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## **A New Look At The Trade Volume Effects of Real Exchange Rate Risk**

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## A New Look At The Trade Volume Effects of Real Exchange Rate Risk

### **Abstract:**

This paper takes a new empirical look at the longstanding question of the effect of exchange rate volatility on international trade flows by studying the case of Taiwan's exports to the United States from 1989-1998. In particular, we employ sectoral level, monthly data and a multivariate GARCH-M estimator with corrections for leptokurtic errors that is consistent with the core hypothesis that traders' forward contracting behavior might be affected by exchange rate risk. We find that real exchange rate risk has insignificant effects in most sectors, although agricultural trade volumes appear highly responsive to real exchange rate volatility. These results differ significantly from those obtained using more conventional and restrictive modeling assumptions.

## I. Introduction

One of the leading conundrums in international economics concerns the relationship between exchange rate volatility and international trade volumes. The widespread popular perception that greater exchange rate risk<sup>1</sup> reduces trade has helped motivate monetary unification in Europe (EU Commission 1990) and is strongly related to currency market intervention by central banks (Bayoumi and Eichengreen 1998). Yet a vast economic literature yields highly inconsistent theoretical and empirical results on this issue.

The conventional argument is that exchange rate uncertainty causes revenue uncertainty that will dampen trade due to risk aversion, irreversible investment in productive capital, or both (Ethier 1973, Demers 1991, Sercu 1992). DeGlaube (1988) nicely illustrates, however, how the relationship between exchange rate volatility and trade flows is analytically indeterminate. The issue has subsequently become a fundamentally empirical one.<sup>2</sup>

The empirical literature is no clearer on this question. A fair number of authors have indeed found evidence that exchange rate uncertainty may induce marginal producers and traders to shift from traded to nontraded goods, thereby dampening international trade volumes (Bahmani-Oskooee and Payesteh 1993, Caballero and Corbo 1989, Caporale and Doroodian 1994, Chowdhury 1993, Cushman 1983, Cushman 1986, Doroodian 1999, Kenen and Rodrik 1986, Kumar and Dhawan 1991, Pozo 1992, Thursby and Thursby 1987). On the other hand, a recent literature has produced evidence that seemingly supports the contrary hypothesis, that exchange rate volatility may stimulate trade (Dellas and Zilberfarb 1993, Franke 1992, Giovannini 1988, Sercu and Vanhulle 1992). Then, inevitably, many empirical studies have failed to establish any significant link between measured exchange rate variability and the volume of international trade (Assery and Peel 1991, Bahmani-Oskooee 1991, Bailey et al 1986, Bailey et al. 1987, Hooper and Kohlhagen 1978, IMF 1984, Gagnon 1993, Gotur 1985, Koray and Lastrapes 1989, Medhora 1990). The empirical evidence on this relationship is as ambiguous as the theoretical evidence.

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<sup>1</sup> We use the terms “risk”, “volatility”, “variability” and “uncertainty” interchangeably in this paper.

<sup>2</sup> McKenzie (1999) offers a detailed review of this literature.

McKenzie's (1999) excellent survey of the literature emphasizes a few key points in charting the empirical road ahead. These include (i) the need for care in specifying the technique by which exchange rate volatility is measured, with increased attention paid to rational expectations-based approaches using generalized autoregressive conditional heteroskedasticity (GARCH) and related methods, (ii) the importance of using data disaggregated by sector, market and time period, and (iii) necessary correction for prospective problems of serial correlation and nonstationarity in time series data. In this paper, we offer a new look at the exchange rate volatility-trade relationship that, for the first time to the best of our knowledge, attends to each of these three issues. We also attend to the problem of non-normality of regression residuals, an issue thus far unaddressed in this particular literature although it has become a source of much attention in financial econometrics. Furthermore, because we study exports from Taiwan, we contribute to a literature of particular salience to contemporary economic policy in low- and middle-income economies heavily dependent on international trade. While there have been other studies of the trade-exchange rate risk relationship using developing country data, this appears to be the first to explore the question using higher frequency (monthly) or sectorally disaggregated data.

The remainder of the paper is structured as follows. The next section briefly motivates our approach to specifying exchange rate volatility. Section III then discusses model specification and related econometric questions. Section IV presents and discusses our estimation results. Section V concludes.

## **II. Estimating Exchange Rate Volatility**

Two crucial issues underpin many of the empirical inconsistencies in this literature. The first is how an exporter<sup>3</sup> conceptualizes exchange rate risk and incorporates it into her trade decisions. We start with the maintained hypotheses that agents are concerned about the real exchange rate, not nominal rates,<sup>4</sup> and that they are forward looking and form expectations rationally. Several studies have demonstrated that the former assumption makes little difference in practice; nominal and real exchange rate

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<sup>3</sup> This could equally be an importer but since we use export data, we refer solely to exporters in the text.

<sup>4</sup> The real exchange rate (RX) is defined as  $E * (P_{\text{foreign}} / P_{\text{home}})$ , where E is the nominal exchange rate with respect to the foreign currency (for example, NT\$/US\$ in our data) and  $P_{\text{foreign}}$  and  $P_{\text{home}}$  represent appropriate price indices, in our case the foreign and Taiwan wholesale price indices, respectively.

series generate nearly identical empirical results (Qian and Varangis 1994, McKenzie and Brooks 1997, McKenzie 1999).

