.08**]	30.43* [25.74]	5.54 [4.69]	0.32 [0.27]
SPA	50.26** [42.53**]	37.11** [31.40**]	12.86 [10.88]	0.07 [0.06]	100.3** [84.86**]	50.03** [42.34**]	12.93 [10.94]	0.07 [0.06]

Notes: The deterministic variables are unrestricted (see the notes of Table 1). For details of the test statistics see for example Johansen (1988, 1991). The parameter r denotes the number of cointegrating vectors (the rank of Π). * indicates significance at the 5% test level and ** significance at the 1 percent test level. The numbers within square brackets in each entry are small sample adjusted test values (Reimers 1992). The critical values are based on a response surface fitted to the results of Osterwald-Lenum (1992).

TABLE 4
ESTIMATED NORMALIZED COINTEGRATING VECTORS
AND LIKELIHOOD RATIO TESTS: JOHANSEN METHOD

Country	Cointegrating vectors	LR tests of $H_0 : \delta_3 = 0$
AUS	$X_t = 3.28Y_t - 0.31P_t - 0.10V_t$	0.61
BEL	$X_t = 0.86Y_t - 3.49P_t - 0.45V_t$	14.71**
FIN	$X_t = 2.80Y_t - 1.01P_t - 0.18V_t$	2.08
FRA	$X_t = 1.87Y_t + 1.29P_t - 0.003V_t$	0.09
GER	$X_t = 19.55Y_t - 21.37P_t + 3.40V_t$	5.28*
IRE	$X_t = 1.91Y_t - 6.29P_t - 0.42V_t$	1.00
ITA	$X_t = 2.07Y_t - 0.53P_t + 0.01V_t$	0.42
NET	$X_t = 1.27Y_t - 0.30P_t - 0.04V_t$	0.07
POR	$X_t = 7.93Y_t + 15.28P_t + 1.16V_t$	4.95*
SPA	$X_t = 3.45Y_t - 1.70P_t - 0.27V_t$	5.02*

Notes: The LR tests test $H_0 : \delta_3 = 0$ in the long-run export equation $X_t = \delta_1 Y_t + \delta_2 P_t + \delta_3 V_t$.

* indicates significance at the 5 percent test level and ** significance at the 1% test level. The critical values are from the $\chi^2(1)$ distribution.

TABLE 5
ESTIMATED NORMALIZED COINTEGRATING VECTORS AND WALD TESTS:
CCR METHOD

Country	Cointegrating vectors	Wald tests of $H_0 : \delta_3 = 0$
AUS	$X_t = -1.76 + 2.98Y_t - 0.86P_t - 0.16V_t$ [OLS: $X_t = -1.20 + 2.84Y_t - 0.92P_t - 0.14V_t$]	31.69**
BEL	$X_t = 14.31 + 1.93Y_t + 0.87P_t - 0.06V_t$ [OLS: $X_t = 11.35 + 1.66Y_t - 0.05P_t - 0.06V_t$]	6.12*
FIN	$X_t = -3.49 + 2.50Y_t - 1.26P_t - 0.002V_t$ [OLS: $X_t = -2.88 + 2.46Y_t - 1.09P_t - 0.13V_t$]	0.00
FRA	$X_t = 13.46 + 1.99Y_t + 1.16P_t + 0.01V_t$ [OLS: $X_t = 12.39 + 1.88Y_t + 0.80P_t + 0.02V_t$]	2.19
GER	$X_t = 8.81 + 1.42Y_t - 0.41P_t - 0.17V_t$ [OLS: $X_t = 8.68 + 1.53Y_t - 0.35P_t - 0.11V_t$]	46.64**
IRE	$X_t = -7.82 + 4.78Y_t + 1.01P_t - 0.13V_t$ [OLS: $X_t = -12.51 + 5.05Y_t + 0.28P_t - 0.13V_t$]	9.75**
ITA	$X_t = 6.68 + 2.23Y_t - 0.99P_t - 0.03V_t$ [OLS: $X_t = 7.57 + 2.09Y_t - 0.95P_t - 0.02V_t$]	12.77**
NET	$X_t = 10.91 + 1.67Y_t + 0.28P_t + 0.12V_t$ [OLS: $X_t = 9.74 + 1.61Y_t + 0.01P_t + 0.05V_t$]	11.14**
POR	$X_t = -2.26 + 4.92Y_t + 1.38P_t - 0.05V_t$ [OLS: $X_t = -8.72 + 4.54Y_t - 0.39P_t - 0.19V_t$]	1.04
SPA	$X_t = -5.45 + 3.78Y_t - 1.06P_t - 0.06V_t$ [OLS: $X_t = -5.52 + 3.70Y_t - 1.13P_t - 0.10V_t$]	3.17

Notes: The sample period is 1969-94. Unmodified OLS estimates are given within square brackets in the entries of column 2. The Wald tests test $H_0 : \delta_3 = 0$ in the long-run export equation $X_t = \delta_1 Y_t + \delta_2 P_t + \delta_3 V_t$ using the CCR method (Park 1992).

* indicates significance at the 5% test level and ** significance at the 1 percent test level. The critical values are from the $\chi^2(1)$ distribution.

