Pacific Economic Review, 9: 2 (2004) pp. 117–129

EXCHANGE RATE RISK AND EXPORT REVENUE IN TAIWAN

WEN SHWO FANG *Feng Chia University, Taiwan*
HENRY THOMPSON* *Auburn University, USA*

Abstract. The effect of exchange rate risk on export revenue in Taiwan between 1979 and 2001 is investigated in a bivariate GARCH-M model that simultaneously estimates time-varying risk. Depreciation is found to stimulate export revenue in domestic currency, but the quantitative impact is small and any associated increase in exchange risk has a negative impact. Implications for economic policy are discussed.

1. INTRODUCTION

The effects of exchange rate risk have been studied since the collapse of fixed exchange rates in the 1970s, but little consensus regarding its effect on export revenue has emerged. Exchange risk could lower export revenue owing to profit risk, as developed by Ethier (1973); but De Grauwe (1988) suggests that exporters might increase volume to offset potential losses, and Broll and Eckwert (1999) note that the price of an option to export increases with risk. The risk profile of exporting firms and currency inventory practices would certainly be relevant. Depreciation might increase export revenue, but the net effect could be negative if there were increased exchange risk. Policy-makers might be advised to remember exchange risk when considering market intervention aimed at stimulating exports.

Pozo (1992) uncovers a negative effect on UK real exports to the United States. Chowdhury (1993) and Arize (1995, 1996a, 1997) find negative impacts of exchange risk on US, European and G7 exports. Weliwita *et al.* (1999) find that Sri Lanka's exports to six developed countries fall with risk. Arize *et al.* (2000) use a moving sample standard deviation and find that risk has a negative effect in 13 LDCs.

There is, however, contrary evidence. Asseery and Peel (1991) find evidence of positive relationships for multilateral exports except for the United Kingdom. Kroner and Lastrapes (1993) find positive effects of conditional variance on exports of France, Germany and Japan, but negative effects for the United Kingdom and the United States. McKenzie and Brooks (1997) uncover positive risk relationships for Germany and the United States.

Generalized autoregressive conditional heteroskedasticity (GARCH) models have been used to model relationships between means and variances as in

**Address for correspondence*: Economics, Comer Hall, Auburn University, AL 36849, USA; Email thomph1@auburn.edu. We would like to thank two anonymous referees for their comments and suggestions. All remaining errors are, of course, our own.

Bollerslev (1986, 1990), Engle *et al.* (1987) and Bollerslev *et al.* (1992). A moving standard deviation of the exchange rate is used as a proxy for risk by various authors including Chowdhury (1993), Arize (1995, 1996a, 1996b, 1997), and Arize *et al.* (2000). Moving standard deviations, however, have a maintained hypothesis of homoskedasticity while being used to construct a proxy for heteroskedasticity.

Exchange risk has been shown to be conditional and time varying by Hodrick and Srivastava (1984). GARCH methodology allows time dependence in Pozo (1992), McKenzie and Brooks (1997) and Weliwita *et al.* (1999), but this two-step procedure may result in inefficient estimation. The present paper applies a bivariate GARCH-M model with simultaneous estimation of time-varying risk including volatility and depreciation as explanatory variables. The effects of the exchange rate and risk depend on how quickly exporters respond, and dynamic features of the present model distinguish it from contemporaneous multivariate GARCH-M models such as Kroner and Lastrapes (1993).

There is motivation to examine Taiwan, where fixed exchange rates were abandoned during 1978, and the present paper examines the evidence from 1979 to 2001. Most previous studies have focused on developed countries, but Taiwan industrialized during this period. Taiwan is a small open economy, with export revenue averaging 45% of GDP over this period. Darrat *et al.* (2000) credit the rapid growth of Taiwan to export promotion, but Chang *et al.* (2000) raise questions about this conclusion. The Asian financial crisis of 1997 caused dramatic depreciation and exchange rate volatility, and the flattening of export growth suggests a negative impact for exchange risk.

2. AN EMPIRICAL MODEL OF EXPORT REVENUE AND EXCHANGE RISK

Real export revenue (x) is specified as a function of real foreign income (y), the real effective exchange rate (q) and effective exchange risk (h_q) as

$$x = f(y, q, h_q) \tag{1}$$

Foreign income would have a positive effect on the demand for normal exports. The exchange rate is defined as the domestic currency price of foreign currency, and the effective exchange rate q is the export weighted average of real exchange rates across trading partners. An increase in the exchange rate is depreciation, implying cheaper exports abroad and increased real export revenue given the Marshall–Lerner condition. The effect of exchange risk is theoretically ambiguous.

To capture the dynamics, the export equation is modeled as an autoregressive distributed lag (ADL) process:

$$\Delta lx_t = a_0 + \sum_{i=1}^{n} a_i \Delta lx_{t-i} + \sum_{i=0}^{n} b_i \Delta ly_{t-i} + \sum_{i=0}^{n} c_i \Delta lq_{t-i} + \sum_{i=0}^{n} d_i \hat{h}_{q,t-i} + \varepsilon_{x,t} \tag{2}$$

where $\Delta lx_t = 100 \times (\ln x_t - \ln x_{t-1})$, $\Delta ly_t = 100 \times (\ln y_t - \ln y_{t-1})$, $\Delta lq_t = 100 \times (\ln q_t - \ln q_{t-1})$, $\hat{h}_{q,t}$ is estimated exchange risk, and $\varepsilon_{x,t}$ is a random disturbance term.