The issue that does seem to matter is the question of how agents form expectations over uncertain exchange rates and over what horizon. There is almost invariably a lag between the time at which an exporter contracts for sale and the time at which delivery, payment, or both are made. In many sectors, production decisions may precede contracting decisions by significant amounts, as in the case of agriculture, wherein farmers typically undertake land allocation and planting decisions well before they sell much, if any, of their crop forward. Most of the extant empirical literature uses realized exchange rate volatility, as proxied by measures such as the absolute percentage changes in the exchange rate, lagged standard deviations or moving average variance around trend. These measures either impose an assumption of adaptive expectations, wherein economic agents use only past information in predicting future exchange rate distributions, or suffer endogeneity, in the case where centered moving averages are used in spite of the fact that future exchange rate movements are surely affected in part by current trading behaviors. Measures that use realized values of exchange rate volatility suffer both conceptual and statistical problems of various sorts (Lanyi and Suss 1982). As we report in section IV, for the sample data studied in this paper, models based on estimates of agents' rational expectations of conditional mean and variance outperform those based on realized level and volatility statistics.

The period over which agents form expectations likewise matters. The literature generally assumes contemporaneous or one period lagged relationships between exchange rates and trade volumes. In part this is due to widespread use of quarterly or annual data, and it would seem reasonable to expect that contracts typically lock in nominal prices only out to a six month horizon or so. But when one uses higher frequency data, as we argue below is preferable, then it becomes less clear what the appropriate lag/lead structure ought to be for these relationships. Our approach is to let the data speak for themselves and to use established statistical methods to test for the optimal lag structure. We find a multiple period lag often yields the best fit to the data. The econometric literature more broadly supports the use of autoregressive moving average (ARMA) specifications as a convenient, reduced form method of capturing rational expectations processes of uncertain lag structure (Feige and Pearce 1976,

Nerlove, Grether, and Carvalho 1979, Wallis 1980). We therefore follow that tradition in this paper.

The second issue is how the econometrician proxies for the exchange rate uncertainty perceived by economic agents. This is closely related to the first issue, but distinct in that even if researchers agree on how agents conceptualize uncertainty and form expectations over exchange rate distributions, there is no generally accepted method for quantifying this risk (McKenzie 1999). Here we follow a burgeoning recent literature that relies on Bollerslev's (1986) generalized autoregressive conditional heteroscedasticity (GARCH) model to allow for time-varying conditional variance (i.e., volatility clustering) in exchange rate series (Pozo 1992, Kroner and Lastrapes 1993, Caporale and Doroodian 1994, Qian and Varangis 1994, McKenzie and Brooks 1997). Unlike most of this literature, however (with the notable exceptions of Kroner and Lastrapes 1993 and Caporale and Doroodian 1994), we estimate the exchange rate process simultaneously with the trade volume equation using a multivariate GARCH-in-mean estimator,<sup>5</sup> thereby avoiding the generated regressors problem that bedevils the rest of the literature that uses GARCH modeling in a two-step process to identify the conditional variance of the (real) exchange rate series (Pagan 1984, McKenzie 1999).

### **III. Model Specification**

In specifying an econometric model for estimating the trade-risk relationship, we take four further econometric issues into consideration: (i) potential intersectoral or temporal aggregation bias, (ii) appropriate lag specification for both the ARMA and distributed lag terms in the model, (iii) prospective time-varying correlation in the trade volume and exchange rate equations' regression errors, and (iv) potential non-normality in the regression errors. We tackle these in turn in introducing our estimation framework.

Most previous studies use data on trade flows aggregated across sectors and overseas markets and on exchange rates averaged over time. This necessarily imposes the strong, undesirable assumption that the impact of exchange rate volatility is uniform across sectors and destination markets. Klein (1990), Bini-Smaghi (1991) and McKenzie (1999) argue strongly for sectorally disaggregated estimation of the trade-risk



relationship and demonstrate that disaggregation uncovers significant intersectoral variation in the effect of exchange rate volatility on trade flows. For example, some sectors, such as agriculture, may be far more sensitive to exchange rate risk than other sectors (Maskus 1986, Pick 1990).

A related aggregation issue concerns the frequency of the data used in estimation. Due largely to data limitations, most studies employ quarterly or annual to examine the trade and risk relationship (McKenzie 1999). However, temporal aggregation necessarily dampens exchange rate variability, which may make identifying any true trade-risk relationship more difficult (Wang, Fawson and Barrett 2002). Moreover, where different sectors have different conventions for contracting and delivery or payment lags, intersectoral and intertemporal aggregation together could necessarily mute real trade-risk effects. For example, casual observation and discussion with industry participants suggest trade in services, electronics and transportation involve relatively short contracting lags as compared to trade in agricultural commodities, metals and intermediate industrial inputs commonly sold on long-term contracts.

For this study we were able to assemble monthly export data over ten years, 1989-1998, from Taiwan to its largest trading partner, the United States, for eight different productive sectors corresponding to the Standard Classification of Commodities (SCC) codes of the Republic of China: 1) animal and vegetable products and prepared foods; 2) textiles and textile articles; 3) wood, paper, pulp and articles; 4) chemicals, plastics, rubber and articles; 5) primary metals and articles; 6) optical and precision instruments; 7) electronic machinery and 8) transportation. We constructed matching export volume series for each sector as the ratio of export values reported in the *Monthly Statistics of Exports and Imports, Republic of China* tapes to the export price reported in the serial *Commodity-Price Statistics Monthly in Taiwan Area of the Republic of China*. Figures 1.1-1.8 display these series. The export volume of industries 2, 3 and 4 exhibit decline over time, while industries 5, 6 and 7 show clear upward trends from January 1989 to December 1998.