TABLE 6
ANALYSIS OF ROBUSTNESS: JOHANSEN METHOD

$\hat{\delta}_3$ [*P* Values for LR tests of $H_0 : \delta_3 = 0$]

Country	Single act dummy	e^V	Fixed weights	Real exchange rates	8-term MA	V_{LAG}
AUS	-0.20 [0.37]	-2.25 [0.42]	-0.85 [0.22]	-0.14 [0.03]*	-0.10 [0.10]	-0.09 [0.17]
BEL	-0.28 [0.01]*	-11.09 [0.00]**	0.23 [0.00]**	1.11 [0.00]**	0.32 [0.05]	-0.10 [0.00]**
FIN	-0.20 [0.14]	-2.76 [0.12]	-0.18 [0.11]	-0.13 [0.21]	-0.44 [0.00]**	-0.85 [0.04]*
FRA	0.00 [0.97]	-0.44 [0.46]	0.02 [0.48]	-0.02 [0.72]	0.002 [0.96]	-0.02 [0.21]
GER	-0.64 [0.04]*	22.27 [0.02]*	0.55 [0.00]**	0.06 [0.05]	0.15 [0.27]	-0.15 [0.00]**
IRE	0.51 [0.49]	-7.69 [0.23]	-0.30 [0.17]	-0.70 [0.19]	-0.22 [0.94]	0.19 [0.05]
ITA	0.01 [0.53]	0.12 [0.71]	0.02 [0.39]	0.02 [0.39]	0.005 [0.89]	-0.04 [0.46]
NET	-4.28 [0.26]	-2.28 [0.70]	-0.10 [0.19]	-0.60 [0.00]**	-0.11 [0.68]	1.83 [0.05]
POR	1.18 [0.06]	32.81 [0.00]**	0.69 [0.04]*	-0.08 [0.56]	-0.71 [0.01]*	0.45 [0.00]**
SPA	-0.28 [0.02]*	1.59 [0.96]	-0.13 [0.47]	-0.16 [0.00]**	-0.30 [0.00]**	-0.22 [0.00]**
Sample period	1969-94	1969-94	1969-95	1969-94	1969-90	1972-97

Notes: See the notes of Table 4. The alternative models are defined as follows:

- Model 'single act dummy' adds an unrestricted dummy variable to the VAR models for each country. This dummy takes on the value of 0 between 1969 and 1985 and the value of 1 between 1986 and 1994.
- Model 'e^V' uses V replaced by e^V.
- Model 'fixed weights' uses 'effective' variables constructed from a fixed- rather than time-varying weighting scheme. The fixed weights are calculated as the sample means of the time-varying weights.
- Model 'real exchange rates' uses real rather than nominal exchange rates to construct the V variables.
- Model '8-term MA' uses $m = 8$ rather than $m = 4$ to construct the V variables.
- Model 'V_{LAG}' uses V_t replaced by $V_t = \ln\left\{\left[(1/m)\sum_{i=1}^{m}(Q_{t-i+1} - Q_{t-i})^2\right]^{1/2}\right\}$.

