Nonlinear Effects of Exchange Rate Volatility on the Volume of Bilateral Exports

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Abstract

In this paper, we empirically investigate the impact of exchange rate volatility on real international trade flows utilizing a 13–country dataset of monthly bilateral real exports for 1980–1998. We compute one–month–ahead exchange rate volatility from the intra–monthly variations in the exchange rate to better quantify this latent variable. We find that the effect of exchange rate volatility on trade flows is nonlin-

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ear, depending on its interaction with the importing country's volatility of economic activity, and that it varies considerably over the set of country pairs considered.

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1 Introduction

Since the breakdown of the Bretton Woods system of fixed exchange rates, both real and nominal exchange rates have fluctuated widely. This volatility has often been cited by the proponents of managed or fixed exchange rates as detrimental, since in their view exchange rate uncertainty will inevitably depress the volume of international trade by increasing the riskiness of trading activity and negatively affecting the optimal allocation of resources. Several theoretical studies (Ethier (1973), Clark (1973), Baron (1976), Cushman (1986), Peree and Steinherr (1989) to mention a few) have shown that an increase in exchange rate volatility will have adverse effects on the volume of international trade. Contrarily, other models (for example Franke (1991), Sercu and Vanhulle (1992)) have shown that exchange rate volatility may have a positive impact on international trade flows, or ambiguous effects depending on aggregate exposure to currency risk (Viaene and deVries (1992)).

Given these contradictory theoretical predictions, one might appeal to empirical analysis to find out which outcome the data would support. Unfortunately, empirical results do not allow us to arrive at a firm conclusion; as Bacchetta and van Wincoop recently stated, "...the substantial empirical literature examining the link between exchange-rate uncertainty and trade has not found a consistent relationship" (2000, p.1093). The empirical results are, in general, sensitive to the choices of sample period, model specification, form of proxies for exchange rate volatility, and countries considered (developed versus developing).[2,3]

[2] Negative effects of exchange rate uncertainty on trade flows are reported by Cushman (1983, 1986, 1988), Akhtar and Hilton (1984), Thursby and Thursby (1987), Kenen and Rodrik (1986), and Peree and Steinherr (1989), among others, while Hooper and Kohlhagen (1978), Gotur (1985), Koray and Lastrapes (1989), and Gagnon (1993) find insignificant effects. Kroner and Lastrapes (1993), using a multivariate GARCH-in-mean model, report that the reduced-form effects of volatility on export volume and prices vary widely. The estimated effects of GARCH conditional variance of the nominal exchange rate on export flows differ in sign and magnitude across the countries studied.

[3] For a survey of theoretical arguments and empirical findings on the relationship between exchange rate volatility

The premise that there is no clear resolution–analytical nor empirical–as to how exchange rate uncertainty impacts trade flows calls for a fresh look at the relationship. It appears that we can forward our knowledge on this puzzle by addressing some of the potential deficiencies in prior work. First, most of the previous research utilizes aggregate US or G7 export data. We investigate the relationship for a broader set of data. Our 13-country data set, which includes U.S., Canada, Germany, U.K., France, Italy, Japan, Finland, Netherlands, Norway, Spain, Sweden, and Switzerland, consists of bilateral real exports for the period 1980-1998 on a monthly basis in each direction. Hence it is possible to examine dozens of bilateral relationships, and avoid the narrow focus on U.S. or the G7 countries' data that has characterised much of the literature.

Second, as Bini-Smaghi (1991) has stressed: there could be methodological problems, as all empirical analysis incorporates a proxy to capture exchange rate volatility. Most of the previous research uses a moving average standard deviation of the past monthly exchange rates, while others use variants of ARCH models. Our study improves upon much of the literature in its method of quantifying exchange rate volatility. We utilize daily spot exchange rates to compute one month-ahead exchange rate volatility (via a method based on Merton (1980), which is also exploited by Klaassen (1999) in the exchange rate context) from the intra-monthly variations in the exchange rate.[4] This approach provides a more representative measure of the perceived volatility avoiding potential problems, such as the high persistence of real exchange rate shocks when moving average representations are used, or low correlation in volatility when ARCH/GARCH models are applied to quantify exchange rate volatility.

Third, there could be a problem of model misspecification, such as inadequate dynamics and omitted variable bias. Our model uses a flexible Poisson lag specification

and trade flows, see Farell et al. (1983), IMF (1984), and Willett (1986) regarding the literature through the mid-1980s, and Côté (1994) regarding more recent works.

[4] Several authors in the finance literature have used high frequency data to obtain volatility measures (e.g., Anderson et al. (2001), French et al. (1987)). Klaassen (1999), using G7 data, demonstrates that proxies obtained from both ARCH models and moving standard deviation measures have conflicting implications for the evolution of risk over time.

to allow the data to determine the appropriate dynamic specification of the time form of explanatory variables' impacts. To address potential omitted–variable bias in similar studies, we introduce a new variable, foreign income uncertainty, and investigate if the impact of exchange rate uncertainty fades or intensifies as uncertainty in foreign income levels varies. The introduction of this variable in the model is in line with the literature which considers entry/exit costs and evaluate producers' "real options" to participate in international trade. In particular, higher volatility of foreign income may signal greater profit opportunities, and thus may have important implications for exporters' behavior. We therefore include a proxy for income volatility, as well as the interaction term of foreign income volatility and exchange rate volatility, which can help us capture any possible nonlinearities and/or indirect effects in the relationship between exchange rate uncertainty and bilateral trade flows.

Our analysis reveals that the relationship between exchange rate volatility and bilateral trade flows appears to be more complicated than the prevalent approach demonstrated by the empirical literature in this field. The data support the concept that the link between exchange rate volatility and volume of exports is clearly not linear. In particular, for some country pairs considered, exchange rate volatility has a meaningful indirect effect on bilateral trade flows through the interaction with income volatility. Furthermore, given the structure of our empirical approach, we show that uncertainty in foreign income may itself play an important role in the determination of trade flows. Although the magnitude and sign of the effect of income volatility on trade varies across the bilateral relationships, it clearly differs from zero in many of those relations, implying that its effect must be considered.

The rest of the paper is constructed as follows. Section 2 presents the model to analyze exchange rate effects on the volume of international trade, and the empirical approach. Section 3 documents our empirical findings and Section 4 concludes and draws implications for future theoretical and empirical research.

2 Modeling exchange rate effects on the volume of trade flows

We consider the bilateral trade flows of an $n-$country world, and focus entirely on the exports of country i to country j, expressed in real terms. A nominal export flow X_{ijt}, denominated in domestic currency, is the product of the export quantity Q_{ijt} and the export price P^x_{ijt}. We work with the logarithm of real exports (using lower-case letters to denote logarithms), deflating the value X_{ijt} by the export price index P^x_{ijt} to generate x_{ijt}.[5] Log real exports are determined by several supply and demand factors. A simple supply relation would include only the price of exports relative to that of domestic output,

$$q^s_{xt} = q^s_x \left(p^x_{ijt}\right),$$

while a more elaborate treatment of producers' behaviour would be less myopic. In our extension, we seek to capture the potential effect of the volatilities of exchange rates and foreign income on exporters' supply decisions. The strands of literature which consider entry/exit costs and evaluate "real options" to participate in export markets give rise to additional factors determining medium-run supply (see, for example, Franke (1991)). The value of a real option, like that of any option, is enhanced by volatility in the underlying relationship, and in this case exporters will be sensitive to both the volatility of foreign income and volatility of the exchange rate.[6] For instance, higher volatility in foreign demand may signal greater opportunities to sell in that market, or to expand one's presence in that market where there are fixed costs (e.g., dealer networks, brand name development) related to such participation. Since these decisions involve resource allocation, suppliers will not react instantaneously to changes in volatility (as would, for instance, options prices in financial markets), even when faced with short-

[5] In our empirical work, we must confront the fact that export price indices are not generally available for bilateral trade flows, but only exist at the aggregate (country) level. Thus, we utilize $P^x_{i.t}$ for the exports of country i to all trading partners. Per a reviewer's suggestion, the dissimilarity of comovements of export prices and the general price level makes this approximation less damaging than would the application of a general price deflator.

[6] As Goldstein and Khan (1985) argue, the time lag between the period in which trade decisions are made and the period when the actual trade takes place implies that uncertainty could affect international trade flows.

term profit opportunities. This consideration implies that suppliers' reactions to income volatility should be modeled with a lag. At the same time, any desire to expand market presence requires a simultaneous consideration of the behavior of exchange rates, and the associated risk. To capture these effects, we include foreign income volatility, by itself and in conjunction with the underlying uncertainty about real exchange rates, and express producers' behavior in the form:

$$q_{xt}^{s} = q_{x}^{s}\left(p_{ijt}^{x}, \sigma_{s,t-\tau}\left[s_t\right], \sigma_{y,t-\tau}\left[y_t\right], \sigma_{s,t-\tau}\left[s_t\right] \times \sigma_{y,t-\tau}\left[y_t\right]\right), \tag{1}$$

where σ_s and σ_y capture real exchange rate and foreign income uncertainty, respectively. In equation (1), suppliers' reactions are modelled with a $\tau-$period lag to capture the potential effect of the volatility of exchange rates, the volatility of foreign output, and their joint effect on exporters' supply decisions.