Risk is specified as time-varying exchange rate volatility constructed with a GARCH(1,1) process following Bollerslev (1986):

$$\Delta lq_t = s_0 + \sum_{i=1}^{P} s_i \Delta lq_{t-i} + \varepsilon_{q,t} \tag{3}$$

$$\varepsilon_{q,t} | \Psi_{t-1} \sim N(0,\ h_{q,t}) \tag{4}$$

$$h_{q,t} = \beta_0 + \beta_1 \varepsilon^2_{q,t-1} + \beta_2 h_{q,t-1}, \tag{5}$$

where the variance $h_{q,t}$ of the error term $\varepsilon_{q,t}$ is conditional on the information set Ψ_{t-1} available at time $t - 1$. The exchange rate is specified as an AR(p) process to capture serial dependence, and $\varepsilon_{q,t}$ should be white noise. Conditions β_0, β_1, $\beta_2 \geq 0$ and $\beta_1 + \beta_2 < 1$ are required to ensure positive finite variance and model stability. The contribution of the GARCH model is to allow the variance to vary. If β_1, $\beta_2 \neq 0$ the variance is time varying, and if $\beta_2 = 0$ the model reduces to an ARCH(1) model as developed by Engle (1982). Variation of the exchange rate is taken to be the conditional variance of $h_{q,t}$, a larger $h_{q,t}$ indicating more risk.

Equations (2)–(5) constitute two-step estimation with $\hat{h}_{q,t}$ generated by (3)–(5) and then used in (2) to estimate its effect on export revenue. A bivariate GARCH(1,1)-M model is used for efficiency in joint estimation:

$$\Delta lx_t = a_0 + \sum_{i=1}^{n} a_i \Delta lx_{t-i} + \sum_{i=0}^{n} b_i \Delta ly_{t-i} + \sum_{i=0}^{n} c_i \Delta lq_{t-i} + \sum_{i=0}^{n} d_i h_{q,t-i} + \varepsilon_{x,t} \tag{6}$$

$$\Delta lq_t = s_0 + \sum_{i=1}^{P} s_i \Delta lq_{t-i} + \varepsilon_{q,t} \tag{7}$$

$$\varepsilon_t = (\varepsilon_{x,t}, \varepsilon_{q,t}) \varepsilon_t | \Psi_{t-1} \sim N(0, H_t) \quad H_t = \begin{bmatrix} h_{x,t} & h_{xq,t} \\ h_{xq,t} & h_{q,t} \end{bmatrix} \tag{8}$$

$$h_{x,t} = \alpha_0 + \alpha_1 \varepsilon^2_{x,t-1} + \alpha_2 h_{x,t-1} \tag{9}$$

$$h_{q,t} = \beta_0 + \beta_1 \varepsilon^2_{q,t-1} + \beta_2 h_{q,t-1} \tag{10}$$

$$h_{xq,t} = \gamma_0 + \gamma_1 \varepsilon_{x,t-1} \varepsilon_{q,t-1} + \gamma_2 h_{xq,t-1} \tag{11}$$

where $h_{x,t}$ is the conditional variance of export revenue, H_t is the conditional covariance matrix, $h_{xq,t}$ is the conditional covariance and $\varepsilon_{x,t}$ and $\varepsilon_{q,t}$ are white noise stochastic processes with $(\varepsilon_{x,t}, \varepsilon_{q,t})$ distributed bivariate normal.

In the GARCH model conditional variances and covariances vary with time. Each element of the covariance matrix follows a univariate GARCH model driven by the corresponding element of the cross-product matrix $\varepsilon_{x,t-1}, \varepsilon_{q,t-1}$. Any shock that increases variances of the two correlated series would raise their covariance. The presence of $h_{q,t-i}$ in the conditional mean export equation implies that the system (6)–(11) is a bivariate GARCH-M model as developed by Engle and Kroner (1995).

Foreign income and the exchange rate are assumed exogenous as in Kroner and Lastrapes (1993). The benefits of endogenizing variables would have to be weighed against the costs of increased complexity of modeling and estimation. All parameters of (6)–(11) are estimated by maximum likelihood with the BHHH algorithm of Berndt *et al*. (1974).

3. DATA AND ESTIMATION OF TIME VARYING VARIANCE

Monthly data run from January 1979 to May 2001, a total of 269 observations. Seasonally adjusted real export revenue is nominal export revenue in domestic currency deflated by the wholesale price index. Foreign income is the export-weighted average of industrial production indexes of Taiwan's ten major exporting partners: the United States, Japan, Korea, Germany, Malaysia, Netherlands, the United Kingdom, France, Canada and Italy. The real effective exchange rate is a similar weighted average. The base year is 1995. All data come from the International Financial Statistics of the IMF, OECD Main Economic Indicators and the AREMOS data bank of Taiwan.