The literature pays relatively little attention to the dynamic specification of the trade-risk relationship. Most studies only consider the contemporaneous or lagged one period effect of the independent variables on the trade decisions without further

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<sup>5</sup> This builds on the seminal paper on ARCH-M estimation by Engle, Lilien, and Robins (1987).

investigating the possibility of any longer lagged response of the endogenous or exogenous variables. This seems an especially important issue when using higher frequency and sectorally disaggregated data, since one month lags may be suitable for some sectors where spot market transactions and rapid payments settlements are common, while longer lags may be more appropriate in other sectors characterized by significant forward contracting, payments delays, or both. If one wishes to reduce aggregation bias in estimation by using more temporally and cross-sectionally disaggregated data, it would seem all the more important to take care in specifying appropriate lag patterns as well. We therefore develop a model with a quite general lag structure, then painstakingly search for the optimal lag specification before estimating the system of equations.

We assume exporters form rational expectations of the real exchange rate series following an ARMA(m,n) process, with conditional variance specified as a GARCH(p,q) process, following equations (1)-(4):

$$\Phi_m(L) DLRX_t = \mu + \Phi_n(L) \varepsilon_{1,t} \quad (1)$$

$$\varepsilon_{1,t} | \varphi_{t-1} \sim N(0, h_t) \quad (2)$$

$$\varepsilon_{1,t} = z_t \sqrt{h_t} \quad (3)$$

$$E(\varepsilon_{1,t}^2 | \Psi_{t-1}) = h_t = w_0 + \sum_{i=1}^q \alpha_i \varepsilon_{1,t-i}^2 + \sum_{j=1}^p \beta_j h_{t-j} \quad (4)$$

DLRX<sub>t</sub> is the first difference in the natural logarithm of the real exchange rate with respect to the previous period, representing monthly percent change in the real exchange rate. It is essential in time series analysis of these relationships to test for stationarity since if trade flows are nonstationary, as is typically the case, yet exchange rate volatility is stationary, as is likewise common, then currency risk necessarily cannot determine trade volumes. We therefore tested for stationarity using the augmented Dickey-Fuller (ADL) test, results of which are available on request. The logarithm of the real exchange rate series was found to be integrated of order one, hence the first differencing used here. L represents a polynomial lag operator used to capture the ARMA properties of the conditional mean equation. We impose the nonnegativity restriction that  $w_0, \alpha_i, \beta_j \geq 0$  in order to ensure strictly positive conditional variance. The residuals from equation (1),

$\varepsilon_{1,t}$ , are a function of the independent and identically distributed  $z_t$ , which have zero mean and unit variance, and of the conditional variance,  $h_t$ , which depends on the information set,  $Q_{t-1}$ , available at time  $t-1$ .

Standard Box-Jenkins diagnostics, available on request, indicate that the NT/\$ real exchange rate follows a AR(1)-GARCH(1,1) process over our sample period. Combined with equations (1) and (4), this generates  $k$ -period ahead expectations (denoted by the superscript  $e$ ) of real exchange rate changes and risk of the form

$$DLRX_t^e = \gamma_0 \sum_{i=0}^{k-1} \gamma_1^i + \gamma_1^k DLRX_{t-k} \quad (5)$$

$$h_t^e = \omega_0 \sum_{i=0}^{k-1} \beta_1^i + \alpha_1 \beta_1^{k-1} \varepsilon_{1,t-k}^2 + \beta_1^k h_{t-k} \quad (6)$$

We use these expectations terms as regressors in our export volume equation. We accept the general consensus in the literature that there is a long run relationship between exports, the level of economic activity, real exchange rate and a measure of exchange rate risk (Pozo 1992, De Grauwe 1988, Kenen and Rodrik 1986, McKenzie 1999).<sup>6</sup> Assuming a linear first-order approximation to the true underlying relationship, we specify a reduced form model as

$$\begin{aligned} \ln(Q_{i,t}) = & \delta + \delta_1 \ln(IP_{t-k_1}) + \delta_2 \ln(RX_{i,t-k_2}^e) + \delta_3 \ln(h_{i,t-k_2}^e) + \sum_{k=1}^3 \delta_{4k} D_k \\ & + \sum_{k=1}^{k_3} \delta_{5k} \ln(Q_{i,t-k}) + \varepsilon_{2i,t} \end{aligned} \quad (7)$$

where  $Q_{i,t}$  is Taiwan's export volume (the export value deflated by the export price index by specific industry) of industry  $i$  to the United States during period  $t$ . Industrial production,  $IP_{t-k_1}$ , is used as the monthly proxy for the economic activity of U.S. in period  $t-k_1$ , where  $k_1$  is a lag to be estimated. We use IP because more conventional proxies for economic activity, such as income, are only available at quarterly frequency. In the export equation (7),  $RX_{i,t-k_2}^e$  is the real exchange rate predicted by equation (5); whereas

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<sup>6</sup> There is considerable variation in the literature as to the control variables used in the export equation. As McKenzie's (1999) survey points out, however, the variables we include appear to suffice as there is rarely any appreciable difference between the parameter estimates obtained using such a parsimonious specification and those arising from models that include a wider range of explanatory variables. Moreover, the higher frequency data we use render many other candidate series unavailable as regressors.

$h_{i,t-k}^e$  is the predicted real exchange rate volatility of equation (6), each with lags to be determined optimally by the sample data. We control for the seasonality readily apparent in the export plots (Figures 1.1-1.8) using quarterly dummy variables,  $D$ . Finally, lagged export volume ( $Q_{i,t-k}$ ) was included in the specification to reflect persistence in export volumes. The regression residual,  $\varepsilon_{2i,t}$ , has the usual Gauss-Markov properties. All the variables except  $D$  are in natural logarithm form, implying a constant elasticity structure.<sup>7</sup>

While the estimated conditional mean and variance of real exchange rate could be substituted into the export equation in a two-step estimation procedure, as several previous authors have done, this can lead to a generated regressors problem of biased estimates of the parameters' standard errors and potentially inconsistent parameter estimates (Murphy and Topel 1985, Pagan 1984, Pagan and Ullah 1988, McKenzie 1999). We obviate this problem by estimating the parameters of the conditional mean conditional variance real exchange rate equations simultaneously with the export volume equation by using full information maximum likelihood (FIML), which ensures both consistency and efficiency.