On the demand side, real foreign income (in log terms) is taken as the scale variable, potentially with a lag to represent the delay between purchase and delivery of the goods. If trade decisions are made contractually τ periods prior to the period in which delivery and payment are made, foreign income τ periods prior should be the relevant determinant.[7] Price effects on demand are considered via the price of traded goods relative to that of domestic goods in the importing country. The relative price in logarithmic terms can be expressed as $p_{ijt}^{x} - s_t$, where $s_t = \log(S_t P_{jt}/P_{it})$ is the logarithm of the real exchange rate, P_{it} the exporting country's domestic price level, P_{jt} the importing country's domestic price level, and S_t the nominal spot exchange rate, measured as the domestic currency price of one unit of foreign exchange. Since importers' decisions are made upon the basis of forecasts of the relative price of imported goods, the conditional mean and standard deviation of $\left(p_{ijt}^{x} - s_t\right)$ are both included in the demand equation, with conditioning on the information set $\Omega_{t-\tau}$ to denote the τ-period lag between decision and

[7] There is no reason to believe that the same $\tau-$period lag should apply to both suppliers' reaction to volatility and the delivery lag for purchases. In our empirical implementation, we utilise a flexible lag specification.

consummation of the transaction. The demand relation is then

$$q_{xt}^{d} = q_{x}^{d}\left(y_{jt-\tau}, E_{t-\tau}\left[p_{ijt}^{x} - s_{t}\right], \sigma_{s,t-\tau}\left[p_{ijt}^{x} - s_{t}\right]\right). \tag{2}$$

Equilibrium in the domestic export market equates supply and demand, with the resulting function for real exports being expressed as

$$x_{it} = x\left(y_{jt-\tau}, E_{t-\tau}\left[s_{t}\right], \sigma_{s,t-\tau}\left[s_{t}\right], \sigma_{y,t-\tau}\left[y_{t}\right], \sigma_{s,t-\tau}\left[s_{t}\right] \times \sigma_{y,t-\tau}\left[y_{t}\right]\right) \tag{3}$$

in its simplest form. This equation may be estimated, given appropriate specifications for the expectations terms, measure of volatility, and inherent dynamics of the relationship. We now present some details of the data employed and of those specifications.

2.1 Data

Our primary empirical investigation is carried out with monthly data on bilateral aggregate real exports, in each direction, over the period between January 1980 and December 1998 for 13 countries: U.S., Canada, Germany, U.K., France, Italy, Japan, Finland, Netherlands, Norway, Spain, Sweden, and Switzerland. These data are constructed from bilateral export series available in the IMF's *Directions of Trade Statistics* (DOTS) and export price deflators, wholesale price indices and monthly spot foreign exchange rates from the IMF's *International Financial Statistics* (IFS). The export data are expressed in current US dollars; they are converted to local currency units (LCU) using the spot exchange rate vis-à-vis the US dollar, and deflated by the country's export price deflator to generate real exports. The dependent variable in our model is the natural logarithm of real exports. The real exchange rate is computed from the spot exchange rate and the local and US wholesale price indices, and is expressed in logarithmic form. Since the series entering the computation of the real exchange rate are not seasonally adjusted, the log(real exchange rate) series is adjusted using seasonal dummies.

To utilize the monthly frequency of export data, we must generate a proxy for foreign income volatility. Although it would be possible to use monthly industrial production itself to generate such a proxy (as has often been done in the literature), we chose not to use industrial production in that context, since it provides a limited measure of overall economic activity. As an alternative, we apply the "proportional Denton" benchmarking technique to quarterly real GDP series in order to produce monthly GDP estimates.[8] The proportional Denton benchmarking technique (Bloem et al., 2001) uses the higher–frequency movements of an associated variable–in our case, monthly industrial production–as an interpolator within the quarter, while enforcing the constraint that the sum of monthly GDP flows equals the observed quarterly total.[9] From the constructed monthly GDP series, the volatility of foreign income is estimated for each country using a moving window technique similar to that employed by Thursby and Thursby (1987, p.491) in which the logarithm of monthly real GDP is regressed on a quadratic trend for a six–month moving window.[10] The root mean squared error of the moving-window regression over observations $(\tau - 1, ..., \tau - 6)$ is used as the estimate of income volatility for period τ. Although this is an adaptive measure of volatility, we believe that it is appropriate for a real–sector volatility series, which is observed only infrequently compared to financial–sector series.

In order to produce a measure of exchange rate volatility that avoids the shortcomings of an ARCH or ARIMA approach fit at the monthly frequency, we employ daily data on spot exchange rates for each country vis-à-vis the US dollar, available from the Board of Governors of the Federal Reserve System. The specific details of that methodology are described next.

[8] The breadth of the data set used in this study is limited by the availability of quarterly GDP series over the sample period.

[9] This calculation was performed by the Stata routine *dentonmq*, written by and available upon request from the first author.

[10] Six months was taken as the fewest number of monthly observations from which a reasonable estimate of the model could be fit.

2.2 Generating volatility measures from daily data

Any attempt to evaluate the effects of exchange rate risk on trade flows requires specification of a measure of risk. The choice of a particular specification may have a considerable impact on the empirical findings; counterintuitive results may be merely reflecting errors of measurement in a proxy for risk. A number of competing specifications may be found in this empirical literature. A commonly employed measure of risk is a moving standard deviation of exchange rate changes, at the same frequency as the data: for instance, including the past 12 or 24 months' exchange rate changes in the context of monthly data. This measure, with equal weights placed on past changes, gives rise to substantial serial correlation in the summary measure. A more sophisticated approach utilises the ability of GARCH models to mimic the "volatility clustering" often found in high-frequency financial series. However, the most successful applications of GARCH modelling are generally at high frequency (daily or intra–daily), and a GARCH model fitted to monthly data may find very weak persistence of shocks. Although an export equation such as (3) must be estimated from monthly (or lower–frequency) data, we use (squared) intra–monthly changes in the exchange rate in order to capture that month's volatility, as originally proposed by Merton (1980). Since this measure may be calculated for real exchange rates between each pair of countries in our dataset, we may focus on the risk faced by country i's exporters selling in the jth export market, which will vary across j, and is quite distinct from the risk faced by country j's exporters selling to country i.

In order to employ Merton's methodology to the problem at hand, we must evaluate the intra–month volatility of the real exchange rate from daily data. Although the spot exchange rate is available at a business–daily frequency, the foreign and US price indices are not. Accordingly, we linearly interpolate the relative price for all countries (the ratio of foreign to US wholesale price indices) within the month, taking account of non-business days in the calculation.[11] The resulting estimate of the daily relative price

[11] Since none of the countries in our sample have experienced severe inflationary episodes during the sample period, it is fair to say that the movements of the real exchange rate, month to month, are dominated by movements in spot

is used to form the daily real exchange rate $\left(s_t^d\right)$, expressed in logarithmic form. The squared first difference of that measure (after dividing by the square root of the number of days intervening) is then defined as the daily contribution to monthly volatility:

$$\varsigma_t^d = \left(100\frac{\Delta s_t^d}{\sqrt{\Delta\phi_t}}\right)^2, \tag{4}$$

where the denominator expresses the effect of calendar time elapsing between observations on the s process. If data were available every calendar day, $\Delta\phi_t$ would always be unity, but since exchange rate data are not available on weekends and holidays, $\Delta\phi_t \in (1,5)$. The estimated monthly volatility of the (log) real exchange series is defined as $\Phi_t\left[s_t\right] = \sqrt{\sum_{t=1}^T \varsigma_t^d}$ where the time index for $\sigma_t\left[s_t\right]$ is at the monthly frequency.

The estimation of equation (3) requires an evaluation of the multiperiod (τ-period-ahead) forecast of $\sigma_{s,t-\tau}\left[s_t\right]$. We employ an AR(2) process in the $\Phi_t\left[s_t\right]$ values, generated from daily data, to define the risk measure:

$$\sigma_{t-1}\left[s_t\right] = \mu + \sum_{i=1}^{2}\gamma_i\left(\Phi_{t-i}\left[s_t\right]-\mu\right), \tag{5}$$

where the parameters μ, γ_1,γ_2 are estimated from the empirical distribution of the $\Phi_t\left[s_t\right]$ values.[12] A $\tau-$period–ahead forecast of $\sigma_{s,t-\tau}\left[s_t\right]$ is then computed by generating a standard multi–period–ahead AR(2) forecast from equation (5), following Hamilton (1994, pp. 80–81).

2.3 Modeling the dynamics of the export relationship

Given that the inclusion of exchange rate volatility in equation (3) arises due to time lags between agents' decisions to purchase and the completion of that transaction, an eclectic

exchange rates, which are in many cases much larger than monthly changes in consumer prices. Furthermore, it is firmly established in the literature that spot exchange rates exhibit high–frequency characteristics. Thus, the absence of a daily measure of relative prices should not be damaging to our calculation of monthly real exchange rate volatility.

[12] Since there is no residual autocorrelation exhibited by the estimates of this model, we conclude that the order of the process is adequate.

specification that allows for considerable freedom in the lag structure is quite important, especially in light of aggregation issues. Consequently, we have implemented a form of the Poisson lag, as described below, for it would appear that the advantages of this flexible lag specification are considerable. Those advantages do not include computational simplicity, so that our specification search for alternative appropriate forms of the model have been constrained by the considerable effort of setting up, estimating and evaluating the model over our broad set of country pairs. However, this approach allows one to estimate the dynamic pattern: the importance of time lags in the effects of income, relative prices, and volatilities upon exports. Therefore, unlike the analytical development in equation (3), we allow the data to specify the time form of the response in the context of this fairly flexible lag structure.