The correct specification of a GARCH model depends on whether variables are cointegrated. If so, the model should include an error correction term. Tests of the order of integration and cointegration, the augmented Dickey–Fuller (ADF) test for unit roots of Dickey and Fuller (1981), are reported in Table 1. None of the series exhibits a time trend. After selecting the minimum lag length required to assure lack of autocorrelation in the ADF regression, non-stationarity cannot be rejected at the 5% level. Non-stationarity can be rejected at the 5% level for every differenced series, implying the series are individually integrated of order one, I(1). Valid inference in GARCH models requires stationarity.

The I(1) series are tested for cointegration and Table 2 presents results from the Johansen (1988, 1991) approach with both maximal eigenvalue (λ_{max}) and trace statistics ('Trace'). The lag length of the VAR system is selected using

Table 1. ADF unit root tests[a]

Level	First difference	
lx_t	−1.6566(14)	−3.4702(13)*
ly_t	0.1668(13)	−4.1576(12)*
lq_t	−1.6231(13)	−4.8412(12)*

[a][Lag length]-selected to assure zero autocorrelation in ADF regression residuals.
*Significance at the 5% level.

Table 2. Cointegration tests (VAR lag = 4)

Eigenvalue	H_0^a	λ_{max}	95% critical value	Trace	95% critical value
0.0627	$r = 0$	17.08	20.97	26.5095	29.68
0.0350	$r \leq 1$	9.42	14.07	9.4292	15.41
0.00002	$r \leq 2$	0.0065	3.76	0.0065	3.76

[a]r is the number of cointegration vectors critical values from Osterwald-Lenum (1992).

AIC and SC criteria, and both suggest four lags. A Ljung–Box Q test on residuals finds no evidence of autocorrelation, suggesting that the model represents the autocorrelation structure. There is no evidence of cointegration. The Johansen test fails to reject the null hypothesis of zero cointegrating vectors, shown by insignificant λ_{max} and Trace statistics at the 5% level. There is no call for an error correction term in the GARCH export model.

The 1997 Asian financial crisis might have caused a structural break leading to the lack of cointegration. Gregory and Hansen (1996) suggest residual-based cointegration tests in the presence of potential structural breaks that could take the form of a change in intercept, a change in intercept with a time trend or a change in cointegrating slope coefficients. Residual-based ADF statistics testing for cointegration for these three structural breaks are –3.11 (11), –2.36 (11), and –3.40 (12), with lag truncation in parentheses selected on the basis of a *t*-test as suggested by Perron and Vogelsang (1992). Maximum lag length is set to 12 and tested downward until the last lag of the first difference is significant at the 5% level. Critical values from Gregory and Hansen (1996) at the 5% level are –4.92, –5.29 and –5.50 for the three cases, and ADF statistics fail to reject the null hypothesis of no cointegration. The 1997 Asian financial crisis does not change the conclusion of rejecting the null hypothesis of cointegration tests.

To specify a more appropriate bivariate GARCH-M model, consider preliminary statistics for log differences of export revenue and the exchange rate. Table 3 reports skewness statistics close to zero for Δlq_t, and the hypothesis that Δlx_t is distributed symmetrically can be rejected at the 5% level. Kurtosis statistics for Δlx_t and Δlq_t are significantly different from 3 at the 5% level, both variables are leptokurtic, and the Jarque–Bera test rejects normality. The Ljung–Box Q statistic tests for autocorrelation and the number of lags (k) affects its performance. Tsay (2002) suggests that the choice of $k = \ln(T)$ would provide better power performance. In the present data $k = \ln 269 = 5.59$ and autocorrelations are tested up to nine lags. Ljung–Box statistics indicate higher order autocorrelations in Δlx_t and Δlq_t. Ljung–Box statistics for the squared series suggest the possible presence of time-varying variance for Δlx_t and Δlq_t.

Squared serially correlated data may favor heteroskedasticity, and time-varying variance with a formal ARCH LM test as in Engle (1982) is reported in Table 4. After considering autocorrelations, LM statistics for an AR(9) process of Δlq_t and an AR(11) process of Δlx_t indicate that Δlq_t has significant higher-order heteroskedasticity, while Δlx_t has only weak first-order heteroskedasticity significant at the 10% level.

GARCH(1,1) models are estimated to identify changing variance for Δlx_t and Δlq_t. Ljung–Box Q-statistics for standardized residuals in Table 5 show no autocorrelations up to nine lags, suggesting that the AR processes are appropriately modeled to obtain white noise. The two GARCH(1,1) models capture heteroskedasticity, as shown by the low Ljung–Box statistics for the squared standardized residuals, LB $Q^2(k)$ up to nine lags. Nevertheless, for Δlq_t the estimate of the AR term β_2 in the conditional variance is not significant at the 5% level. For Δlx_t, both MA and AR terms β_1 and β_2 are insignificant.