Specification of the FIML covariance matrix then becomes important. We allow for time-varying conditional variance for the real exchange rate series. We do, however, impose the assumption of time-invariant conditional variance on the export volume series because preliminary statistical analysis, available on request, revealed that the variances of each sector's export volume series in our sample proved time invariant. This finding is consistent with previous studies (Meng-Chun Tsai 2000, Kroner and Lastrapes 1993).

Real exchange rates and international trade comove in general equilibrium. We therefore allow for time-varying covariance among the two regressions' error terms, obviating the potential inefficiency that comes from ignoring the time varying covariance terms (Holt and Aradhyula, 1998). Although the variance of export volume does not vary across periods, the covariance between export volume and the real exchange rate likely does vary since the conditional variance of the latter series is clearly time-varying. We therefore specify a covariance matrix for the FIML model that includes a constant variance for export volume,  $\sigma_{22}$ , but allows for time-varying conditional variance of the real exchange rate returns following the GARCH process and, hence, time-varying

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<sup>7</sup> The industrial production, nominal exchange rate, and wholesale price index series come from the International Monetary Fund Economic Information System (IMFEIS) and Taiwan AREMOS system.

covariance ( $\sigma_{12} = \sigma_{21}$ ) between export volumes and the real exchange rate. We assume the conditional correlation between the two variances,  $\rho \in [-1,1]$ , is constant through time.<sup>8</sup> The time varying covariance is proportional to the square root of the product of the two conditional variances,  $h_t$  and  $\sigma_{22}$ . Under standard regularity conditions, the error terms  $\varepsilon_{1,t}$  and  $\varepsilon_{2,t}$  are distributed multivariate normal distribution with zero mean and the time-varying variance-covariance matrix  $H_t$ . The system could be described as:

$$\hat{\varepsilon}_t = \begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix} \quad (8)$$

$$\hat{\varepsilon}_t | \mathcal{Q}_{t-1} \sim N(0, H_t) \quad (9)$$

$$H_t = \begin{bmatrix} h_t & \sigma_{12,t} \\ \sigma_{21,t} & \sigma_{22} \end{bmatrix} \quad (10)$$

$$\sigma_{12,t} = \sigma_{21,t} = \rho \sqrt{(h_t \sigma_{22})} \quad (11)$$

Our model thus involves simultaneous nonlinear estimation of equations (1)-(11). We used the Berndt, Hall, Hall, and Hausman (BHHH) algorithm in the Gauss Constrained Maximum Likelihood (CML) module. Let  $\theta$  denote the unknown parameters in  $\hat{\varepsilon}_t$  and  $H_t$ . The log-likelihood function under general heteroskedasticity with a multivariate normal distribution then becomes

$$L(\theta) = -\sum_{t=1}^n (\ln |H_t| + \hat{\varepsilon}_t' H_t^{-1} \hat{\varepsilon}_t) \quad (12).$$

However, conventional estimation methods in this literature tend to understate the effects of exchange rate variability on trade volumes due to the nonnormal properties of exchange rate changes (Arize, 1997).<sup>9</sup> Pagan and Sabau (1987) demonstrate that both efficiency and, in the case of maximum likelihood estimation, consistency of parameter estimates require correct specification of that conditional distribution. We therefore test

<sup>8</sup> The constant conditional correlation assumption simplifies computation and inference. Moreover, it has been proved reasonable in many previous applications (Bollerslev 1990, Theodossiou and Lee 1993).

<sup>9</sup> Exchange rate change distributions typically exhibit leptokurtosis (heavy tails), as shown by Westerfield (1977), Engle and Bollerslev (1986), Milhøj (1987), Baillie and Bollerslev (1989), Hsieh (1989), Baillie and DeGennaro (1990) and Wang et al. (2001).

explicitly for non-normality and, where appropriate, relax the usual multivariate normal distribution assumption to accommodate greater leptokurtosis using a multivariate Student-t distribution. With this assumption, the marginal distribution of each term is univariate Student t, including the Cauchy and normal distribution as special cases. We find that the substitution of a conditional heavy-tailed multivariate Student-t distribution for the conditional multivariate normal distribution helps improve the estimation performance when the data exhibits leptokurtosis. The likelihood function of the k-variate Student-t distribution with unknown  $\nu$  degrees of freedom and n observations is given by

$$L(\theta) = \ln \Gamma\left(\frac{\nu+k}{2}\right) - \ln \Gamma\left(\frac{\nu}{2}\right) - \frac{1}{2} \ln((\nu-2)\pi) - \frac{1}{2} \sum_{t=1}^n \left( \ln |H_t| + (\nu+k) \ln \left(1 + \frac{\hat{\varepsilon}_t' H_t^{-1} \hat{\varepsilon}_t}{\nu-2}\right) \right) \quad (13)$$

Where  $\Gamma$  denotes the gamma function.