Under the assumption of linearity, the model that we estimate for the exports from country i to country j (country subscripts suppressed) takes the following specification:

$$\begin{aligned} x_t &= \beta_0 + \\ & \sum_{\tau=1}^{\infty} \left[\beta_{1\tau} y_{t-\tau} + \beta_{2\tau} E_{t-\tau}\left[s_t\right] + \beta_{3\tau}\sigma_{s,t-\tau}\left[s_t\right] + \beta_{4\tau}\sigma_{s,t-\tau}\left[s_t\right] \times \sigma_{y,t-\tau}\left[y_t\right] + \beta_{5\tau}\sigma_{y,t-\tau}\left[y_t\right]\right] + \epsilon_t. \end{aligned} \tag{6}$$

To render the model estimable, an explicit distributed lag structure must be employed to constrain the infinite sequences of β coefficients: for instance, to consider only L periods' past values in the estimation. Restrictive specifications–such as those imposing monotonicity, linearity, or exponential decay upon the lag coefficients–may be quite harmful, and inadequate dynamics embedded in the model's specification will almost surely result in damaging omitted-variable bias. In a model with interrelated regressors, it may be necessary to impose similar lag structures on related variables (particularly in our expanded specification, which makes use of an interaction term). To provide for a reasonably parsimonious structure, we employ the following form of a Poisson lag:

$$\beta_{k\tau} = \beta_k^* \left[\frac{(\lambda_k - 1)^{\tau-1}}{(\tau-1)!} e^{-(\lambda_k - 1)} \right], \tag{7}$$

where $\lambda_k > 1$ and k indexes the explanatory variables, each of which is associated with a β^* parameter. This lag structure encompasses several more restrictive alternatives, such as the geometric lag. In the estimation, constraints are imposed on the vector of λ parameters related to the exchange rate so that the same λ is used for the expected real exchange rate $E_{t-\tau}[s_t]$, its volatility $\sigma_{s,t-\tau}[s_t]$, the volatility of foreign income $\sigma_{y,t-\tau}[y_t]$, and the volatility interaction term $\sigma_{s,t-\tau}[s_t] \times \sigma_{y,t-\tau}[y_t]$.[13]

This approach allows us to parsimoniously capture declining as well as hump-shaped lag structures, and permits the mean lag length of the distributed lag to be estimated endogenously rather than imposed on the data. A tradeoff exists between the length of lag allowed in the Poisson specification and the sample size over which the model is fit; we found that allowing up to L =30 months' lag generated sensible results in almost all cases. The bracketed term in equation (7) can be considered the weight placed on period $t-\tau$'s value of the regressor: ϖ_τ, where $\varpi_\tau > 0$ and $\sum_{\tau=1}^{\infty} \varpi_\tau = 1$. The latter constraint is not imposed in the estimation–as it is a constraint on the infinite sequence of Poisson lag coefficients, not the finite subset utilised in the model–but may be evaluated from the estimated parameters, in terms of the value of $\sum_{\tau=1}^{L} \hat{\varpi}_\tau$.[14]

Our estimation, performed with Stata's nonlinear least squares (*nl*) estimation algorithm, also incorporates a trend term as well as three "ERM dummies" to capture trend and structural changes in the long-term behavior of real exchange rates over the years, respectively. The "ERM dummies" pick up the potential structural changes caused by

[13] We also estimated a model allowing for separate λ coefficients for foreign income, exchange rate, and the volatility terms. The λ coefficient for foreign income is similar to what we report in Table 5, and the λ coefficients for exchange rate and volatility terms are not significantly different from each other. These results are available from the authors upon request.

[14] In the estimation of 149 models' Poisson lag distributions for the time form of foreign income effects, only nine had an empirical sum of weights less than 0.99, with five of those less than 0.95. For the other λ parameter, imposed on the time form of the real exchange rate, its volatility, foreign income volatility and their interaction term, 25 of the 149 models had a sum of weights less than 0.99, and 16 of those were less than 0.95. For the preponderance of models, the empirical lag distribution conforms to the theoretical restriction that the weights sum to unity.

three salient events during our sample period: the October, 1990 German unification and entry of the UK; the September, 1992 suspension of ERM participation by Italy and the UK; and the August, 1993 widening of ERM margins to 15%. These events were abstracted from the timeline developed by dell'Aricia (1999, p. 331-332). To conserve space, the coefficients on those terms and their significance are not reported below, but their importance in the specification holds up throughout the set of models estimated.

3 Empirical findings

3.1 Descriptive measures

The nature of our dataset, containing bilateral measures of real exports and exchange rate volatility, permits scrutiny of the common features, as well as the dissimilarities, of these data over the sample period. Since much of the empirical literature has focused only on aggregate exports rather than bilateral trade flows, we first present evidence of the degree to which aggregation might be masking the idiosyncracies present in the bilateral relationships. We take into account measures of exchange rate volatility and income volatility as well as the interaction between these two terms.

We first consider an aggregate measure of exchange rate volatility, constructed from the bilateral measures as a trade-weighted per-period average of the volatilities pertaining to each partner country's exchange rate. These trade-weighted measures have been calculated for the entire set of 13 countries; Table 1 presents the correlations among those measures for the G7 countries. These correlations show that similar volatility patterns are experienced by, for instance, Canada and Japan, perhaps reflecting both countries' sizable exports to the US and European markets. Some intra–European correlations are also quite high, but it may be seen that the trade–weighted volatilities faced by European exporters–including the UK–are quite dissimilar to those faced by US exporters.

1. Correlations among trade-weighted exchange rate volatility measures

	UK	US	FR	DE	IT	CA	JP
UK	1.000						
US	0.528	1.000					
FR	0.594	0.450	1.000				
DE	0.777	0.631	0.840	1.000			
IT	0.572	0.407	0.710	0.803	1.000		
CA	0.380	0.802	0.332	0.465	0.334	1.000	
JP	0.389	0.817	0.303	0.465	0.239	0.616	1.000

To evaluate how these trade-weighted aggregate volatility measures may differ from their bilateral components, we consider one exporting country, the UK, and its G7 trading partners. Table 2 presents the correlations between the UK's trade-weighted series and the bilateral series from each of these trading partners. These correlations show that the UK's trade-weighted volatility is quite closely related to the bilateral components associated with the major European trading partners, and less closely related to those of the overseas G7 partners. Within the bilateral correlations, we see a very high correlation between the risk faced by UK exporters to France and Germany, whereas those risks are not highly correlated with the volatilities faced in exports to the US.

By comparison, Table 3 presents the equivalent correlations for the UK versus six small open economies in the sample: the Netherlands, Norway, Sweden, Switzerland, Finland and Spain. Despite the relative insignificance of trade flows between those smaller countries, Table 3 also presents evidence that volatility in exports to the Netherlands and Switzerland are quite closely related to the trade-weighted measure; Spain and Norway to a lesser degree. Overall, considerable variability in the bilateral series correlations is evident from the table.

In order to provide a visual interpretation of these volatility measures, Figure 1 presents the trade-weighted series for four countries: US, UK, Japan and Spain. The solid line illustrates a smoothed prediction of an AR(1)-with-trend model of the monthly

2. Correlations among bilateral exchange rate volatility measures: G7 vs UK

	UKtw	US	FR	DE	IT	CA	JP
UKtw	1.000						
US	0.640	1.000					
FR	0.871	0.348	1.000				
DE	0.911	0.349	0.894	1.000			
IT	0.815	0.429	0.696	0.736	1.000		
CA	0.681	0.905	0.391	0.420	0.499	1.000	
JP	0.688	0.387	0.511	0.577	0.475	0.505	1.000

3. Correlations among bilateral exchange rate volatility measures: non-G7 vs UK

	UKtw	NL	NO	SU	CH	FI	ES
UKtw	1.000						
NL	0.912	1.000					
NO	0.759	0.674	1.000				
SU	0.569	0.457	0.569	1.000			
CH	0.855	0.907	0.653	0.471	1.000		
FI	0.672	0.566	0.497	0.546	0.523	1.000	
ES	0.836	0.767	0.597	0.407	0.712	0.585	1.000

volatility series.[15] Although there are some similarities among the figures, the differences across countries appears evident. Figure 2 presents four of the bilateral series, relating to the UK's exports to the US, France, Japan and Spain. Examination of the smoothed predictions shows the secular trends in the volatility vis-à-vis the US, falling after mid-1993, and vis-à-vis Japan, steadily rising over the last years of the sample. The spikes in intra-European volatilities relating to the 1993 disruptions of the ERM are also quite apparent. Although each of these features may be visible in the corresponding panel of Figure 1, it should be clear that there is considerable information in the bilateral volatility series that represents idiosyncracies of that country pair. Our modeling of the bilateral trade flows, as a function of bilateral volatilities, takes that information into account.

Continuing our focus on UK exports to the US, France, Japan and Spain, we now present descriptive measures for foreign income volatility and for the volatility interaction term introduced in our model: the product of the exchange rate volatility and foreign income volatility measures. The foreign income volatility measures for these four importing countries are presented in Figure 3, from whence one may see considerably different patterns, reflecting the economic cycles in those three economies. The horizontal line is drawn at $3\hat{\sigma}$ of the empirical distribution for that importing country. Finally, the interaction term itself–the product of the two volatility terms–is presented in Figure 4 for each of these importing countries. Comparing with Figure 2 for the exchange rate volatility alone, it should be evident that the information contained in the interaction term is quite different from that reflected in the exchange rate volatility measure. Hence, one might expect that the interaction term, by taking both of its components into account, would have the ability to explain some phenomena that are not captured by exchange rate volatility alone. Our results in the next section support this conjecture.

[15] We tested higher-order AR models, and found that the parsimonious AR(1) representation, with trend, is superior to more complex AR structures.

3.2 Estimation of the real export equation

Estimation for each directed country pair−that is, the exports of country i to country j−involves the construction of Poisson lag terms, conditional on the values of the λ parameters, within the nonlinear estimation routine. As discussed above, we estimate only two λ_k parameters: λ_y, for the foreign demand variable, and λ_s for the mean and variance of the real exchange rate, as well as the interaction term between exchange rate volatility and foreign income volatility and foreign income volatility itself.[16] The feasible counterpart to equation (6) can then be written as

$$\begin{aligned} x_{ijt} = \; & \alpha_0 + \beta_1 f(y_j; \lambda_y) + \beta_2 f(E_{t-\tau}[s_{ijt}]; \lambda_s) + \beta_3 f(\sigma_{ijs,t-\tau}[s_{ijt}]; \lambda_s) \\ & + \beta_4 f(\sigma_{ijs,t-\tau}[s_{ijt}] \times \sigma_{jy,t-\tau}[y_{jt}]; \lambda_s) + \beta_5 f(\sigma_{jy,t-\tau}[y_{jt}]; \lambda_s) \\ & + \phi_0 t + \phi_1 ERM_1 + \phi_2 ERM_2 + \phi_3 ERM_3 + \epsilon_t, \end{aligned} \tag{8}$$

where $f(\cdot)$ is the Poisson lag function, defined in equation (7), with $\tau \in (1, 30)$. The ERM_k dummy variables are those described above, while the t variable is a time trend to absorb the secular behavior of real exports.