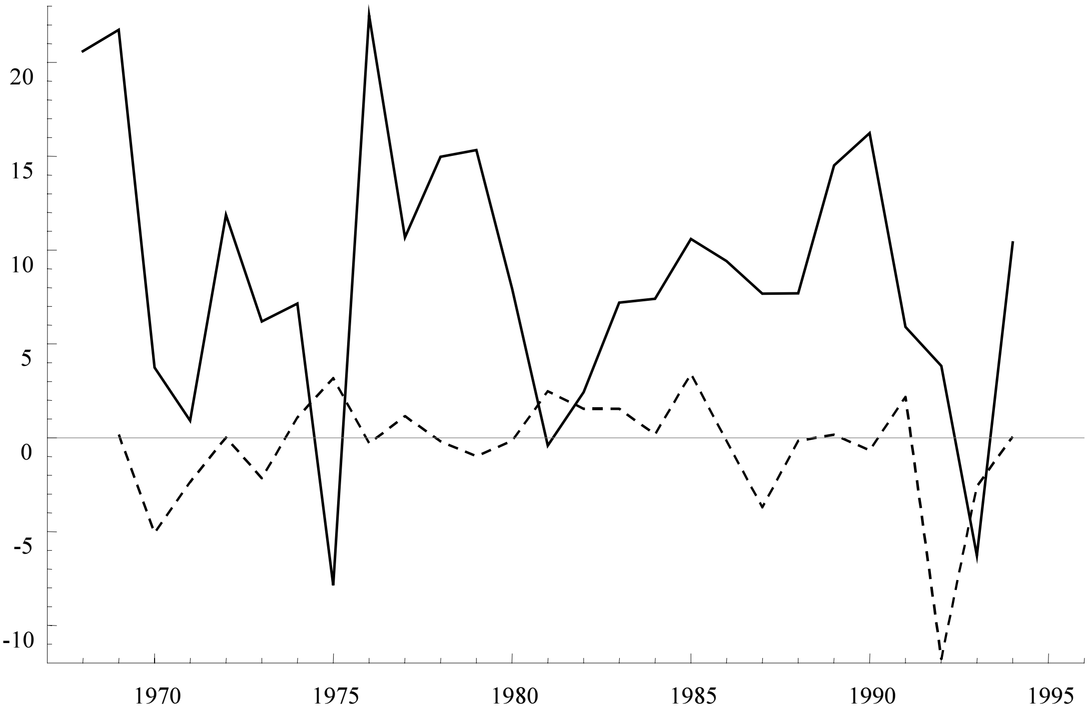
TABLE 7
ANALYSIS OF ROBUSTNESS: QUARTERLY DATA AND JOHANSEN METHOD

$\hat{\delta}_3$ [P Values for LR tests of $H_0 : \delta_3 = 0$]						
Country	V	e^V	Fixed weights	Real exchange rates	8-term MA	V_{LAG}
FIN	-0.04 [0.58]	-0.22 [0.97]	-0.11 [0.69]	0.02 [0.78]	-0.04 [0.36]	0.02 [0.60]
FRA	-0.17 [0.00]**	-20.38 [0.00]**	-0.21 [0.00]**	-0.16 [0.00]**	-0.09 [0.00]**	-0.13 [0.20]
GER	0.29 [0.00]**	29.27 [0.00]**	0.24 [0.00]**	0.28 [0.00]**	0.18 [0.00]**	0.34 [0.00]**
ITA	-0.23 [0.02]*	-13.94 [0.01]*	-0.21 [0.02]*	-0.23 [0.08]	-0.24 [0.00]**	-0.20 [0.05]
NET	0.09 [0.00]**	13.47 [0.00]**	0.08 [0.00]**	0.29 [0.03]*	0.08 [0.03]*	0.20 [0.02]*
POR	0.20 [0.49]	28.17 [0.00]**	0.27 [0.19]	0.22 [0.59]	-0.39 [0.75]	-0.05 [0.49]
SPA	-0.09 [0.09]	-5.11 [0.08]	-0.10 [0.05]	-0.07 [0.36]	0.10 [0.29]	-0.08 [0.15]
Sample period	83:3-96:1	83:3-96:1	83:3-96:1	83:3-96:1	83:3-95:1	84:2-96:4

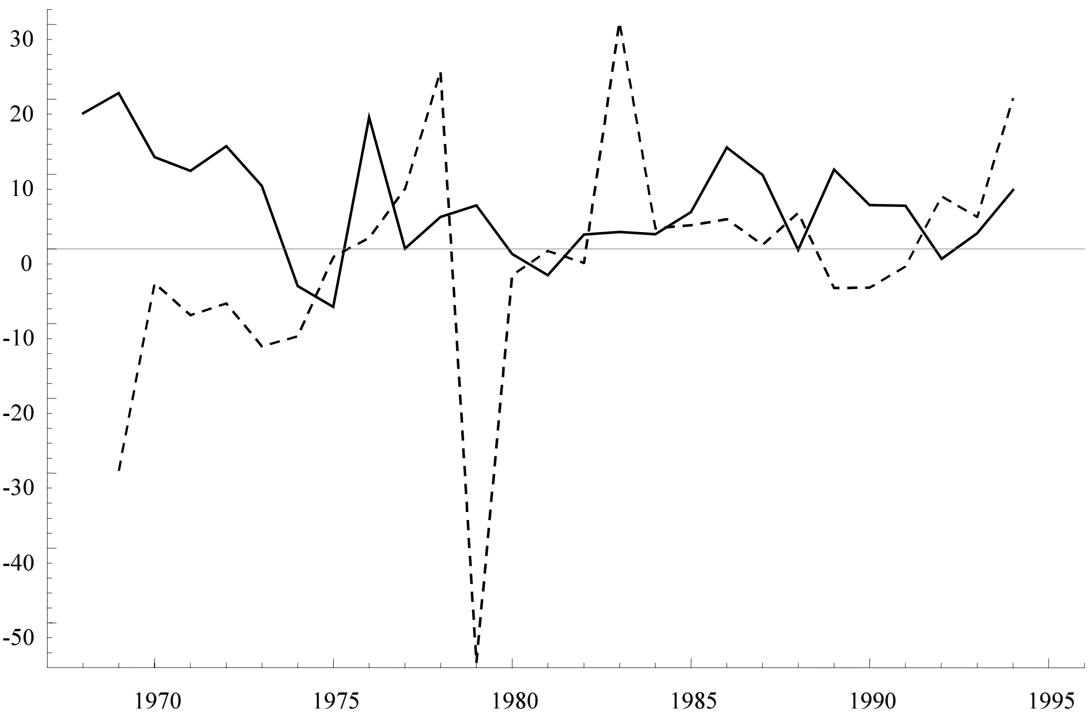
Notes: See the notes of Table 4 and 6. Model 'V' uses formula (1) in the text. The data on exports, exchange rates, and export price indices are from OECD, (1999), the data on consumer price indices are from IMF (1999), and the data on GDP are from *Datastream*, except for Belgium, where data were provided by the Belgian central bank.

FIGURE 1
ACTUAL EXPORT GROWTH IN PER CENT ($100\Delta X_t$, SOLID LINE) AND SCALED EXCHANGE RATE UNCERTAINTY IN PERCENTAGE POINTS ($\hat{\delta}_3 100\Delta V_t$, DASHED LINE)