Table 3. Preliminary statistics for export revenue and the exchange rate[a]

	Δlx_t	Δlq_t
Sample size	268	268
Mean	0.6045	−0.0183
SD	8.7018	1.5473
Maximum	37.3463	6.3571
Minimum	−25.9363	−4.6615
Skewness	0.4665*	0.0390
	(0.1496)	(0.1496)
Kurtosis	4.9229*	4.1091*
	(0.2993)	(0.2993)
J-B N	51.0076*	13.8065*
LB $Q(3)$	95.001*	4.2121
LB $Q(6)$	95.468*	13.927*
LB $Q(9)$	106.46*	23.034*
LB Q^2 (3)	34.151*	24.565*
LB Q^2 (6)	35.349*	25.436*
LB Q^2 (9)	36.439*	25.787*

[a]SD: standard deviation.
J-B N: Jacque–Bera normality test.
LB $Q(k)$, LB $Q^2(k)$: Ljung–Box statistics for level, squared terms for autocorrelations up to k lags.
*5% significance.

Table 4. The ARCH LM test[a]

k	Δlx_t	Δlq_t
1	3.6298**	12.3190*
2	3.6274	12.8388*
3	3.8991	15.3642*
4	3.9503	16.8353*
5	3.8097	17.5003*
6	7.3103	17.0761*

[a]LM(k) statistic follows a χ^2 distribution with k degrees of freedom, where $k = 1, 2, \ldots, 6$.
*5%; **10%.

More parsimonious ARCH(1) models are estimated for the two series as reported in Table 5, and Δlq_t behaves well. The two estimates in the conditional variance equation are significant at the 5% level, implying time-varying variance with short memory. The variance process is positive, finite and stationary, with $\beta_0 > 0$ and $1 > \beta_1 = 0.25 > 0$. There is no autocorrelation or heteroskedasticity in $\varepsilon_{q,t}$ as shown by Ljung–Box statistics for both the standardized (LB Q) and the squared standardized (LB Q^2) residuals up to nine lags. For Δlx_t there is no ARCH effect – not surprising, given the weak evidence of an ARCH effect in Table 4. ADF test statistics show that GARCH or ARCH processes are stationary, a common characteristic because volatility does not diverge and is often stationary. Finally, the likelihood ratio statistic tests whether the AR(1) term in the GARCH(1,1) process is zero, that is whether $\beta_2 = 0$ in (10), and it has a χ^2 distribution with one degree of freedom. The low LR value suggests the ARCH(1) specification sufficiently captures variance processes for Δlq_t or Δlx_t, although evidence shows no GARCH or ARCH

Table 5. Estimates of univariate GARCH models[a]

	Δlx_t		Δlq_t	
	GARCH(1,1)	ARCH(1)	GARCH(1,1)	ARCH(1)
s_0	0.8985**	0.9709*	0.0535	0.0758
	(0.4766)	(0.4934)	(0.0976)	(0.0994)
s_1	−0.7368*	−0.8351*	0.0528	0.0716
	(0.0758)	(0.0741)	(0.0768)	(0.0769)
s_2	−0.5227*	−0.5186*	0.0018	0.0022
	(0.0889)	(0.0831)	(0.0778)	(0.0715)
s_3	−0.1524	−0.1590**	0.0526	0.0810
	(0.0963)	(0.0939)	(0.0709)	(0.0566)
s_4	0.1017	0.1062	−0.0008	0.0037
	(0.0894)	(0.0882)	(0.0691)	(0.0697)
s_5	0.2796*	0.2735*	0.1242**	0.1171**
	(0.0904)	(0.0907)	(0.0644)	(0.0659)
s_6	0.2796*	0.2804*	0.1121*	0.1046**
	(0.0983)	(0.0960)	(0.0513)	(0.0535)
s_7	0.2022*	0.2067*	−0.0053	0.0074
	(0.0973)	(0.0926)	(0.0564)	(0.0547)
s_8	0.03610	0.0356	0.0803	0.0921
	(0.0995)	(0.1000)	(0.0684)	(0.0689)
s_9	0.0121	0.0094	−0.1193*	−0.1050**
	(0.0887)	(0.0888)	(0.0587)	(0.0616)
s_{10}	−0.1830*	−0.1822*		
	(0.0732)	(0.0755)		
β_0	13.9408	33.1212*	1.0858*	1.6595*
	(10.1381)	(3.4277)	(0.3496)	(0.1549)
β_1	0.1035	0.0921	0.2664*	0.2461*
	(0.0790)	(0.0789)	(0.1183)	(0.1135)
β_2	0.5157		0.2505	
	(0.3184)		(0.1857)	
LB $Q(3)$	0.8125	0.6385	0.1018	0.0484
LB $Q(6)$	0.8925	0.6591	0.9224	0.9816
LB $Q(9)$	4.4776	4.1265	2.7154	2.0304
LB $Q^2(3)$	0.0673	0.3679	1.0998	4.3018
LB $Q^2(6)$	3.9506	4.6510	2.3015	5.0553
LB $Q^2(9)$	9.2192	10.4413	2.9167	5.7229
ADF(n)	−8.27339(0)*	−15.0501(0)*	−9.4583(0)*	−12.2605(0)*
LR(1)		1.1192		1.8068

[a]Standard errors are in parentheses.
LB $Q(k)$, LB $Q^2(k)$: Ljung–Box statistics for autocorrelations up to k lags.
ADF tests for stationarity of the GARCH process.
LR(1) likelihood ratio statistic: χ^2 distribution with one degree of freedom that tests $\beta_2 = 0$.
*5%; **10%.

effect. The mean of the derived risk variable $h_{q,t}$ in this ARCH model is 2.21 with a standard deviation of 0.98 and a range of 1.66–11.1.