We estimate equation (8) for each i and j, $i \neq j$: a total of 156 models. In a small number of cases, the nonlinear estimation algorithm does not produce a usable solution with a positive definite covariance matrix of the parameter estimates. Since these cases cannot be included in further scrutiny of the estimates, they have been discarded, leaving us with 149 estimated models. Given the sizable number of independently estimated models, we treat the point and interval estimates of the parameters of interest as derived data, and consider their empirical distributions. We present six sets of estimates: those derived from bilateral flows (a) among the G7 countries, (b) among the non-G7 countries, (c) G7 exports to non-G7 countries, (d) non-G7 exports to G7 countries, (e) ERM mem-

[16] This approach allows us to estimate the modal lag at which the exchange rate and volatility measures have the greatest impact on the dependent variable. In earlier estimation of the model, we did not find that separate λ coefficints were distinguishable from each other, but they led to problems of convergence in the nonlinear estimation routine for a number of country–pairs.

ber countries[17] to ERM member countries, and (f) the full set of 149 estimated models among the 13 trading partners.

3.2.1 Exchange rate and income effects on real exports

The modal lag at which the associated regressor(s) has the largest impact on the dependent variable is captured by the estimates of λ. These estimates suggest that there are lengthy lags associated with the effects of both foreign economic activity and the volatility measures on bilateral real exports, implying that these dynamics must be properly taken into account if the relationship is to be modeled appropriately.

Summary estimates of λ are presented in Table 4 for the five subsets and full set of models described above. The first panel of the table shows that over the 149 models, the mean lag inherent in the effects of foreign economic activity on bilateral trade flows ($\hat{\lambda}_y$) is 5.69 months with a 95% interval estimate of $(4.40, 6.97)$, and a corresponding median value of 2.69 months. The mean values for the five subsets do not differ widely from these estimates; although G7–to–G7 exports reflect a slightly smaller mean and median lag. The sequence of lag weights implied by this mean value of $\hat{\lambda}_y$ is illustrated in Figure 5. Note that the lag weights are reasonably symmetric around the mean value, with little effect beyond 9 or 10 months.

The estimates of $\hat{\lambda}_s$ for each of these subsets and the full set of models are presented in summary form in the second panel of Table 4. For $\hat{\lambda}_s$, which represents the lag embedded in the effects of shifts in the distribution of the real exchange rate and volatility measures, the mean value is 14.11 months, with a 95% interval estimate of $(13.13, 15.10)$ and a median value of 14.25 months. Thus, the data indicate that these effects achieve their greatest effect on real exports with more than a year's lag. Interestingly, G7 exports to non–G7 countries show the longest lag—on average 15.22 months, more than six weeks' greater than the G7–to–G7 exports. The mean and median values for intra–ERM trade

[17] ERM member countries are taken as those six who were ERM members at any time during our estimation sample: the UK, France, Germany, Italy, the Netherlands and Spain.

are significantly smaller than those of the other categories. The sequence of lag weights implied by this mean value, illustrated in Figure 5, is symmetric, with a noticeable impact appearing after 7–8 months and dissipating after 21–22 months' delay.

Exporters Importers	G7 G7	G7 nonG7	nonG7 G7	nonG7 nonG7	ERM ERM	All All
λ_y median	2.08	3.41	2.52	3.74	2.90	2.69
λ_y mean	3.46	7.45	5.10	7.01	4.94	5.69
std. error	0.54	1.77	1.02	1.48	1.05	0.65
95% conf.	2.36	3.87	3.04	3.99	2.79	4.40
interval	4.56	11.03	7.15	10.04	7.09	6.97
λ_y minimum	1.00	1.00	1.00	1.00	1.00	1.00
λ_y maximum	15.04	65.58	27.75	27.96	27.96	65.58
λ_s median	13.71	14.47	14.90	14.03	13.49	14.25
λ_s mean	13.56	15.22	13.52	14.12	12.85	14.11
std. error	1.05	0.97	0.91	1.07	1.18	0.50
95% conf.	11.43	13.26	11.68	11.93	10.44	13.13
interval	15.69	17.18	15.37	16.31	15.26	15.10
λ_s minimum	4.34	4.35	2.34	1.00	2.34	1.00
λ_s maximum	28.50	36.40	26.37	30.49	30.49	36.40
N of models	38	41	42	28	30	149

4. Estimation Results: λ_y, λ_s

3.2.2 Effects of exchange rate and income volatilities on real exports

We focus now on the estimated effects of volatility in the real exchange rate and foreign income on real exports, which depend on coefficients β_3, β_4 and β_5 in equation (8). In the presence of the interaction term between real exchange rate volatility and foreign income volatility, these effects are not constant, but rather depend on the values of each form of volatility. That is,

$$\Gamma_{ij} = \frac{\partial x_{ijt}}{\partial \sigma_{ijs,t} \left[s_{ijt} \right]} = \hat{\beta}_3 + \hat{\beta}_4 \sigma_{jy,t} \left[y_{jt} \right], \tag{9}$$

so that, given $\hat{\beta}_4 \neq 0$, the effect of real exchange rate volatility depends on the level of foreign income volatility, with the sign of the interaction being that of $\hat{\beta}_4$. That is, greater foreign income volatility could either enhance or diminish the direct effect of exchange rate volatility, depending on the signs of the two estimated coefficients. We calculate this expression at the mean value of $\sigma_{jy,t}\left[y_{jt}\right]$, and test the hypothesis that it equals zero using a Wald statistic.[18] Likewise, the effect of foreign income volatility depends upon the level of exchange rate volatility:

$$\Lambda_{ij} = \frac{\partial x_{ijt}}{\partial \sigma_{jy,t}\left[y_{jt}\right]} = \hat{\beta}_4 \sigma_{ijs,t}\left[s_{ijt}\right] + \hat{\beta}_5 \tag{10}$$

which we calculate at the mean value of $\sigma_{ijs,t}\left[s_{ijt}\right]$ so that the hypothesis that it equals zero may be tested. Given $\hat{\beta}_4 \neq 0$, the effect of foreign income volatility depends on the level of real exchange rate volatility, with the sign of the interaction again being that of $\hat{\beta}_4$. Greater real exchange rate volatility could either enhance or diminish the direct effect of foreign income volatility, depending on the signs of the two estimated coefficients. Summary statistics for Γ_{ij} and Λ_{ij} estimates, for the five subsets and the full set of models, are presented in Table 5, and the individual estimates are presented in Table 8 in the appendix.

If we consider the entire set of 149 bilateral models, the estimated effect (semi–elasticity) of exchange rate volatility, $\hat{\Gamma}_{ij}$, is generally positive, with a mean value of 0.157, a 95% confidence interval of $(0.079, 0.235)$, and a median value of 0.014. The results for the five subsets have broadly similar values, with the median value in each subset of a similar magnitude. The mean value is positive in all subsets, and significantly different from zero for the nonG7–G7 subset. However, these summary statistics mask the variation in these effects across the 13 exporting countries. Since the theoretical literature maintains that exchange rate volatility may have either positive or negative effects on particular bilateral trade flows, attention should be focused on the degree to which these effects are distinguishable from zero. Hence, in Table 6 we summarize the prevalence of these effects from the exporter's perspective: that is, from each exporting country i, how many of the bilateral relations to importers $j = 1, ..., 13$ exhibit sensitiv-

[18] The test of this hypothesis from the estimated nonlinear model is computed by Stata's *testnl* procedure, which is based on a statistic described in Greene (2000, p. 153-154).

5. Estimation Results: $\Gamma_{ij}, \Lambda_{ij}$

Exporters Importers	G7 G7	G7 nonG7	nonG7 G7	nonG7 nonG7	ERM ERM	All All
Γ_{ij} median	0.015	0.014	0.025	0.003	0.018	0.014
Γ_{ij} mean	0.103	0.144	0.296	0.041	0.251	0.157
std. error	0.060	0.084	0.089	0.062	0.130	0.040
95% conf.	-0.019	-0.025	0.116	-0.086	-0.014	0.079
interval	0.224	0.313	0.477	0.168	0.517	0.235
Γ_{ij} minimum	-0.163	-0.458	-0.139	-0.192	-0.204	-0.458
Γ_{ij} maximum	1.836	2.950	2.088	1.667	2.950	2.950
Λ_{ij} median	-0.005	-0.065	-0.061	-0.063	-0.019	-0.050
Λ_{ij} mean	0.115	0.109	-0.072	-0.033	0.138	0.033
std. error	0.222	0.202	0.299	0.158	0.247	0.119
95% conf.	-0.335	-0.300	-0.676	-0.357	-0.368	-0.202
interval	0.566	0.517	0.532	0.291	0.644	0.267
Λ_{ij} minimum	-2.908	-2.150	-4.598	-1.757	-2.394	-4.598
Λ_{ij} maximum	3.473	3.670	5.680	2.421	3.691	5.680
N of models	38	41	42	28	30	149

ity to exchange rate volatility or to the volatility interaction term. Of the 149 models considered, 37 (or 25%) have values significantly different from zero at the 95% level of significance, with 29 positive and only 8 negative. Scrutiny of the table indicates that the greatest number of significant effects (four) is registered by Switzerland and Japan, each of which have four positive and two negative estimates of $\hat{\Gamma}_{ij}$. In summary, these results suggest that the effect of exchange rate uncertainty on trade is generally positive.

6. Frequency of significant Γ_{ij} estimates from exporters' perspective

	G7: # $\Gamma_{ij} > 0$	G7: # $\Gamma_{ij} < 0$	All: # $\Gamma_{ij} > 0$	All: # $\Gamma_{ij} < 0$
US	2	0	3	0
UK	0	1	2	1
France	1	0	2	0
Germany	0	0	0	0
Italy	0	0	1	0
Netherlands	0	0	2	0
Norway	0	0	1	2
Sweden	0	0	4	0
Switzerland	0	0	4	2
Canada	1	0	1	0
Japan	2	2	4	2
Finland	0	0	2	1
Spain	0	0	3	0

Notes: frequencies are taken from 38 bilateral estimated models of intra-G7 exports (a maximum of 6 models per exporting country) and 149 bilateral estimated models of exports among the full 13-country set (a maximum of 12 models per exporting country).