We began estimation by identifying and estimating a common ARMA(m,n) process for the DLRX series following a three-step procedure proposed by Wang et al. (2001). First, Box-Jenkins iterative techniques are used to reduce the set of prospective ARMA specifications. Next, we further screen among the resulting candidates ARMA specifications to eliminate those having a p-value for the Ljung-Box portmanteau Q(12) statistic less than 0.3, a significance level clearly supporting the assumption of white noise. Finally, from among the candidate models having passed the Box-Jenkins and Q(12) screens we chose the optimal conditional mean specification based on the Schwarz Bayesian criterion (SBC).

As asserted earlier, this procedure established that an AR(1) model best represents the conditional mean of the DLRX series, equation (1).<sup>10</sup> The squared residuals from the AR(1) process were then found to exhibit serial correlation (Q=29.14, p-value=0.004), indicating a need to accommodate time varying conditional variance. We then tested a variety of symmetric GARCH and asymmetric GJR GARCH specifications. There proved to be no statistically significant asymmetry in the conditional variance process nor between the GJR GARCH and standard, symmetric GARCH(1,1) model, so we use the latter, more parsimonious specification for equation (4)<sup>11</sup>.

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<sup>10</sup> A table of estimated parameters and diagnostic checks on these preliminary specification checks is available by request. The Ljung-Box Q-statistic of residual from AR(1) process proves insignificant at all time horizons and reasonable levels of significance, signaling the absence of residual serial correlation.

<sup>11</sup> Diagnostic statistics and estimation results for these specification checks are likewise available by request. Of particular note, we could not reject the null hypothesis of normality for the residuals using a

Having determined the optimal specification of equations (1)-(4), we next determined the optimal lag structure for equations (5)-(7). The predicted exchange rates and exchange rate volatility generated from equation (5) and (6) were allowed to range from one to six months ahead for each industry. Expected exchange rate change and exchange rate risk were jointly estimated with the sectoral export equations sequentially for different predicted lag structures.<sup>12</sup> We selected the optimal lag structure based on the minimal log likelihood value. This yielded one month lag specifications in sector 5 (primary metals and articles), two month lag specifications in sectors 3 (wood, paper, pulp and articles), 6 (optical and precision instruments) and 8 (transportation), three month lag specifications in sectors 1 (animal and vegetable products and prepared foods), 2 (textiles and textile articles), 4 (chemicals, plastics, rubber and articles), and a four month lag in sector 7 (electronic machinery).<sup>13</sup> With the ARMA, GARCH and lag structures for equations (1)-(7) established, we could begin estimation of the FIML model.

#### **IV. Estimation Results**

Table 1 reports the estimated parameters of the multivariate GARCH-M model estimated using the FIML methods described above. As one would expect, the parameter estimates show considerable variability across sectors. Exports are significantly increasing in the condition mean of the real exchange rate series (RX) in six of the eight sectors, with sectoral elasticity estimates ranging from 0.85-15.55, and are negative but statistically insignificantly different from zero in the other two sectors. With the exception of anomalous negative and statistically significant estimates for sectors 2 and 3, export volumes are likewise significantly increasing in United States' industrial production, with parameter estimates in the 0.37-1.27 range. As Figures 1.2 and 1.3 show, export volumes consistently declined in these two anomalous sectors over the

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Jarque-Bera test nor the null hypothesis of no serial correlation in the squared residuals using the Ljung-Box Q statistic from the AR(1)-GARCH(1,1) process for the DLRX series.

<sup>12</sup> For each sector, the export volume, industrial production, real exchange rate and its conditional variance were all tested for stationarity and found to be integrated of order 1. We therefore used Johansen's multivariate cointegration method to check the number of cointegrating vectors for the nonstationary time series. Detailed results are omitted for the sake of brevity, but we found at least one cointegrating vector for each sectoral export volume-exchange rate system, clearly suggesting the existence of long run equilibrium relationships among the export volume, foreign income, real exchange rate, and exchange rate volatility. Thus, the spurious regression problem associated with nonstationary data does not affect our estimation.

<sup>13</sup> A table of results of these tests is available by request.

sample period as increasing labor costs, rising land prices, and stricter environmental protection laws forced many Taiwanese textile and wood, paper and pulp firms to shut down or relocate abroad, mainly to mainland China and Southeast Asia. The negative estimated coefficient on United States industrial production thus most likely reflects induced structural change for which the current specification does not control satisfactorily. There are strong seasonality effects evident in each sector's D1-D3 parameter estimates. Sectors 3-6 likewise exhibit significant autoregression.

Of primary interest to us, the estimated effects of expected exchange rate volatility on trade prove statistically insignificantly different from zero in seven of eight sectors. There are several likely reasons why exchange rate risk seems to have little effect on Taiwanese exports. First, the U.S. dollar is the major currency for Taiwan's international transactions. Second, forward currency markets are well developed and widely used by Taiwanese exporters to eliminate exchange rate risk. Third, longstanding business relations between many American and Taiwanese trading partners include arrangements to help eliminate exchange rate risk, such as open account agreements, especially for intra-firm trade between divisions of multinational firms, a widespread phenomenon in Taiwan, especially with respect to transportation and high technology industries. Finally, Taiwan's central bank holds unusually large foreign exchange reserves – the second largest in the world on average over the sample period, behind only Japan – and it routinely uses its reserves to stabilize the exchange rate. Taiwanese exporters therefore likely expect the central bank to be able and willing to intervene in currency markets if fluctuations become excessive, effectively providing exporters with insurance and perhaps making it somewhat easier for firms to predict exchange rate movements.

However, exports from the agricultural sector, where the above characteristics apply least, appear to respond negatively and significantly to expected exchange rate volatility. This is consistent with both an extant literature which argues that the agricultural sector is most susceptible to exchange rate uncertainty (Maskus 1986, Pick 1990) and to empirical evidence that Pacific Basin agricultural markets of importance to Taiwan are highly competitive in terms of price (Barrett et al. 1999, Barrett and Li 2002).