4. ESTIMATION OF THE EXCHANGE RISK MODEL

The lag length of the ADL specification of the export equation has to be determined, and Table 3 shows that Δlx_t has autocorrelation. Too few lags might not resolve autocorrelation in residuals, but too many would reduce the

Table 6. LR, AIC and SC values for choosing the order of the model

K/k – 1	*LR*	*k*	*AIC*	*SC*
6/5	10.4238*	6	6.5202	6.9016
5/4	8.2616	5	6.5046	6.8306
4/3	20.5646*	4	6.4810 min.[a]	6.7519
3/2	35.2884*	3	6.5040	6.7201 min.[a]
2/1	57.6106*	2	6.5821	6.7438
1/0	133.357*	1	6.7433	6.8507

[a]min: minimum value of AIC or SC.
*5% LR statistic.

degrees of freedom and estimation efficiency. According to the model in (2), the univariate ARCH process (3)–(5) generates $\hat{h}_{q,t-i}$. Three model selection criteria – LR, AIC and SC values – are used to determine the optimal ADL form. Alternatively, the bivariate GARCH model (6)–(11) can be used to choose lag length, but the reported approach is much more time efficient and the results are nearly the same. Table 6 reports the AIC and SC values along with LR statistics for testing lag k against k – 1 beginning at k = 6. The minimum AIC and SC are found at k = 4 and 3, respectively, and the first significant LR statistic is k = 5 at the 5% level. Models with k = 4 and 3 are considered, the model shows autocorrelations in residuals when k = 3, and k = 4 are chosen as the lag length.

Summarizing, there is statistical evidence of stationarity, lack of cointegration, leptokurticity, and heteroskedasticity in Tables 1–4. Univariate GARCH estimation in Table 5 and ADL lag length selection in Table 6 suggest the following eclectic bivariate GARCH-M model:

$$\Delta lx_t = a_0 + \sum_{i=1}^{4} a_i \Delta lx_{t-i} + \sum_{i=0}^{4} b_i \Delta ly_{t-i} + \sum_{i=0}^{4} c_i \Delta lq_{t-i} + \sum_{i=0}^{4} d_i h_{q,t-i} + \varepsilon_{x,t} \tag{12}$$

$$\Delta lq_t = s_0 + \sum_{i=1}^{9} s_i \Delta lq_{t-i} + \varepsilon_{q,t} \tag{13}$$

$$h_{x,t} = \alpha_0 \tag{14}$$

$$h_{q,t} = \beta_0 + \beta_1 \varepsilon^2_{q,t-1} \tag{15}$$

$$h_{xq,t} = \gamma_0 + \gamma_1 \varepsilon_{x,t-1} \varepsilon_{q,t-1} + \gamma_2 h_{xq,t-1} \tag{16}$$

This model simultaneously estimates time-varying risk ($h_{q,t-i}$). Variance of the export series is constant, and the variance of the exchange rate is identified as an ARCH process in Table 5, reducing the bivariate GARCH-M model to an ARCH-M model but retaining the property that exchange rate variance appears in the mean of the export equation. The information matrix of the system is not block diagonal, and joint estimation is efficient as developed by Kroner and Lastrapes (1993).

Table 7. Estimates for the bivariate ARCH-M model

	General		*Simple*	
	Coefficient	*Std error*	*Coefficient*	*Std error*
a_0	1.9966	2.5410	1.1676	1.3661
a_1	−0.8719*	0.0611	−0.9092*	0.0609
a_2	−0.5976*	0.0844	−0.6039*	0.0830
a_3	−0.2933*	0.0928	−0.3325*	0.0984
a_4	−0.0755	0.0712	−0.0975	0.0673
b_0	2.2643*	0.5954	1.9394*	0.5586
b_1	0.2221	0.7163		
b_2	2.0350*	0.6365	2.3582*	0.6248
b_3	−1.0865**	0.5905		
b_4	0.7998	0.6477		
c_0	0.3358	0.4512		
c_1	0.4947**	0.2753	0.5012*	0.2547
c_2	0.2151	0.2780		
c_3	0.2852	0.2781		
c_4	0.3605	0.2533		
d_0	−0.0905	0.6548		
d_1	0.3452	0.7110		
d_2	−2.0988	1.4848	−1.3003**	0.7747
d_3	2.2158	1.5098	1.2285**	0.7034
d_4	−0.8400	0.8413		
s_0	0.0698	0.0991	0.0598	0.0960
s_1	0.0647	0.0638		
s_2	0.0333	0.0716		
s_3	0.1011	0.0633	0.0978**	0.0565
s_4	0.1106**	0.0623		
s_5	0.1090**	0.0630		
s_6	0.1051*	0.0482	0.1051**	0.0539
s_7	−0.0277	0.0571		
s_8	0.1492*	0.0671	0.1288**	0.0693
s_9	−0.1722*	0.0571	−0.1101*	0.0561
α_0	34.0171*	3.2936	33.9229*	2.8489
β_0	1.833*	0.1889	1.7320*	0.1481
β_1	0.1870**	0.1106	0.2269*	0.1037
γ_0	−1.9652	1.4773	−0.7006	0.9287
γ_1	−0.2169*	0.0642	−0.0681	0.0970
γ_2	−0.1763	0.2247	−0.3333	0.8563
LR(15)[a]			18.5512	
LR(2)[b]			9.061*	

[a]LR(15): likelihood statistic that tests restriction from general to the simple model.
[b]LR(2): statistic tests significance of the two estimates of exchange rate risk in simple model.
*5% **10%.