Our findings, as summarized in the lower panel of Table 5, also indicate that income volatility plays an important role in the determination of bilateral real exports. We find a mean value of 0.033 and a 95% confidence interval of $(-0.202, 0.267)$, with a median value of -0.050, for the full set of results. Although this mean effect cannot be distinguished from zero in its summary form, a sizable number of the estimated coefficients are significantly different from zero, almost evenly split between positive and negative effects. Table 7 presents the frequencies of significant values of Λ_{ij}, the estimated effect of foreign income volatility, for the five subsets and for the full set of trade flows. The effects of foreign income volatility, incorporating the interaction of exchange rate volatility, appear to be important determinants of real exports for 35 models (23%) of the full set, with 19 positive and 16 negative effects significantly different from zero. The largest number of significant effects is that recorded for the US, for which three coefficients are significantly positive, and four negative.

7. Frequency of significant Λ_{ij} estimates from exporters' perspective

	G7: # $\Lambda_{ij} > 0$	G7: # $\Lambda_{ij} < 0$	All: # $\Lambda_{ij} > 0$	All: # $\Lambda_{ij} < 0$
US	1	2	3	4
UK	1	1	4	1
France	0	0	0	2
Germany	0	2	0	3
Italy	2	0	2	0
Netherlands	0	0	0	0
Norway	0	0	1	1
Sweden	0	0	1	0
Switzerland	0	0	0	2
Canada	1	0	2	0
Japan	2	1	3	1
Finland	0	0	1	2
Spain	0	0	2	0

Notes: frequencies are taken from 38 bilateral estimated models of intra-G7 exports (a maximum of 6 models per exporting country) and 149 bilateral estimated models of exports among the full 13-country set (a maximum of 12 models per exporting country).

We provide detailed estimates of Γ_{ij} and Λ_{ij} in Table 8 in the Appendix, but for a visual interpretation of their magnitudes, we present these point estimates in Figure 6 for those 62 models in which one or both estimated effects are distinguishable from zero. It is clear from this plot along with Tables 5 and 6 that the full impact of exchange rate uncertainty on trade flows is generally positive. Nevertheless, we cannot infer a clear sign for the impact of income volatility. There are almost an equal number of cases where the impact is positive or negative, but as the hypothesis that its effect equals zero is clearly refuted by the data, we claim that foreign income volatility plays an important role in the determination of trade flows.

Having laid out our main empirical findings, it might be useful to consider the impact of the interaction term $\hat{\beta}_4$ on the impact of exchange rate uncertainty on trade flows. We find that in 47 of the estimated models, the coefficient on the interaction term ($\hat{\beta}_4$) is significantly different from zero, with 36 positive and 11 negative values. In comparison, the direct effect of exchange rate uncertainty, as captured by $\hat{\beta}_3$, is significantly different from zero in 35 models, with 19 positive and 16 negative values. It is clear that the interaction term is playing an important role on determination of the total impact of exchange rate uncertainty on trade flows. Of those 35 models, there are 28 in which both $\hat{\beta}_3$ and $\hat{\beta}_4$ are significantly different from zero.[19] In 14 of those models, a positive $\hat{\beta}_4$ appears with a negative $\hat{\beta}_3$, suggesting that income volatility is offsetting the negative impact of greater exchange rate volatility on trade. In 10 of the remaining models, both coefficients are positive, so that income volatility reinforces the positive impact of greater exchange rate volatility on trade. In summary, the interaction between the two volatility terms appears strong in most of the models where exchange rate volatility has a distinguishable effect on trade. In line with the disparity in the theoretical literature, that effect may be stimulative or depressive, but where the effect may be detected, its magnitude appears to be influenced by the volatility in foreign demand.

4 Conclusions

In this paper, we attempt to forward our understanding of the impact of exchange rate uncertainty on real exports by addressing some of the potential deficiencies in prior empirical work. We utilize a broader set of data, which allows us to independently examine dozens of bilateral relationships, and employ an alternative measure for exchange rate uncertainty derived from daily spot exchange rates à la Merton (1980) to work with a more plausible measure of perceived volatility. Most importantly, we entertain the idea that exchange rate uncertainty could indirectly impact trade flows working through income volatility, which could be considered as either inhibiting (through revenue

[19] Note that the precision of $\hat{\Gamma}_{ij}$ (from equation (9)) also depends upon the estimated covariance between $\hat{\beta}_3$ and $\hat{\beta}_4$.

uncertainty) or actually enhancing trade prospects (taking a real options perspective on the potential to establish operations in that market).

The diversity of our findings from the estimation of a simple model using a flexible distributed lag structure suggests a clear message: research making use of aggregate measures, which assume that a single, linear relation exists at the aggregate level, is not likely to be successful. The effect of exchange rate uncertainty on trade flows appears complex, working through the interaction of exchange rate volatility with foreign income volatility, with the latter variable playing an important role in its own right. Using monthly bilateral data for 13 developed countries between 1980-1998 we demonstrate that exchange rate volatility has a significant impact on real exports on all but one country (Germany) in our dataset, which can intensify or diminish through changes in foreign income volatility. Yet, we find that on average the total effect of exchange rate uncertainty is positive. Furthermore, given the structure of our empirical approach, we show that uncertainty in foreign demand may itself play an important role in determination of trade flows. However, the total effect of demand volatility on trade is unclear and warrants for further research.

These findings may provide a potential explanation for many studies' difficulties in finding a consistent link between exchange rate volatility and trade flows using a more simplistic measure of exchange rate volatility. We believe that future theoretical and empirical research should further investigate the effect of income volatility (as well as other sources of uncertainty surrounding fundamental variables) on trade flows, while entertaining the notion that the impact of exchange rate uncertainty could be identified as working through these stochastic elements. Furthermore, applying a similar methodology to developing country data would be useful.

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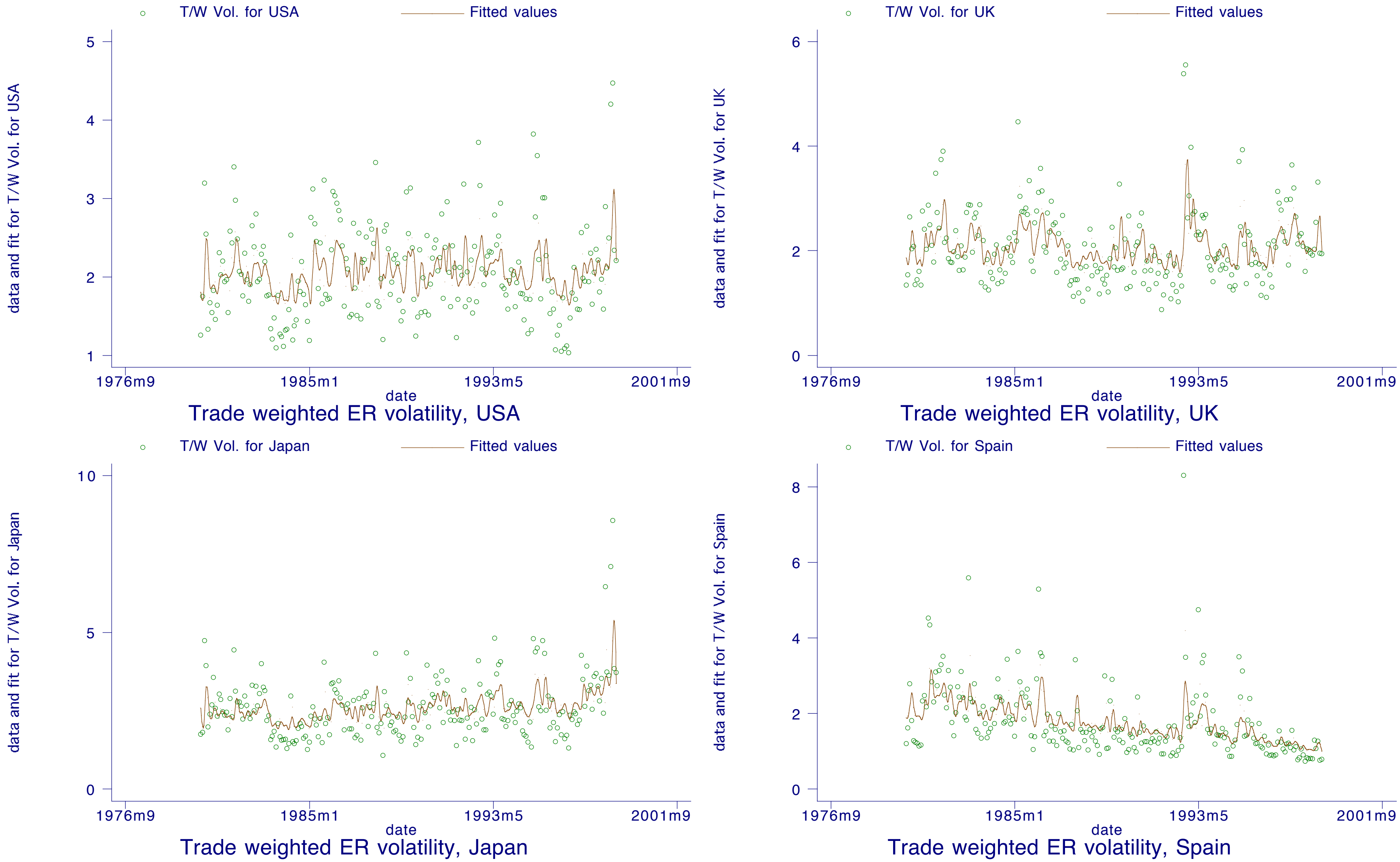

T/W Vol. for USA
Fitted values
data and fit for T/W Vol. for USA
5
4
3
2
1
1976m9
1985m1
1993m5
2001m9
date
Trade weighted ER volatility, USA
T/W Vol. for UK
Fitted values
data and fit for T/W Vol. for UK
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Trade weighted ER volatility, UK
T/W Vol. for Japan
Fitted values
data and fit for T/W Vol. for Japan
10
5
0
1976m9
1985m1
1993m5
2001m9
date
Trade weighted ER volatility, Japan
T/W Vol. for Spain
Fitted values
data and fit for T/W Vol. for Spain
8
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Trade weighted ER volatility, Spain