The exchange rate series indeed exhibit significant GARCH effects, as reflected in the coefficient estimates on  $h_{t-1}$  in the conditional variance equation. The estimated



variance ( $\sigma_{22}$ ) of natural resource-dependent exports in sectors 1 (agriculture), 2 (textiles) and 3 (wood, paper and pulp) were substantially higher than those of the other five sectors, with the estimated variance of exports from sectors 2 and 3 roughly double those of the other sectors and the variance of agricultural exports about five times that of the higher technology manufactured goods and services sectors. Most of the estimated cross-equation correlation parameters ( $\rho$ ) were statistically insignificantly different from zero, with the agricultural sector once again a notable exception.

Table 2 reports a battery of diagnostic test statistics from these regressions. The results generally confirm the satisfactory specification of each sectoral multivariate GARCH-in-mean model, as reflected in goodness of fit, tests for serial correlation in the residuals and squared residuals, for residual heteroskedasticity or GARCH effects, and for normality and structural stability.<sup>14</sup> The only significant failure of the multivariate normal GARCH-in-mean model seems to be the evident non-normality of the residuals in the model for sector 1, agriculture, where the risk-trade effect was most pronounced. Since non-normality corrupts inference with respect to this parameter estimate of primary interest, we reestimated the model now using the multivariate Student-t distribution for the error term. Table 2 indicates that excess kurtosis was the main source of nonnormality and the Student-t distribution directly accommodates such leptokurtosis. The parameter estimates under the multivariate Student-t GARCH-M model for sector 1 are presented in the leftmost column of Table 1, as sector 1\*. The primary effects of this alteration in the regression model are that the estimated effect of United States industrial production on Taiwanese agricultural exports increases from 0.54 to 0.72, albeit with considerable loss in precision, such that the latter estimate is now statistically insignificantly different from zero at conventional significance levels, and a sharp reduction in the estimated agricultural export volume dampening effect of expected exchange rate volatility. Where the multivariate normal model estimates an elasticity of agricultural exports with respect to expected exchange rate volatility of -35.3, controlling for apparent leptokurtosis drops that point estimate to -11.6. Exchange rate volatility still seems to exert significant negative effect on agricultural exports – and not on exports from any other sectors in Taiwan’s economy – but those trade-dampening effects are

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<sup>14</sup>Although the Chow test indicates that the estimated parameters of the sector 5 model are not constant over the full range of the data (we used a breakpoint in the middle of the sample), subsample-specific parameter estimates yielded qualitatively identical estimates, in particular that exchange rate risk has no significant effect on metals exports.

plainly exaggerated by leptokurtosis associated with misspecification of the multivariate error term distribution.

The likelihood ratio test statistic for the multivariate Student-t distribution against multivariate normal distribution is 3.44, suggesting that the accommodation of leptokurtosis indeed yields modest but statistically significant gains in model performance. Recall that the exchange rate data used here do not exhibit leptokurtosis, so these effects are almost surely dampened in this sample. Since leptokurtosis has been commonly observed in exchange rate data, the multivariate Student-t GARCH-M model could likewise offer significant improvements in other samples.

## **V. Conclusions**

This paper investigates the impact of exchange rate volatility on Taiwan's exports by estimating a multivariate GARCH-M model using sector- and market-specific monthly data. By using more disaggregated data and attending to a variety of econometric issues that bedevil much of the extant literature on this high profile issue, we offer a new look at this longstanding question. The most striking finding is that we cannot find any significant relationship between expected exchange rate volatility and trade volumes outside of the agricultural sector, where export variability is great and exchange rate volatility has a strong dampening effect. Even in the agricultural sector, however, our results show that failure to attend to issues of non-normality in the regression residuals seems to lead to substantial overstatement of the negative effect of exchange rate risk on trade flows.

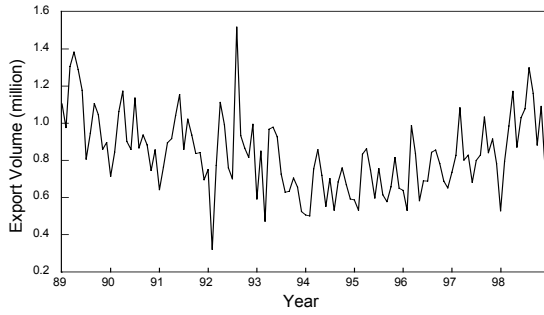
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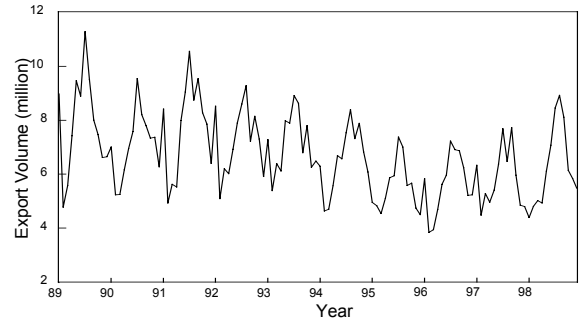
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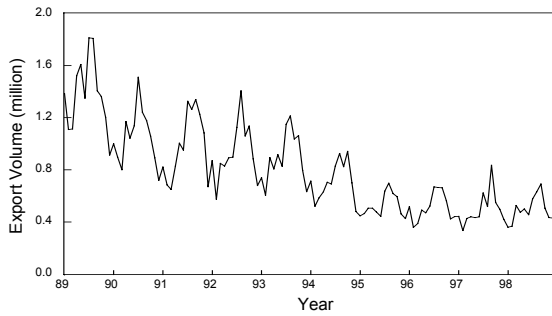
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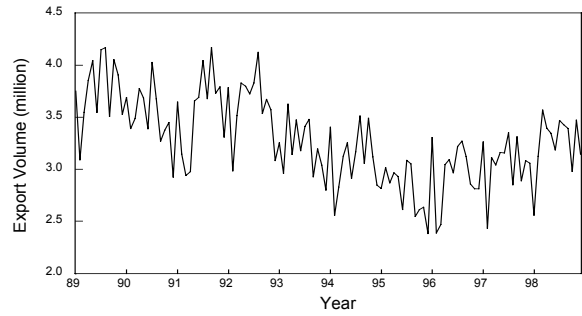
**Figure 1.1** Animal and vegetable products and prepared foods