Table 7 reports estimated coefficients and asymptotic *t*-statistics for a general unrestricted model and for a simple restricted version with insignificant variables deleted. A likelihood ratio statistic with χ^2 distribution and degrees of freedom equal to the number of restrictions is used to test validity of deleting variables.

The general model with lag length $k = 4$ is estimated first in Table 7. Although there is neither autocorrelation nor heteroskedasticity, there are insignificant coefficients making it difficult to gauge the impact of the risk. Following Hendry's (1985) 'general to simple' approach, fifteen insignificant variables are eliminated. Advantages of parsimony include higher precision of estimates resulting from reduced multicollinearity, increased degrees of freedom, more reliable estimates, greater power of tests and a simpler model. In this parsimonious process, the insignificant likelihood ratio statistic LR(15) = 18.5 at the 5% level suggests no explanatory difference between models. The variance process of exchange rate depreciation is positive and convergent with $\beta_0 > 0$ and $1 > 0.23 = \beta_1 > 0$. The mean of the risk variable $h_{q,t}$ in this ARCH-M model is 2.24 with a standard deviation of 0.92 and a range of 1.73–11.0.

Diagnostic tests support the statistical appropriateness of the bivariate ARCH-M model. Coefficients on exchange risk in the export equation are significantly different from zero at the 10% level. The likelihood ratio statistic LR(2) = 9.06 jointly tests the significance of $h_{q,t-2}$ and $h_{q,t-3}$. The null hypothesis that the export variable is independent of ARCH exchange risk is rejected at the 5% level. The LR test is more powerful than asymptotic *t*-tests as addressed by Kroner and Lastrapes (1993). Lagrange multiplier (LM) statistics test for autocorrelation and heteroskedasticity up to nine lags. A low LM statistic fails to reject the null hypothesis of no autocorrelation in residuals of $\varepsilon_{x,t}$ with LM(9) = 13.3 and in those of $\varepsilon_{q,t}$ with LM(9) = 6.35 at the 5% level, and both are white noise. The LM statistics for the squared standardized residuals, LM(9) = 2.50 for $\varepsilon_{x,t}$ and LM(9) = 6.55 for $\varepsilon_{qx,t}$, suggest a lack of heteroskedasticity.

5. ESTIMATED EFFECTS OF EXCHANGE RISK

The estimated export revenue function is

$$\begin{array}{llllll}
\Delta lx_t = & 1.17 & -0.91\Delta lx_{t-1} & -0.60\Delta lx_{t-2} & -0.33\Delta lx_{t-3} & -0.10\Delta lx_{t-4} \\
 & (0.08) & (-14.9)^{*} & (-7.28)^{*} & (-3.38)^{*} & (1.45) \\
 & +1.94\Delta ly_t & +2.36\Delta ly_{t-2} & +0.50\Delta lq_{t-1} & -1.30h_{t-2} & +1.23h_{t-3} \\
 & (3.47)^{*} & (37.7)^{*} & (1.97)^{*} & (-1.68)^{**} & (1.75)^{**}
\end{array} \tag{17}$$

Foreign income increases export revenue, as is characteristic of developed countries. The contemporaneous foreign income elasticity of export revenue equals the coefficient in (17) given the zero intercept term, $\%\Delta lx_t/\%\Delta ly_t = \Delta lx_t/\Delta ly_t = 1.94$, assuming no changes in other variables and no risk. The two-month-lagged effect is even larger and the cumulative foreign income elasticity is 4.30, reflecting the small open property of the economy. This quick adjustment is consistent with Arize's (1996a, 1997) evidence that exports increase instantaneously for G7 and eight European countries.

Quantitative impacts in (17) can be gauged by properties of exogenous variables. The mean value of the percentage change in foreign income Δly_t is 0.20%, implying an average cumulative effect after two months of $4.30 \times 0.20\%$ = 0.86%. The maximum percentage change in foreign income of 1.82% over the period implies a jump of 7.8% in export revenue. The largest monthly decline in foreign income was −2.31%, implying a shock of −9.9% to export revenue. Foreign recessions would evidently lead to declines in export revenue, and