Figure 1

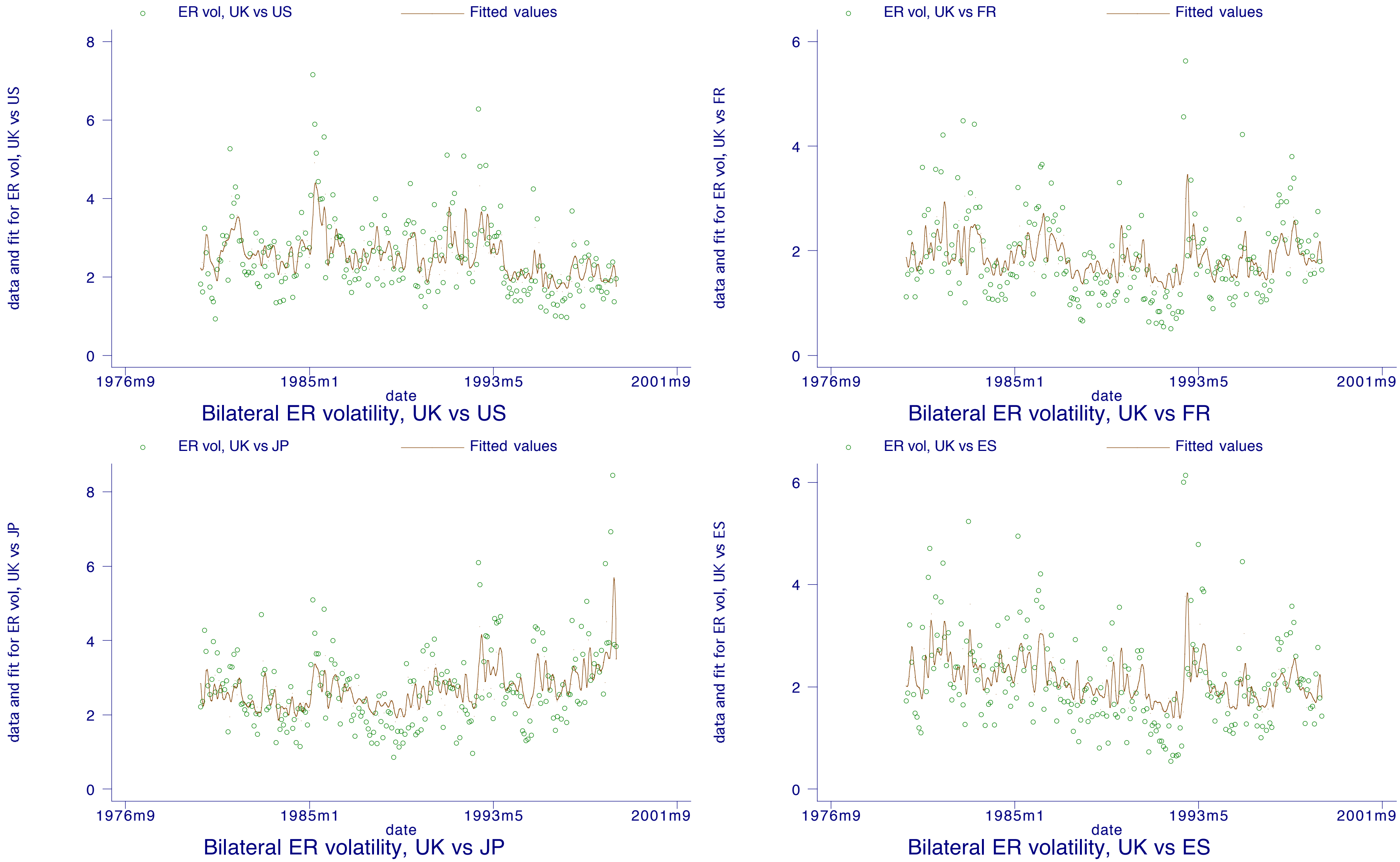

ER vol, UK vs US
Fitted values
data and fit for ER vol, UK vs US
8
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Bilateral ER volatility, UK vs US
ER vol, UK vs FR
Fitted values
data and fit for ER vol, UK vs FR
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Bilateral ER volatility, UK vs FR
ER vol, UK vs JP
Fitted values
data and fit for ER vol, UK vs JP
8
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Bilateral ER volatility, UK vs JP
ER vol, UK vs ES
Fitted values
data and fit for ER vol, UK vs ES
6
4
2
0
1976m9
1985m1
1993m5
2001m9
date
Bilateral ER volatility, UK vs ES

Figure 2

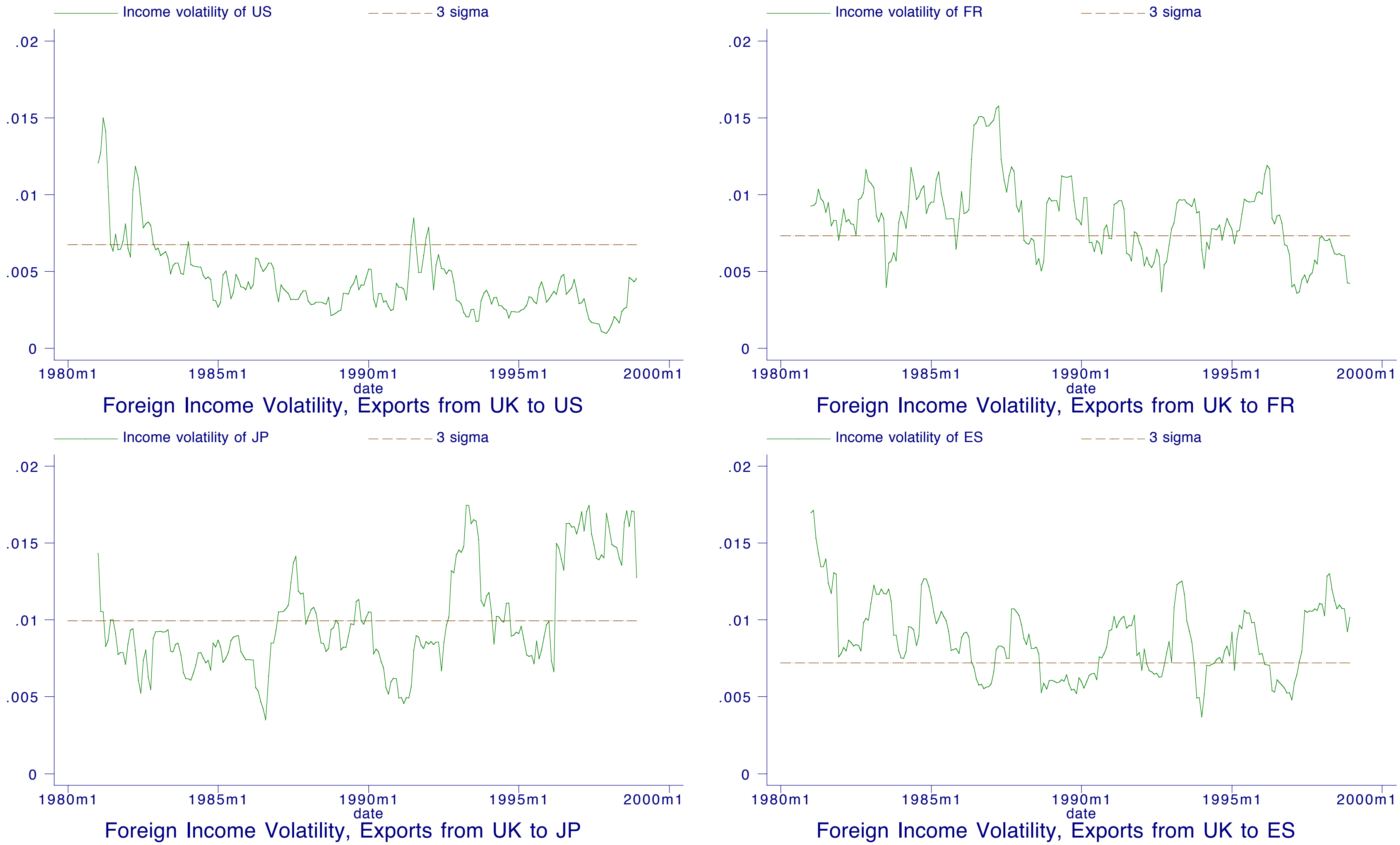

Income volatility of US
3 sigma
.02
.015
.01
.005
0
1980m1
1985m1
1990m1
1995m1
2000m1
date
Foreign Income Volatility, Exports from UK to US
Income volatility of FR
3 sigma
.02
.015
.01
.005
0
1980m1
1985m1
1990m1
1995m1
2000m1
date
Foreign Income Volatility, Exports from UK to FR
Income volatility of JP
3 sigma
.02
.015
.01
.005
0
1980m1
1985m1
1990m1
1995m1
2000m1
date
Foreign Income Volatility, Exports from UK to JP
Income volatility of ES
3 sigma
.02
.015
.01
.005
0
1980m1
1985m1
1990m1
1995m1
2000m1
date
Foreign Income Volatility, Exports from UK to ES