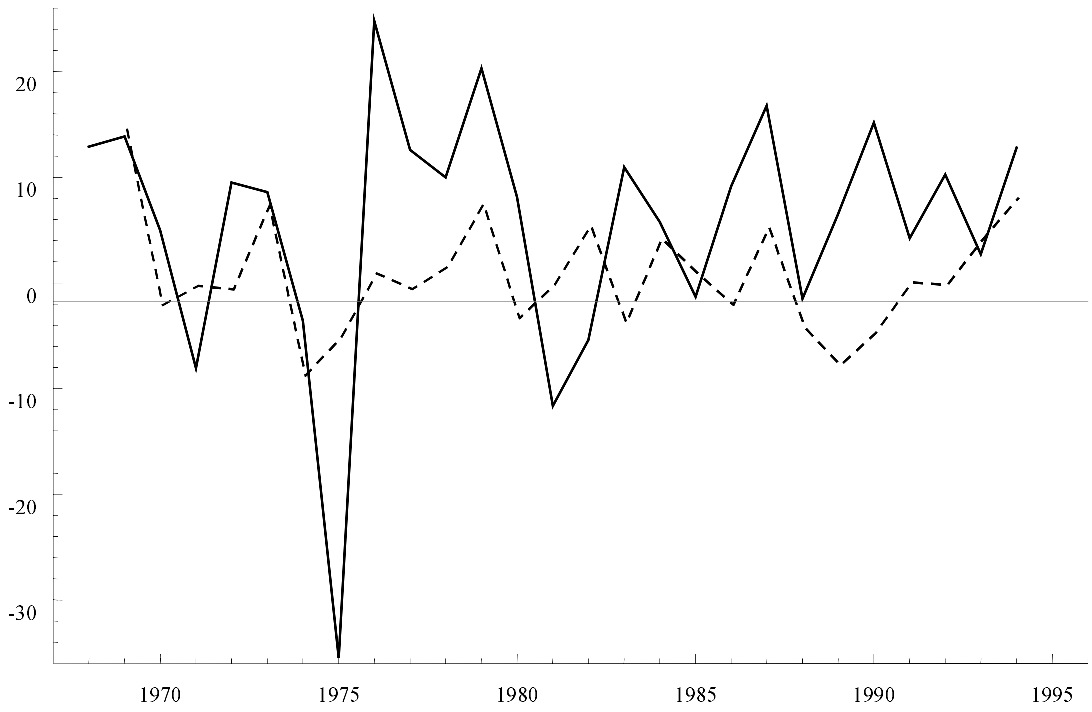
AUS: $\hat{\delta}_3 = -0.10$



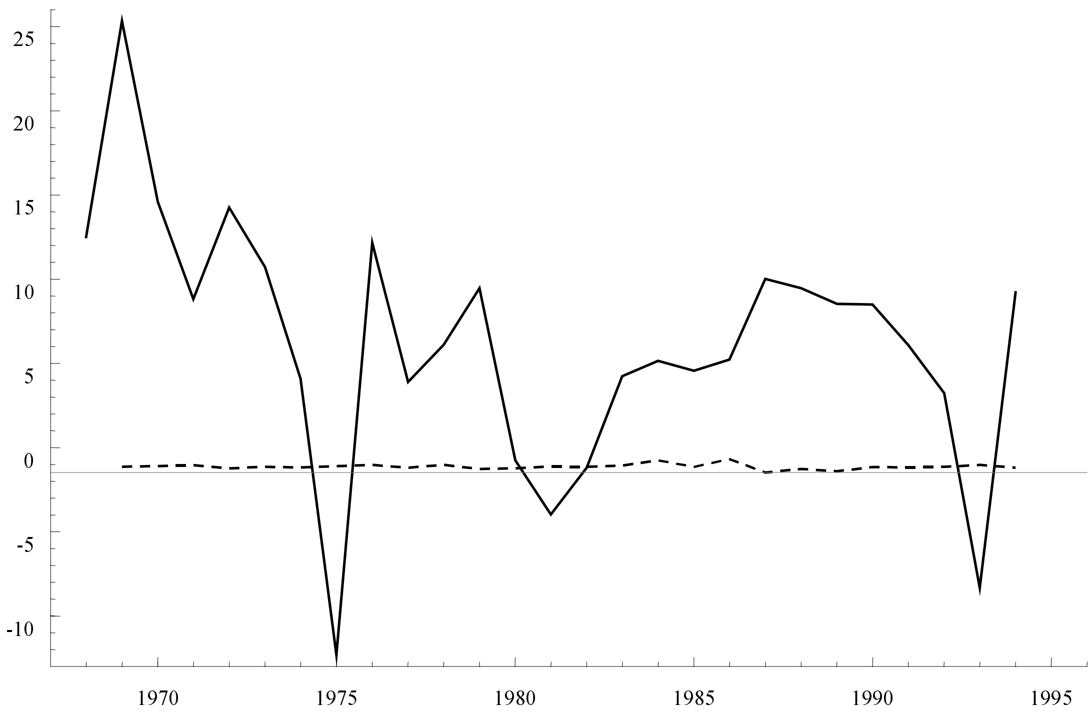
BEL: $\hat{\delta}_3 = -0.45$



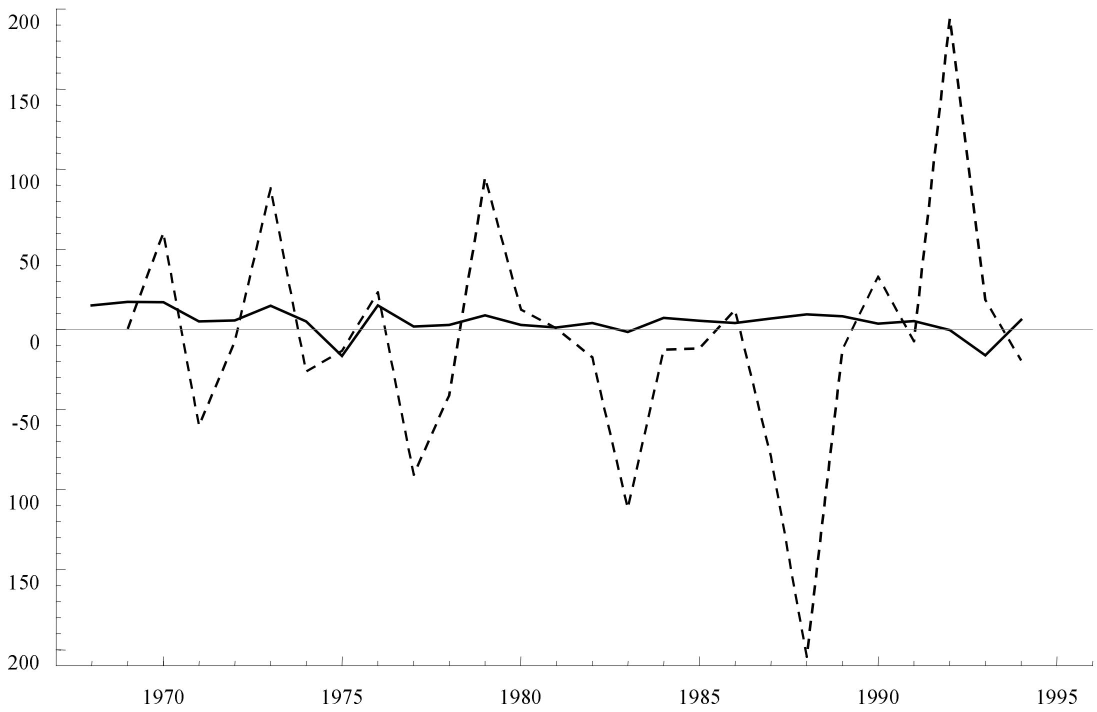
FIN: $\hat{\delta}_3 = -0.18$



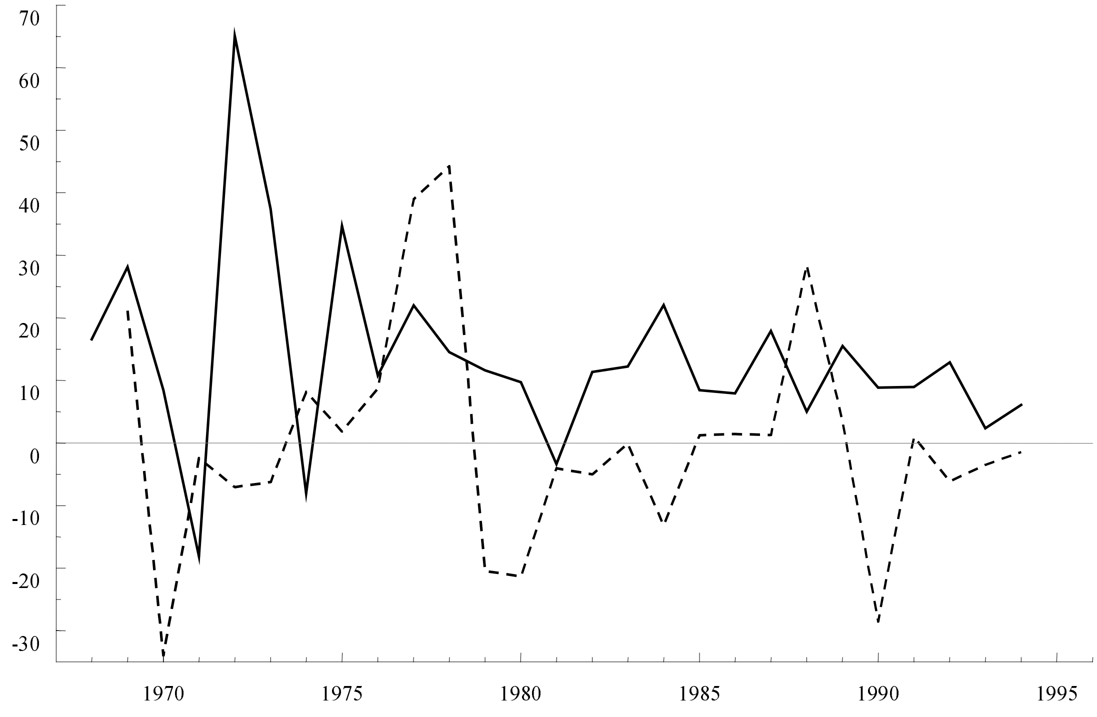
FRA: $\hat{\delta}_3 = -0.003$



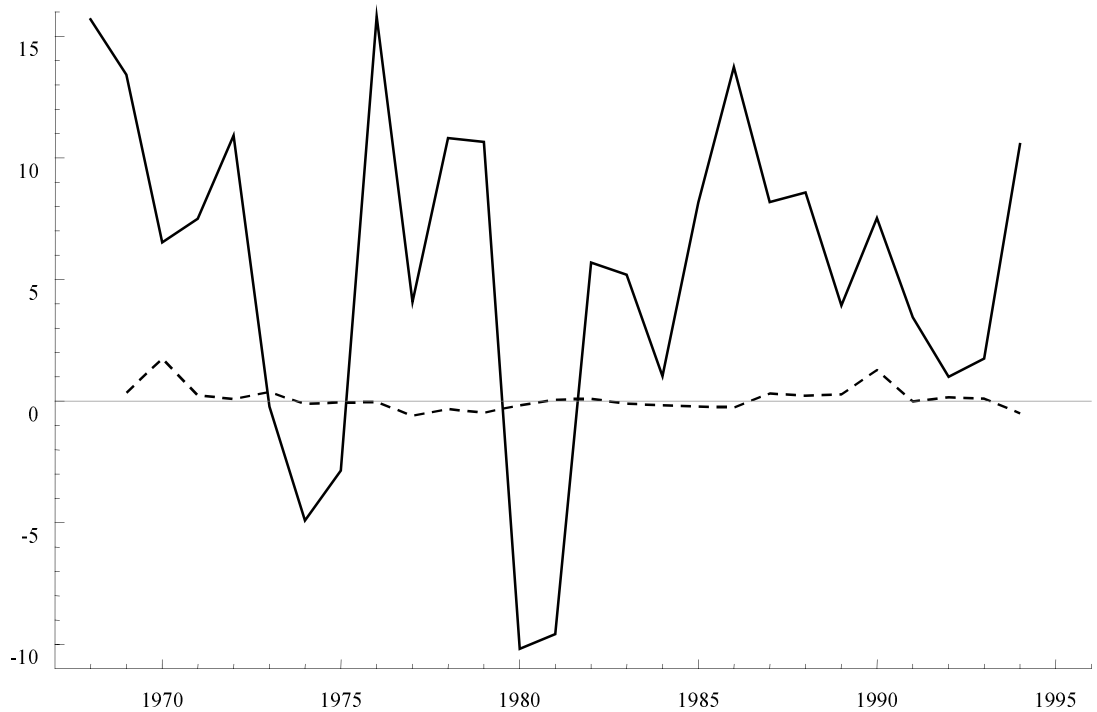
GER: $\hat{\delta}_3 = 3.40$



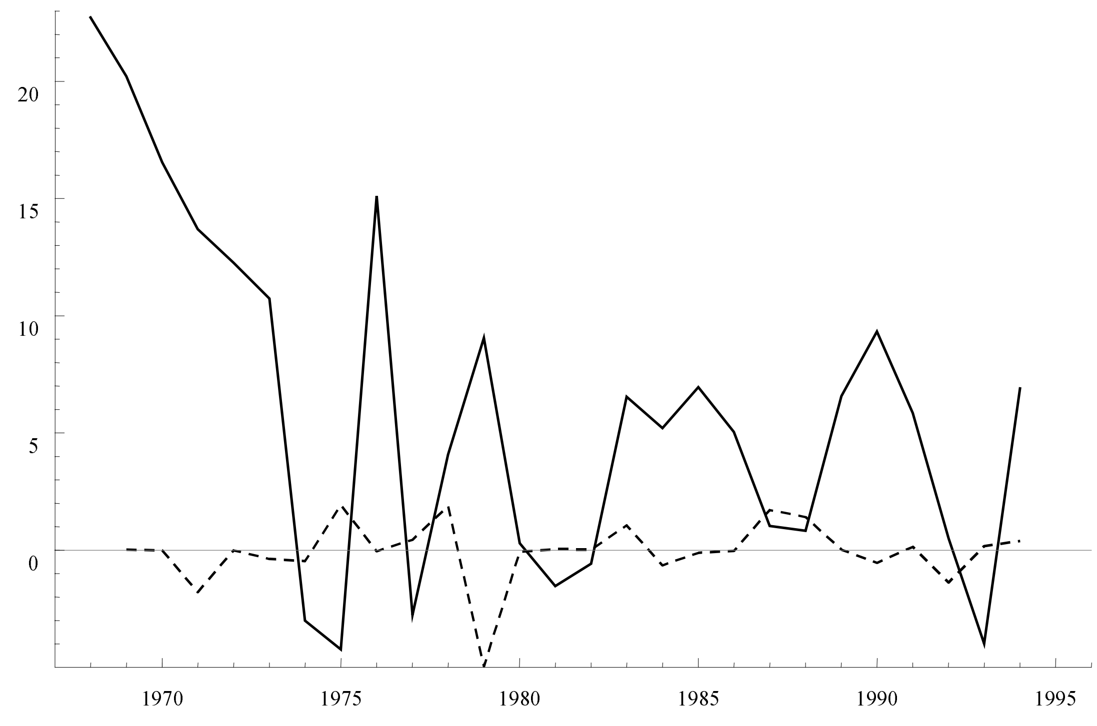
IRE: $\hat{\delta}_3 = -0.42$



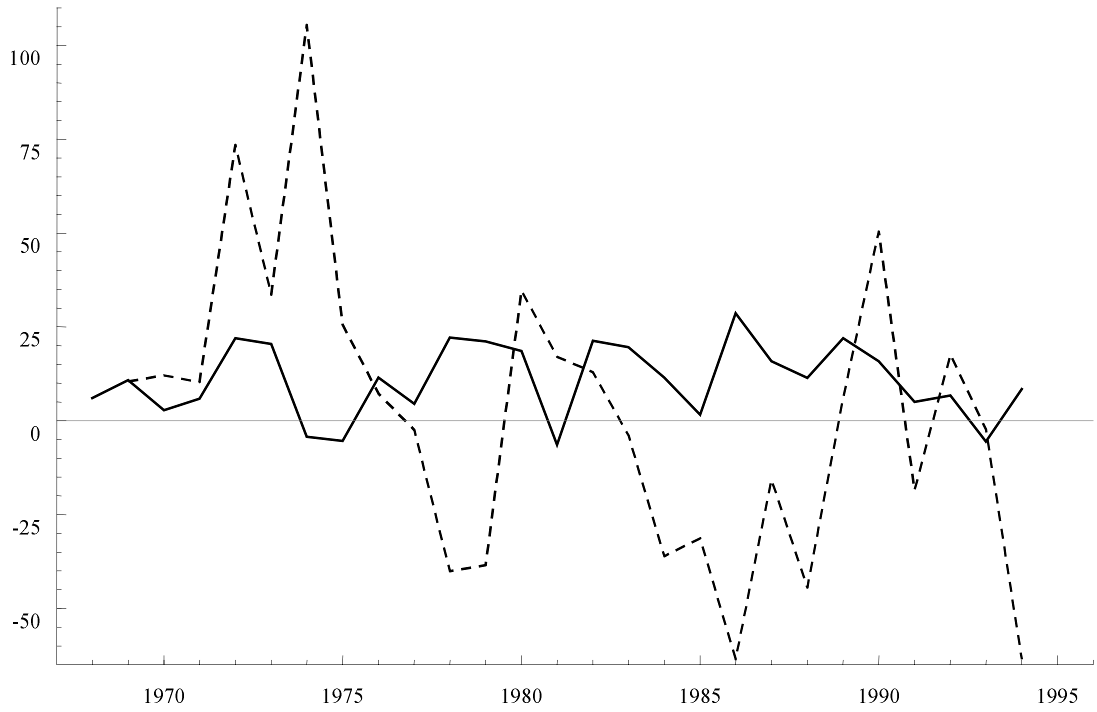
ITA: $\hat{\delta}_3 = 0.01$



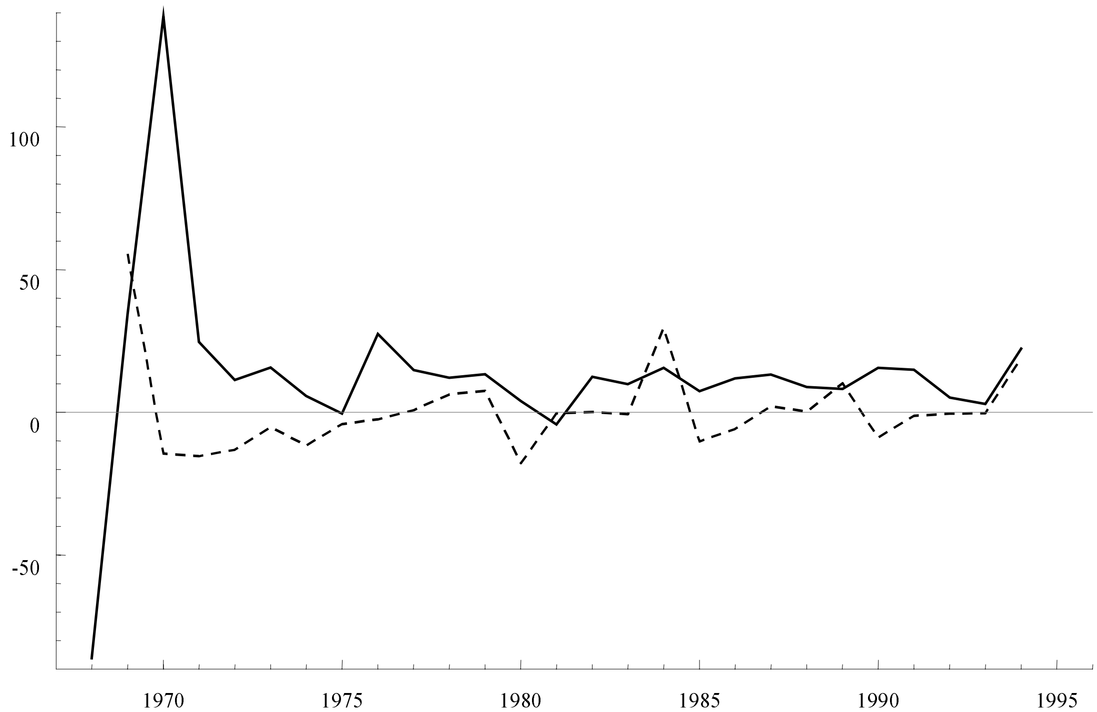
NET: $\hat{\delta}_3 = -0.04$