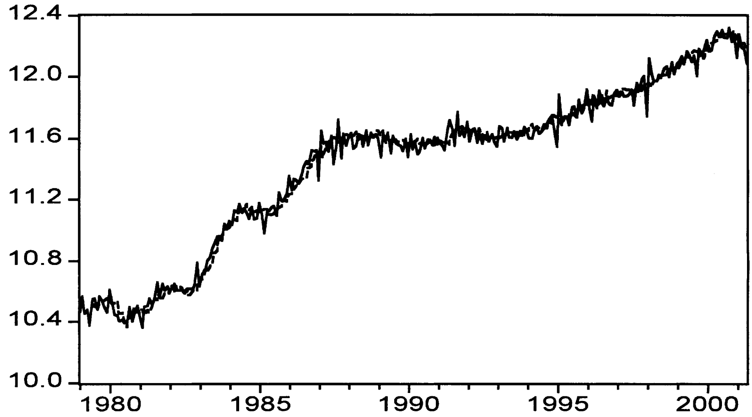

Figure 1. Export revenue
– Actual
– Fundamental

possibly to recessions, in Taiwan. The standard deviation of foreign income is 0.67%, implying that most observations fall in the range of 0.87% to −0.47% with associated effects on export revenue in the range of 3.7% to –2.0%. The average effect of foreign income is small, but it creates a lot of noise in export revenue.

Depreciation raises export revenue after only one month. Export adjustment speed in Taiwan is relatively quick, as Weliwita *et al.* (1999) find that at least eight months are required for a positive effect in Sri Lanka. The small exchange rate elasticity, however, implies that a 1% depreciation would raise export revenue by only 0.5%, holding foreign income and exchange risk constant. Depreciation of the nominal exchange rate would raise export revenue in terms of domestic currency but decrease it in terms of foreign currency, and there is no evidence of a subsequent J-curve.

The mean value of the change in the exchange rate is –0.018%. The coefficient for the exchange rate in (17) implies an almost trivial average monthly impact of –0.018% × 0.5 = –0.009%, although the leptokutic variable has long tails. The maximum and minimum Δlq_t yield a range of potential monthly impacts from 6.36 × 0.5 = 3.18% to −4.66 × 0.5 = –2.33%. The standard deviation of 1.55 suggests that most exchange rate changes fall in the range of 1.57 to –1.53, implying a range of effects on the exchange rate of 0.79% to −0.77%. While depreciation raises export revenue, effects of this size would hardly be noticed. Taiwan specializes in electronic products in small to medium-sized firms that can readily respond to the exchange rate.

Regarding exchange risk, the likelihood ratio statistic LR(2) indicates significant coefficients at the 5% level. The second and third month lagged effects of −1.30 and 1.23 are offsetting and sum to only −0.07. Regarding the size of this effect, the mean value of exchange risk $h_{q,t}$ in the ARCH-M model is 2.24. The ceteris paribus average impacts of risk on export revenue are 2.24 ×

(–1.30) = –2.91% after two months and 2.24 × 1.23 = 2.76% after three months. While these two effects are sizeable separately, their net impact is only −0.15%.

The maximum value of exchange risk over the period is 11.0, implying very large monthly impacts of −14.4% and 13.5%. These large effects are offsetting and sum to only 0.9%. The standard deviation of $h_{q,t}$ of 0.92 implies that most observations of risk fall in the range of 3.16 to 1.32. The implied range of effects after two months is −4.11% to –1.72%, and after three months 3.89% to 1.62%. While the transitory effects of exchange risk are large, exporters appear to be able to adjust rather quickly and completely to exchange rate risk.

As an experiment, a time trend of 'fundamental' export revenue is simulated to gauge the net effect of the exchange rate and its risk. Initial export revenue is used with the estimated coefficients in (17), except for the exchange rate and risk, to generate a fundamental model of export revenue. Figure 1 displays time plots for the actual and simulated fundamental series, and the difference is white noise. The mean of –0.001, a Ljung–Box Q-statistic of 15.6, and an LM ARCH test of 14.1, both up to the ninth lag, indicate that the difference is an independently identically distributed sequence with finite mean and variance at the 5% level. The exchange rate and its risk only add noise to the fundamentals of export revenue.

6. CONCLUSION

Time-varying real exchange rate risk has a negative impact on Taiwan's export revenue in a dynamic model, holding real foreign income and the real effective exchange rate constant. The negative impact of risk suggests that policy-makers should consider carefully any exchange market intervention that might be viewed by the market as transitory and unpredictable. Further, depreciation appears to lower export revenue in foreign currency, holding exchange risk constant.

Increased exchange rate risk has a large negative impact on export revenue in Taiwan after two months, but exporters appear to be able to adjust by the third month. Forward exchange rate cover is evidently incomplete, although adjustment to risk is quick and the net effect of risk is negligible. The exchange rate and its risk only add noise to underlying export revenue fundamentals, suggesting that exporters are able to sift through the noise of exchange rates and exchange risk.

REFERENCES

Arize, A. C. (1995) 'The Effects of Exchange-Rate Volatility on US Exports: An Empirical Investigation,' *Southern Economic Journal* 62, 34–43.

Arize, A. C. (1996a) 'Real Exchange Rate Volatility and Trade Flows: The Experience of Eight European Economies,' *International Review of Economics and Finance* 5, 187–205.

Arize, A. C. (1996b) 'The Impact of Exchange Rate Uncertainty on Export Growth: Evidence from Korean Data,' *International Economic Journal* 10, 49–60.