Figure 3

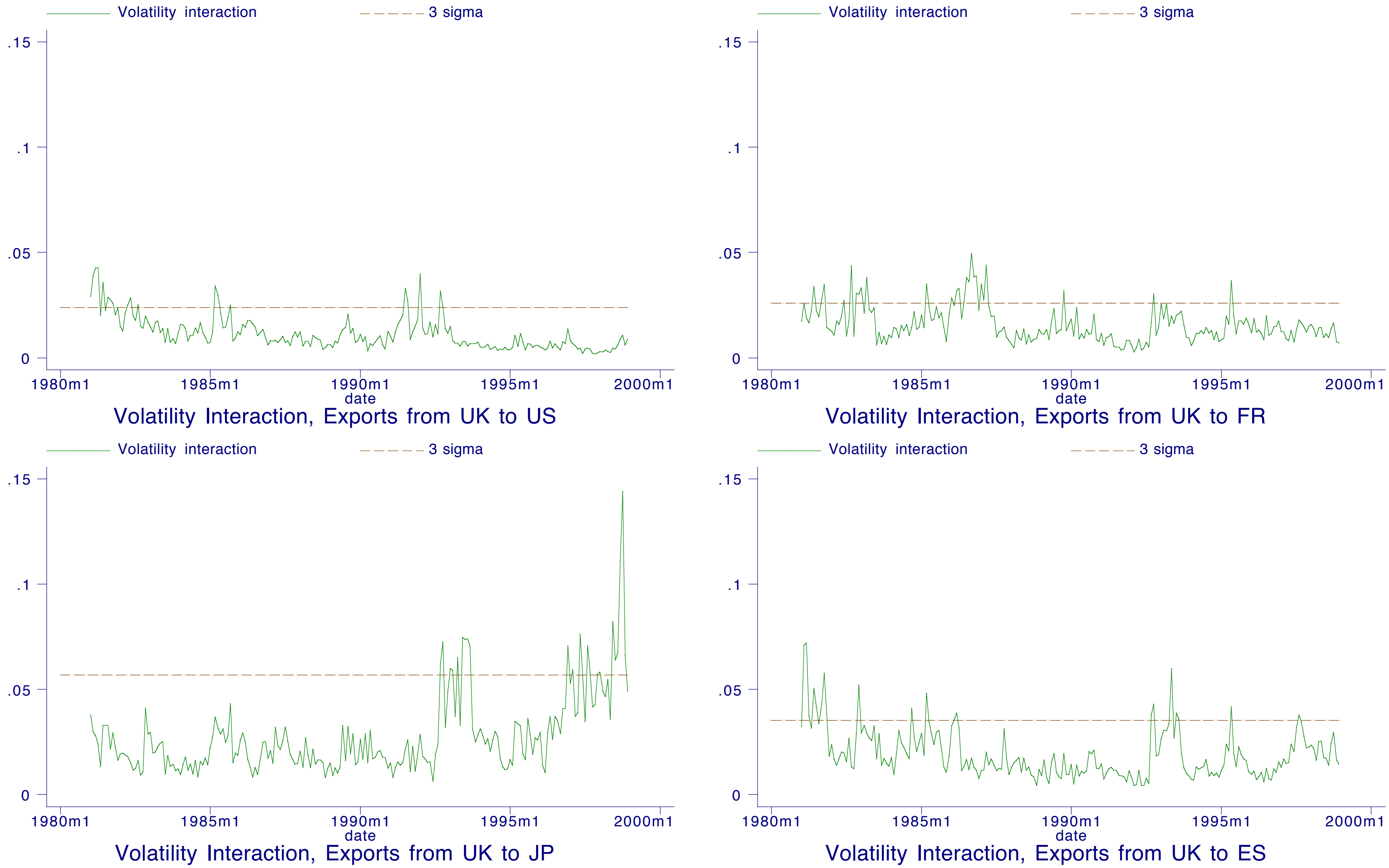
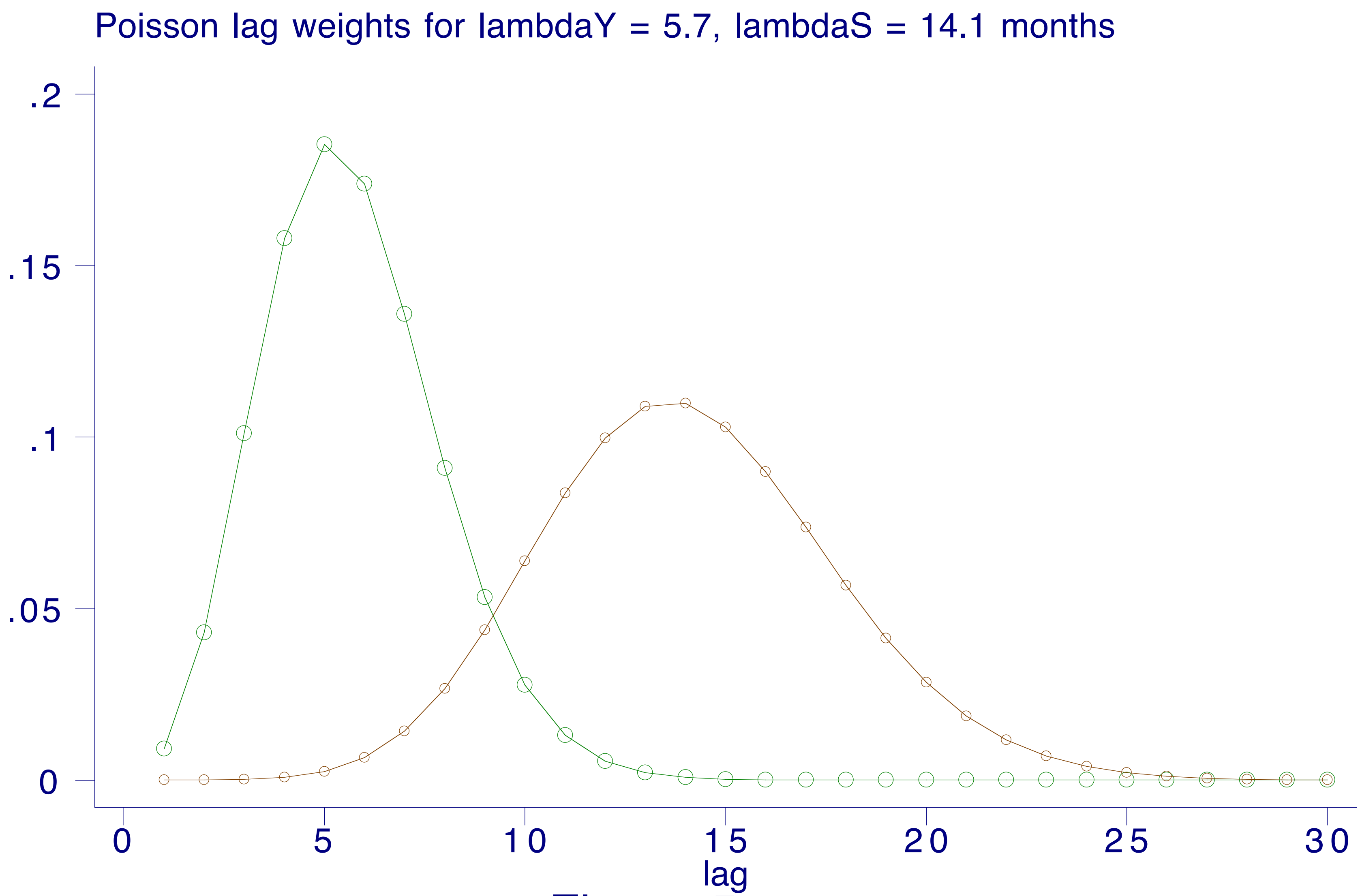

Volatility interaction
3 sigma
.15
.1
.05
0
1980m1
1985m1
1990m1
1995m1
2000m1
date
Volatility Interaction, Exports from UK to US
Volatility Interaction, Exports from UK to FR
Volatility Interaction, Exports from UK to JP
Volatility Interaction, Exports from UK to ES

Figure 4


Poisson lag weights for lambdaY = 5.7, lambdaS = 14.1 months
.2
.15
.1
.05
0
0
5
10
15
20
25
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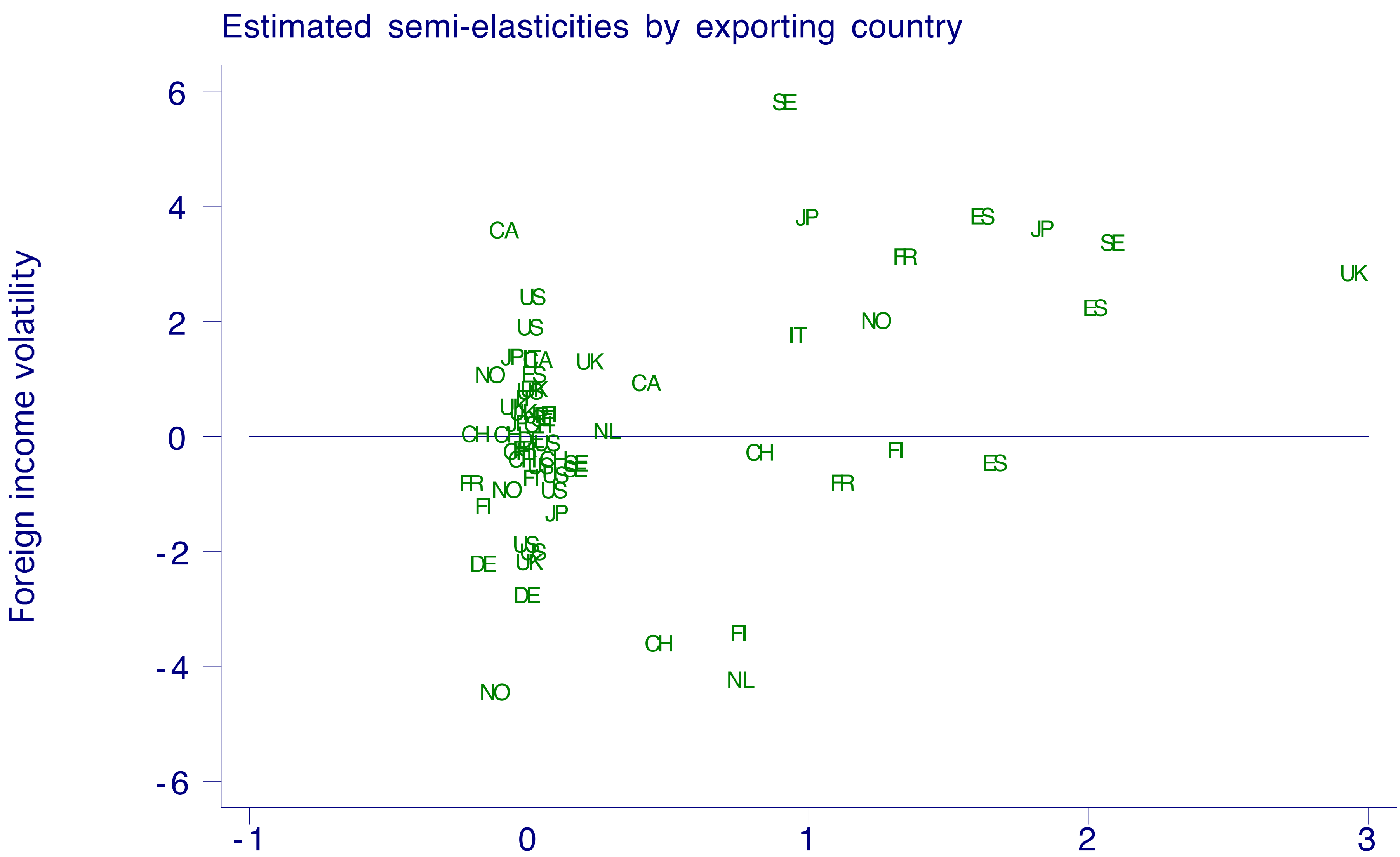
lag