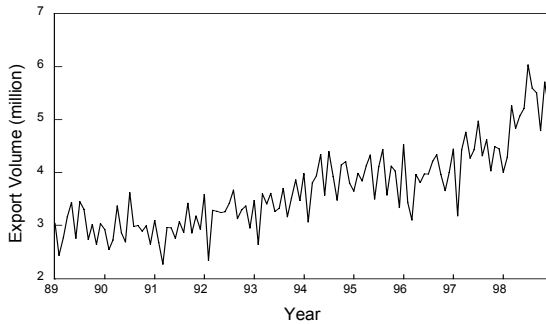
**Figure 1.2** Textiles & Textile Articles



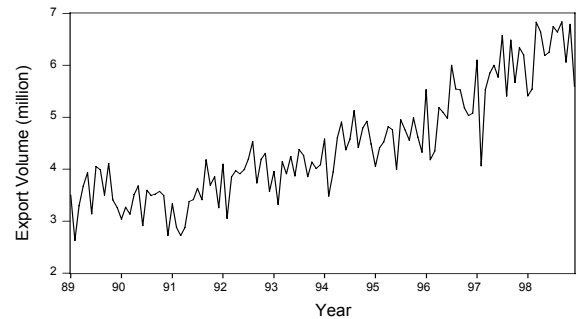
**Figure 1.3** Wood, Paper, Pulp & Articles



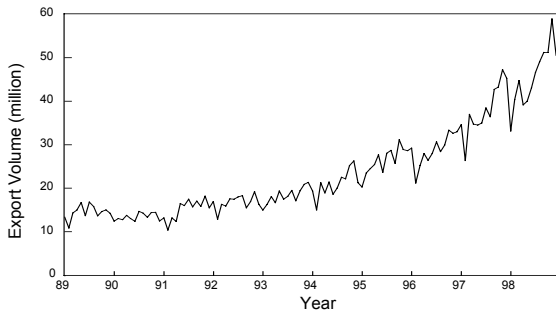
**Figure 1.4** Chemicals, Plastics, Rubber & Articles



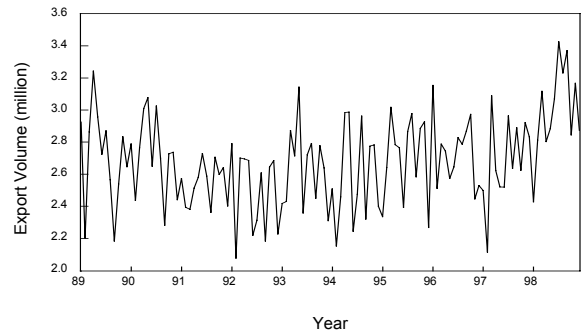
**Figure 1.5** Primary Metals & Articles