Arize, A. C. (1997) 'Foreign Trade and Exchange Rate Risk in the G7 Countries: Cointegration and Error Correction Models,' *Review of Financial Economics* 6, 95–112.

Arize, A. C., T. Osang and D. J. Slottje (2000) 'Exchange Rate Volatility and. Foreign Trade: Evidence from Thirteen LDCs,' *Journal of Business and Economic Statistics* 18, 10–7.

Asseery, A. and D. A. Peel (1991) 'The Effects of Exchange Rate Volatility on Exports: Some New Estimates,' *Economics Letters* 37, 173–7.

Berndt, E. K., B. H. Hall, R. E. Hall and J. A. Hausman (1974) 'Estimation and Inference in Nonlinear Structural Models,' *Annals of Economic Social Measurement* 4, 653–65.

Bollerslev, T. (1986) 'Generalized Autoregressive Conditional Heteroskedasticity,' *Journal of Econometrics* 31, 307–27.

Bollerslev, T. (1990) 'Modelling the Coherence in Short-run Nominal Exchange Rates: A Multivariate Generalized ARCH Model,' *Review of Economics and Statistics* 72, 498–505.

Bollerslev, T., R. J. Chou and K. F. Kroner (1992) 'ARCH Modeling in Finance: A Review of the Theory and Empirical Evidence,' *Journal of Econometrics* 52, 5–59.

Broll, U. and B. Eckwert (1999) 'Exchange Rate Volatility and International Trade,' *Southern Economic Journal* 66, 178–85.

Chang, T. Y., W. S. Fang, W. R. Liu and H. Thompson (2000) 'Exports, Imports and Income in Taiwan: An Examination of the Export Led Growth Hypothesis,' *International Economic Journal* 14, 151–60.

Chowdhury, A. R. (1993) 'Does Exchange Rate Volatility Depress Trade Flows? Evidence from Error Correction Models,' *Review of Economics and Statistics* 75, 700–6.

Darrat, A. F., M. K. Hsu and M. Zhong (2000) 'Testing Export Exogeneity in Taiwan: Further Evidence,' *Applied Economics Letters* 7, 563–7.

De Grauwe, P. (1988) 'Exchange Rate Variability and the Slowdown in Growth of International Trade,' *IMF Staff Papers* 35, 63–84.

Dickey, D. A. and W. A. Fuller (1981) 'Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root,' *Econometrica* 49, 1057–72.

Engle, R. F. (1982) 'Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation,' *Econometrica* 50, 987–1007.

Engle, R. F. and K. F. Kroner (1995) 'Multivariate Simultaneous Generalized ARCH,' *Econometric Theory* 11, 122–50.

Engle, R. F., D. M. Lilien and R. P. Robins (1987) 'Estimating Time-Varying Risk Premia in the Term Structure: The ARCH-M Model,' *Econometrica* 55, 391–407.

Ethier, W. (1973) 'International Trade and the Forward Exchange Market,' *American Economic Review* 63, 494–503.

Gregory, A. W. and B. E. Hansen (1996) 'Residual-based Tests for Cointegration in Models with Regime Shifts,' *Journal of Econometrics* 70, 99–126.

Hendry, D. (1985) 'Econometric Methodology'. Paper presented to the Econometric Society Fifth World Congress, MIT, Cambridge, MA.

Hodrick, R. J. and S. Srivastava (1984) 'An Investigation of Risk and Return in Forward Foreign Exchange,' *Journal of International Money and Finance* 3, 5–29.

Johansen, S. (1988) 'Statistical Analysis of Cointegration', *Journal of Economic Dynamics and Control* 12, 231–54.

Johansen, S. (1991) 'Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models,' *Econometrica* 59, 1551–80.

Kroner, K. F. and W. D. Lastrapes (1993) 'The Impact of Exchange Rate Volatility on International Trade: Reduced Form Estimates using the GARCH in Mean Model,' *Journal of International Money and Finance* 12, 298–318.

McKenzie, M. D. and R. D. Brooks (1997) 'The Impact of Exchange Rate Volatility on German–US Trade Flow,' *Journal of International Financial Markets, Institutions and Money* 7, 73–87.

Osterwald-Lenum, M. (1992) 'A Note with Quantiles of the Asymptotic Distribution if the Maximum Likelihood Cointegration Rank Test Statistics,' *Oxford Bulletin of Economics and Statistics* 54, 461–72.

Perron, P. and T. J. Vogelsang (1992) 'Non-stationarity and Level Shifts with an Application to Purchasing Power Parity,' *Journal of Business and Economic Statistics* 10, 301–20.

Pozo, S. (1992) 'Conditional Exchange Rate Volatility and the Volume of International Trade: Evidence from the early 1990s,' *Review of Economics and Statistics* 74, 325–9.

Tsay, R. S. (2002) *Analysis of Financial Time Series*. New York: John Wiley.

Weliwita, A., E. M. Ekanayake and H. Tsujii (1999) 'Real Exchange Rate Volatility and Sri Lanka's Exports to the Developed Countries, 1978–96,' *Journal of Economic Development* 24, 147–65.