POR: $\hat{\delta}_3 = 1.16$



SPA: $\hat{\delta}_3 = -0.27$



REFERENCES

Arize, A. (1995). 'The effects of exchange rate volatility on US exports: An empirical investigation'. *Southern Economic Journal,* 62: 34-43.

Arize, A. (1996). 'Real exchange rate volatility and trade flows: The experience of eight European economies'. *International Review of Economics and Finance,* 5: 187-205.

Arize, A. (1997). 'Foreign trade and exchange rate risk in the G-7 countries: Cointegration and error-correction models'. *Review of Financial Economics,* 6: 95-113.

Arize, A. and S. Shwiff (1998). 'Does exchange rate-rate volatility affect import flows in G-7 countries? Evidence from cointegration models'. *Applied Economics,* 30: 1269-76.

Bailey, M. and G. Tavlas (1991). 'Exchange rate volatility and direct investment'. *The Annals AAPSS,* 516: 106-16.

Bini-Smaghi, L. (1991). 'Exchange rate volatility and trade: Why is it so difficult to find any empirical relationship?' *Applied Economics,* 23: 927-36.

Calmfors, L., H. Flam, N. Gottfries, J. Haaland Matlary, M. Jerneck, R. Lindahl, C. Nordh Berntsson, E. Rabinowicz and A. Vredin (1997). *EMU. A Swedish Perspective.* Boston: Kluwer Academic Publishers.

Chowdhury, A. (1993). Does exchange rate volatility depress trade flows? Evidence from error correction models'. *Review of Economics and Statistics,* 75: 700-06.

Committee for the Study of Economic and Monetary Union (Delors Report) (1989). Report on economic and monetary union in the European Community. Luxembourg: Office of publications of the European Communities.

Cushman, D. (1983). 'The effect of real exchange rate risk on international trade'. *Journal of International Economics,* 15: 45-63.