Figure 5



Estimated semi-elasticities by exporting country
Foreign income volatility
Exchange rate volatility
6
4
2
0
-2
-4
-6
-1
0
1
2
3

Figure 6

Appendix

8a. Estimation Results for exports from US, UK, FR

	Γ_{ij}	std. err.	Λ_{ij}	std. err.
US->UK	0.038	0.023	1.293	0.824
US->FR	-0.009	0.035	-2.020	0.477
US->DE	0.005	0.034	0.640	0.199
US->IT	0.095	0.035	-0.810	0.280
US->NL	0.042	0.041	-0.655	0.309
US->NO	0.032	0.036	-0.089	0.172
US->SU	0.091	0.045	-1.081	0.995
US->CH	0.014	0.085	2.284	0.998
US->CA	0.064	0.032	-0.259	0.191
US->JP	-0.055	0.045	0.118	0.689
US->FI	0.016	0.055	-2.150	0.686
US->ES	0.005	0.036	1.759	0.697
UK->US	0.001	0.014	-2.327	0.769
UK->FR	-0.056	0.024	0.371	0.264
UK->DE	0.019	0.034	0.678	0.290
UK->IT	0.066	0.034	0.104	0.500
UK->NL	2.950	0.051	2.699	0.495
UK->NO	-0.020	0.020	0.280	0.082
UK->SU	0.219	0.048	1.159	0.419
UK->CH	-0.081	0.043	-0.539	0.377
UK->FI	0.025	0.094	0.766	0.794
UK->ES	0.025	0.038	-0.149	0.970
FR->US	1.345	0.043	2.990	2.534
FR->UK	0.000	0.023	-0.404	0.210
FR->DE	-0.117	0.085	0.343	0.289
FR->IT	0.030	0.029	-0.012	0.109
FR->NL	-0.204	0.108	-0.959	0.256
FR->NO	0.012	0.060	0.348	0.207
FR->SU	-0.016	0.036	-0.364	0.157
FR->CH	-0.458	1.041	-1.561	13.546
FR->CA	0.090	0.112	1.241	2.017
FR->JP	0.094	0.062	3.327	1.748
FR->FI	1.122	0.062	-0.951	0.905
FR->ES	0.120	0.074	-1.280	1.114

8b. Estimation Results for exports from DE, IT, NL

	Γ_{ij}	std. err.	Λ_{ij}	std. err.
DE->US	-0.005	0.032	-2.908	1.087
DE->UK	0.024	0.018	-0.158	0.129
DE->FR	-0.163	0.103	-2.363	1.023
DE->IT	0.025	0.018	-0.026	0.076
DE->NL	0.102	0.088	0.144	0.160
DE->NO	-0.012	0.026	-0.017	0.103
DE->SU	0.008	0.021	-0.198	0.101
DE->CH	0.060	0.079	-0.018	0.477
DE->CA	0.066	0.052	0.431	1.045
DE->JP	0.028	0.018	-0.141	0.303
DE->FI	-0.062	0.048	-0.084	0.523
DE->ES	0.002	0.022	0.318	0.396
IT->US	-0.001	0.015	-0.479	0.369
IT->UK	0.013	0.022	0.001	0.204
IT->FR	0.018	0.048	-0.766	1.329
IT->DE	-0.018	0.019	0.519	0.217
IT->NL	-0.037	0.024	-0.065	0.240
IT->NO	-0.014	0.020	0.083	0.064
IT->SU	0.014	0.021	0.010	0.067
IT->CA	0.058	0.031	-0.418	0.679
IT->JP	0.010	0.022	1.214	0.613
IT->FI	-0.060	0.033	0.346	0.257
IT->ES	0.960	0.042	1.622	1.947
NL->US	0.756	0.042	-4.378	2.596
NL->UK	0.023	0.023	-0.077	0.086
NL->FR	-0.001	0.096	-0.898	1.010
NL->DE	0.278	0.101	-0.050	0.203
NL->IT	-0.003	0.020	-0.072	0.084
NL->SU	-0.027	0.030	-0.284	0.145
NL->CH	0.123	0.087	0.055	0.794
NL->CA	0.085	0.082	0.418	3.213
NL->JP	0.034	0.023	-0.174	0.442
NL->FI	-0.092	0.070	-0.853	1.076
NL->ES	0.017	0.040	-1.757	1.029

8c. Estimation Results for exports from NO, SU, CH

	Γ_{ij}	std. err.	Λ_{ij}	std. err.
NO->US	-0.122	0.040	-4.598	2.415
NO->UK	0.013	0.015	-0.243	0.574
NO->FR	-0.139	0.051	0.925	1.115
NO->DE	-0.023	0.021	0.245	0.168
NO->IT	0.006	0.022	-0.010	0.113
NO->NL	-0.008	0.045	0.005	0.020
NO->SU	-0.079	0.043	-1.074	0.330
NO->CH	0.012	0.065	0.791	1.054
NO->CA	-0.002	0.083	-1.162	2.161
NO->JP	1.242	0.027	1.873	0.751
NO->FI	0.032	0.055	-0.483	0.333
NO->ES	-0.050	0.042	1.209	1.081
SU->US	0.914	0.046	5.680	2.907
SU->UK	0.029	0.038	0.382	0.518
SU->FR	2.088	0.057	3.228	1.649
SU->DE	0.028	0.031	0.152	0.151
SU->IT	0.171	0.060	-0.609	0.475
SU->NL	0.045	0.050	-0.072	0.267
SU->NO	0.053	0.029	0.178	0.074
SU->CH	0.033	0.057	-0.446	0.747
SU->CA	0.167	0.081	-0.708	1.487
SU->JP	-0.018	0.030	0.853	0.548
SU->FI	0.093	0.048	-0.125	0.765
SU->ES	-0.034	0.032	-0.910	0.508
CH->US	0.466	0.034	-3.742	2.104
CH->UK	-0.024	0.026	-0.544	0.206
CH->FR	0.826	0.060	-0.420	1.186
CH->DE	-0.114	0.060	0.003	0.072
CH->IT	-0.002	0.026	0.037	0.173
CH->NL	-0.002	0.092	-0.054	0.244
CH->NO	-0.075	0.032	-0.109	0.110
CH->SU	-0.042	0.026	-0.401	0.153
CH->CA	0.089	0.040	-0.545	0.498
CH->JP	0.034	0.017	0.061	0.451
CH->FI	-0.192	0.077	-0.099	0.868
CH->ES	-0.036	0.056	0.198	2.004

8d. Estimation Results for exports from CA, JP

	Γ_{ij}	std. err.	Λ_{ij}	std. err.
CA->US	0.419	0.044	0.789	1.313
CA->FR	0.031	0.053	1.193	0.396
CA->IT	-0.020	0.026	-0.230	0.219
CA->NL	-0.092	0.055	-0.499	0.653
CA->NO	-0.050	0.152	-0.510	0.698
CA->SU	0.062	0.057	-0.419	0.532
CA->CH	-0.089	0.125	3.448	1.405
CA->JP	-0.041	0.025	-0.337	0.484
CA->FI	0.013	0.115	0.436	2.331
CA->ES	-0.095	0.095	-1.725	1.412
JP->US	-0.057	0.015	1.238	0.396
JP->UK	0.000	0.014	-0.447	0.446
JP->FR	0.102	0.023	-1.475	0.565
JP->DE	1.836	0.015	3.473	0.357
JP->IT	0.009	0.016	-0.094	0.124
JP->NL	0.032	0.016	0.228	0.160
JP->NO	0.192	0.156	-2.128	1.474
JP->SU	0.013	0.020	-0.201	0.266
JP->CH	0.996	0.017	3.670	0.789
JP->CA	-0.039	0.018	0.090	0.339
JP->FI	0.035	0.032	0.614	0.466
JP->ES	0.031	0.022	-0.122	0.588

8e. Estimation Results for exports from FI, ES

	Γ_{ij}	std. err.	Λ_{ij}	std. err.
FI->US	0.749	0.075	-3.562	4.954
FI->UK	0.000	0.052	-0.500	0.955
FI->FR	1.311	0.058	-0.383	1.578
FI->DE	0.014	0.038	0.289	0.344
FI->IT	0.007	0.032	-0.864	0.240
FI->NL	-0.164	0.072	-1.353	0.338
FI->NO	0.071	0.059	0.247	0.125
FI->SU	0.028	0.041	0.116	0.210
FI->CA	0.036	0.066	-1.154	0.712
FI->JP	-0.071	0.074	2.448	1.596
FI->ES	0.007	0.041	1.009	0.679
ES->US	-0.016	0.024	0.765	0.600
ES->UK	1.622	0.139	3.691	5.943
ES->FR	-0.042	0.057	-2.394	1.452
ES->DE	2.025	0.047	2.096	0.585
ES->IT	0.057	0.050	0.565	0.368
ES->NL	-0.188	0.173	2.421	2.305
ES->NO	0.028	0.074	0.203	0.249
ES->SU	0.018	0.061	0.939	0.400
ES->CH	-0.098	0.066	0.331	0.675
ES->CA	-0.003	0.096	1.139	5.084
ES->JP	-0.036	0.045	-0.789	0.811
ES->FI	1.667	0.074	-0.603	1.392