**Figure 1.6** Optical & Precision Instruments



**Figure 1.7** Electronic Machinery



**Figure 1.8** Transportation

Table 1: FIML Estimates of Sector-Specific Multivariate GARCH-M Models

Variables/Lags	Sector Number								
	1*	1	2	3	4	5	6	7	8
<b>Export Supply Equation Parameters</b>									
C	-16.05 (19.43)	6.91*** (1.03)	-44.65*** (3.62)	-23.39 (115.54)	1.00 (1.47)	-5.14*** (1.13)	-6.42*** (1.72)	-31.14 (98.84)	-11.84*** (5.34)
IP <sub>t</sub> (-1)	0.72 (0.76)	0.54** (0.23)							
IP <sub>t</sub> (-2)					0.37 (0.36)	0.65*** (0.27)	0.96*** (0.33)	1.27*** (0.58)	0.87*** (0.29)
IP <sub>t</sub> (-3)			-0.99*** (0.37)	-4.02*** (0.60)					
RX <sub>t</sub> <sup>c</sup> (-1)						0.85*** (0.36)			
RX <sub>t</sub> <sup>c</sup> (-2)				4.92*** (1.76)			0.90* (0.53)		-0.01 0.04
RX <sub>t</sub> <sup>c</sup> (-3)	6.61** (3.03)	5.44*** (0.41)	15.55*** (1.22)		-0.72 (0.70)				
RX <sub>t</sub> <sup>c</sup> (-4)								3.19*** (1.13)	
h <sub>t</sub> <sup>c</sup> (-1)						-0.01 (0.04)			
h <sub>t</sub> <sup>c</sup> (-2)				33.32 (147.60)			-0.06 (0.07)		2.74 (1.61)
h <sub>t</sub> <sup>c</sup> (-3)	-11.64*** (1.48)	-35.3*** (2.25)	-0.001 (0.003)		0.02 (0.06)				
h <sub>t</sub> <sup>c</sup> (-4)								23.2 (125.59)	
D1	-0.07 (0.06)	-0.09 (0.06)	-0.18 (0.04)	-0.09*** (0.05)	0.04 (0.03)	0.00 (0.03)	-0.01 (0.02)	-0.10*** (0.04)	0.01 (0.03)
D2	0.15** (0.06)	0.14*** (0.05)	-0.01 (0.03)	0.09* (0.06)	0.10*** (0.03)	0.13*** (0.03)	0.08*** (0.03)	-0.05 (0.04)	0.05** (0.03)
D3	0.08* (0.05)	0.08* (0.05)	0.23*** (0.03)	0.31*** (0.04)	0.10*** (0.03)	0.11*** (0.02)	0.10*** (0.02)	0.004 (0.03)	0.04 (0.03)
Q(-1)						-0.31*** (0.09)	-0.19** (0.09)		
Q(-2)				0.44*** (0.09)	0.31*** (0.10)		0.19** (0.08)		
Q(-3)					0.37*** (0.08)	0.38*** (0.08)	0.21** (0.09)		
Q(-4)						0.27*** (0.08)			
<b>Exchange Rate Equation Parameters</b>									
<b>Conditional Mean</b>									
C*10 <sup>2</sup>	0.11 (0.08)	0.12*** (0.01)	-0.007 (0.005)	0.16** (0.08)	0.15** (0.08)	0.14** (0.07)	0.15** (0.07)	0.18*** (0.06)	-0.06** (0.03)
AR(1)	0.40** (0.08)	0.41*** (0.002)	0.43*** (0.03)	0.38*** (0.07)	0.37*** (0.08)	0.35*** (0.09)	0.37*** (0.08)	0.37*** (0.06)	0.40*** (0.08)
<b>Conditional Variance</b>									
C*10 <sup>4</sup>	0.01 (.)	0.01 (.)	0.22*** (0.06)	0.03*** (0.01)	0.20 (0.34)	0.18* (0.10)	0.22 (0.21)	0.12 (0.10)	0.27 (0.28)
$\alpha$	0.00 (.)	0.00 (.)	0.39*** (0.08)	0.00 (.)	0.36** (0.17)	0.15** (0.08)	0.20 (0.25)	0.00 (0.00)	0.39* (0.22)
$\beta$	0.99*** (0.004)	0.99*** (0.003)	0.61*** (0.08)	0.96*** (0.02)	0.63*** (0.57)	0.61*** (0.16)	0.62** (0.33)	0.83*** (0.13)	0.53*** (0.14)
<b>Variance-Covariance</b>									
$\sigma_{22}$	352.24*** (47.06)	368.52*** (68.89)	173.71*** (3.31)	137.81*** (20.45)	70.38*** (10.84)	70.78*** (8.50)	67.89*** (8.96)	79.02*** (12.15)	77.62*** (12.09)
$\rho$	0.11 (0.12)	0.13*** (0.03)	0.03 (0.09)	0.12 (0.12)	-0.05 (0.13)	-0.19** (0.09)	0.03 (0.10)	-0.04 (0.10)	0.02 (0.10)

Note: Sector 1\* reflects estimates from the rational expected multivariate GARCH-M model based on a multivariate Student-t distribution, while sectors 1-8 use a multivariate normal distribution.



Table 2: Diagnostic Test Statistics on Sector-Specific Export Volume Equations

	Sector Number							
	1	2	3	4	5	6	7	8
R <sup>2</sup>	0.63	0.64	0.91	0.57	0.82	0.87	0.96	0.56
Q(12)	8.23 [0.22]	5.38 [0.50]	14.17 [0.12]	6.65 [0.35]	5.18 [0.52]	6.49 [0.26]	5.03 [0.54]	6.61 [0.36]
Q <sup>2</sup> (12)	2.17 [0.83]	8.88 [0.18]	10.41 [0.11]	6.12 [0.41]	5.28 [0.51]	5.29 [0.51]	3.64 [0.73]	1.19 [0.98]
B-G	1.36 [0.24]	0.93 [0.63]	2.78 [0.25]	1.43 [0.49]	0.12 [0.73]	0.42 [0.52]	4.01 [0.13]	0.77 [0.38]
ARCH	0.49 [0.48]	0.95 [0.39]	0.1 [0.76]	0.12 [0.73]	0.70 [0.41]	1.31 [0.27]	0.003 [0.95]	0.50 [0.48]
Skewness	-0.07 (0.24)	0.21 (0.24)	0.12 (0.24)	0.16 (0.24)	-0.10 (0.24)	-0.12 (0.24)	-0.21 (0.24)	-0.20 (0.24)
Kurtosis	5.11 (0.47)	2.70 (0.47)	3.14 (0.47)	2.76 (0.47)	2.66 (0.47)	3.22 (0.47)	2.80 (0.47)	2.49 (0.47)
J-B	22.04 [0.00]	1.31 [0.52]	1.84 [0.40]	0.78 [0.68]	0.73 [0.69]	0.51 [0.77]	1.09 [0.58]	2.07 [0.35]
Chow	0.81 [0.78]	0.74 [0.87]	0.98 [0.53]	1.10 [0.37]	3.37 [0.00]	0.71 [0.90]	1.00 [0.50]	1.38 [0.12]

(1) Figures in bracket are p-values

(2) Q and Q<sup>2</sup> indicate the Liung-Box portmanteau test statistics on the residuals and square residuals

(3) B-G is the Breusch – Godfrey serial correlation test statistic

(4) ARCH is the White test statistic for autoregressive conditional heteroscedasticity

(5) J-B is the Jarque-Bera normality test statistic for skewness and excess kurtosis

(6) Chow is the Chow stability test statistic