Cushman, D. (1985). 'Real exchange rate risk, expectations, and the level of direct investment'. *Review of Economics and Statistics,* 67: 297-308.

Cushman, D. (1986). 'Exchange rate uncertainty and foreign direct investment in the US'. *Weltwirtschaftliches Archiv,* 124: 322-34.

Cushman, D. (1988). 'US bilateral trade flows and exchange risk during the floating period'. *Journal of International Economics,* 24: 317-30.

Daly, K. (1998). 'Does exchange rate volatility impede the volume of Japan's bilateral trade?'. *Japan and the World Economy,* 10: 333-48.

Darby, J., A. Hughes, J. Ireland and L. Piscitelli (1999). 'The impact of exchange rate uncertainty on the level of investment'. *The Economic Journal,* 109: C55-C67.

De Grauwe, P. (1988). 'Exchange rate volatility and the slowdown in growth in international trade'. *IMF Staff Papers,* 35: 63-84.

Delors, J. (1989). Report on economic and monetary union in the European Community (Delors Report). Committee for the Study of Economic and Monetary Union. Luxembourg: Office of publications of the European Communities.

Dixit, A. and R. Pindyck (1994). *Investment Under Uncertainty.* Princeton: Princeton University Press.

Doornik, J. and H. Hansen (1994). 'A practical test for univariate and multivariate normality'. Discussion Paper. Nuffield College.

Doornik, J. and D. Hendry (1997). *Modelling Dynamic Systems Using PcFiml 9.0 for Windows.* International Thomson Business Press.

Engel, C. and J. Rogers (1996). 'How wide is the border?' *American Economic Review,* 86: 1112-25.

Engle, R., Hendry, D., and J.-F. Richard (1983). 'Exogeneity'. *Econometrica,* 51: 277-304.

European Commission (1990). *European Economy.* 44.

Frankel, J. and A. Rose (1997). 'The endogeneity of the optimum currency area criteria'. *The Economic Journal,* 108: 1009-25.

Gagnon, J. (1993). 'Exchange rate variability and the level of international trade'. *Journal of International Economics,* 34: 269-87.

Goldberg, L. and C. Kolstad (1995). 'Foreign direct investment, exchange rate volatility and demand uncertainty'. *International Economic Review,* 36: 855-73.

Goldstein, M. and M. S. Khan (1985). 'Income and price effects in foreign trade', in R. Jones and P. Kenen (eds.), *Handbook of International Economics,* volume II. Amsterdam: North-Holland.

Gonzalo, J. (1994). 'Five alternative methods of estimating long-run equilibrium relationships'. *Journal of Econometrics,* 60: 203-33.

Gotur, P. (1985). 'Effects of exchange rate volatility on trade: Some further evidence'. *IMF Staff Papers,* 32: 475-512.

Hargreaves, C. (1994). 'A review of methods of estimating cointegrating relationships', in C. Hargreaves (ed.), *Nonstationary Time Series Analysis and Cointegration*. Advanced Texts in Econometrics. Oxford: Oxford University Press.

Hooper, P. and S. Kohlhagen (1978). 'The effect of exchange rate uncertainty on the prices and volumes of international trade'. *Journal of International Trade,* 8: 483-511.

IMF (1999). *International Financial Statistics.* September. Washington, DC: IMF.

Johansen, S. (1988). 'Statistical analysis of cointegration vectors'. *Journal of Economic Dynamics and Control,* 12: 231-54.

Johansen, S. (1991). 'Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models'. *Econometrica,* 59: 1551-80.

Kenen, P. and D. Rodrik (1986). 'Measuring and analyzing the effects of short-term volatility in real exchange rates'. *Review of Economics and Statistics,* 68: 311-9.

Koray, F. and W. Lastrapes (1989). 'Real exchange rate volatility and US bilateral trade: A VAR approach'. *The Review of Economics and Statistics,* 71: 708-12.

Lastrapes, W. and F. Koray (1990). 'Exchange rate volatility and US multilateral trade flows'. *Journal of Macroeconomics,* 12: 341-63.

McCallum, J. (1995). 'National borders matter: Canada-US regional trade patterns'. *The American Economic Review,* 85: 615-23.

Mckenzie, M. (1998). 'The impact of exchange rate volatility on Australian trade flows'. *Journal of International Financial Markets, Institutions and Money,* 8: 21-38.

Ng, S. and P. Perron (1993). 'Unit root tests in ARMA models with data dependent methods for the selection of the truncation lag'. Montreal: Department of Economics, University of Montreal. Mimeo.

OECD (1999). *Monthly Statistics of Foreign Trade*. October. Paris: OECD.

Osterwald-Lenum, M. (1992). 'A note with quantiles of the asymptotic distribution of the ML cointegration rank test statistics'. *Oxford Bulletin of Economics and Statistics,* 54: 461-72.

Park, J. (1992). 'Canonical cointegrating regressions'. *Econometrica,* 60: 119-43.

Prati, A. and G. Schinasi (1997). 'European monetary union and international capital markets: Structural implications and risks'. IMF Working Paper No. 97/62. Washington, DC: IMF.

Reimers, H.-E. (1992). 'Comparisons of tests for multivariate cointegration'. *Statistical Papers,* 33: 335-59.

Rose, A. (2000). 'One money, one market: Estimating the effect of common currencies on trade'. *Economic Policy*, forthcoming.

Sercu, P. and C. Vanhulle (1992). 'Exchange rate volatility, international trade, and the value of exporting firms'. *Journal of Banking and Finance,* 16: 155-82.