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ABSTRACT

Australia has one of the more volatile set of export prices among OECD countries. This paper examines the extent to which Australia's export prices relate to the world prices using quarterly time-series data spanning the period 1969q4-2002q3. The empirical results based on dynamic least squares method show that Australia's export prices are cointegrated with the global export prices. A short-term dynamic ARCH-in Mean model, which captures the time varying nature of price volatility, has been used to explain the growth rate of Australia's export prices. It is found that (a) changes in Australia's export prices are highly associated with systematic changes in world export prices; (b) the diversification of Australia's export base has contributed to a significant reduction in the volatility of export prices during the study period; and (c) the time varying volatility has not undermined, in a significant manner, the growth rate of Australia's export prices.

EXPORT PRICE VOLATILITY IN AUSTRALIA: AN APPLICATION OF ARCH AND GARCH MODELS

I. INTRODUCTION

The volatility in Australia's export prices is an important source of national macroeconomic disturbance mainly due to the importance of exports in the composition of Australia's endowment bundle. Generally speaking, price volatility may be viewed as derived from country specific factors as well as from vicarious influences emanating from the global market place. This paper exposits a modeling framework for a prototype commodity exporting nation such as Australia which can capture the overall export price volatility highlighting the twin effects of idiosyncratic country specific factors as well as effects generated by the competitive global market place forces. This distinction can be useful from the policy debate point-of-view.

Motivation for the paper stems from recent economic debate in Australia concerning whether the country's export bundle is too narrowly based. As an open economy, Australia's exports constitutes about 22 per cent of GDP in 2002 and, in keeping with any open economy devoting a substantial proportion of its resources to export production, prices received for such exports are a crucial determinant of aggregate income and social welfare.

There are a significant number of empirical analyses which have investigated the effect of the terms of trade on Australia's economy (*e.g.* McTaggart and Rogers, 1990, Harvie and Tran, 1993, 1994, Gruen and Wilkinson, 1994, Fisher, 1996, Gruen, and Kortian, 1996). For example, Hoque (1995) examines the relationship between the terms of trade and current account outcomes in Australia. Based on his empirical findings, he asserts that the terms of trade impacted on Australia's current account balance during the fixed exchange rate regime but not during the flexible exchange rate era. Such a finding is consistent with insulation property perspective of floating exchange rate under the Mundell-Fleming model (Mundell, 1963 and Fleming, 1962).

In a more comprehensive study, Gruen and Dwyer (1996) investigate the interaction among the terms of trade, the real exchange rate and inflation and, *inter alia* they find that an increase in the terms of trade can result in inflationary pressures if the corresponding rise in the real exchange rate is less than 1/3-1/2 of the rise in the terms of trade. Kent (1997) and Cashin and McDermott (2002) in their cross-country analyses argue that, depending on the degree of persistence, the current account responds differently to the shocks associated with the countries' terms of trade. It is also posited that terms of trade shocks account for a considerable proportion of the volatility of current account balances in Australia and New Zealand (Cashin and McDermott, 2002).

Furthermore, some analysts examine the interplay between export prices and macroeconomic variables such as the exchange rate and its volatilities, the current account outcomes and the demand for imports and exports (*inter alia* see Chen and Devereux, 1994, Caselli, 1996, Arize, 1996 and Mahdavi, 2000). For instance, Chen and Devereux (1994) in their empirical investigation highlight the importance of the asymmetry between the effects of temporary import and export price shocks on the current account in the U.S and the U.K. In another study Caselli (1996) examines the relationship between the exchange rates and the export prices of various commodities in several European countries. He specifies an export price equation to measure the effects of nominal exchange rate movements on the exporters' profit margins and prices. Arize (1996) argues that the potential effect of the uncertainty associated with relative prices on export demand is important as he finds that in the U.K this uncertainty has had a negative effect on exports. Although these studies make substantial contribution to our knowledge of the effect of export prices on the economy, little is known about major determinants of Australia's export price index at a macro level.

Australia's export bundle mainly consists of primary products, whereas its imports are mostly manufactures with possibly more stable prices, supporting the proposition that changes in Australia's terms of trade "are largely the result of export prices changing by more than import prices" (McTaggart and Rogers, 1990, p.38). Hence, given the small-country price-taking assumption, this entails that the purchasing power of its exports (in terms of imports) can be subject to considerable fluctuations. For this reason, one may argue that Australian authorities should institute policies (*e.g.* tax incentives etc.) to encourage an expansion of the nation's export base into higher value-added industries. This could infuse more stability into the country's terms of trade, thereby reducing adverse economic effects from exogenous disturbances (Layton and Valadkhani, 2004).

The paper itself is divided into four main areas. The second section describes the data employed in the analysis and presents the unit root test results. The third section discusses the methodology employed to examine empirically the long and short-run determinants of Australia's export price. Various estimates of a short-term dynamic model capturing the growth rate of Australia's export price are presented in the fourth section. The paper ends with some brief concluding remarks.

II. THE DATA

Export prices are usually measured in index number form in terms of some selected base year. A given year's index value measures the level of the average price of an export bundle in that year as a proportion of the average price of the base year bundle. One source of such international price index data is International Financial Statistics (http://ifs.apdi.net/imf/logon.aspx) which publishes export unit value series (having the interpretation of implicit price deflators) for a wide selection of countries all expressed in US dollars. The most recent base used in IFS for various countries is 1995 and is the base used in this study. We have also used another variable in this paper denoted by Z which is the ratio of the exports of goods and services generated by high value-added non-primary industries (such as services and manufacturing) to total exports as a measure of the diversification of export base. Exports of mining and agricultural goods are assumed to be major primary industries in our definition. Figure 1 presents the graphs of the three variables employed in this study, namely P=Australia's export price index (1995=100); P^w=the world export price index (1995=100); and Z or the measure of diversification of Australian export base. The first two variables are obtained from the IFS website and are available for the period 1957q1-2002q3 and the last variable from Australian Bureau of Statistics (2004, Table 55) and this series is available for the period 1969q3-2003q3.

In order to make robust conclusion about the time series properties of the data we have used two unit root tests, *i.e* the ADF test and the Elliott-Rothenberg-Stock DF-GLS test. In this paper the lowest value of the Schwarz Criterion (SC) has been used to determine the optimal lag length in the testing procedure. These lags (reported in Table 1) augment the relevant regressions to ensure that the error term is white noise and free of serial correlation. Based on the results from the unit root tests presented in Table 1, we can conclude that all the three variables employed in this paper, *i.e.* $ln(P)$ and $ln(P^w)$ and $ln(Z)$, are I(1).

Table 1: Unit root test results

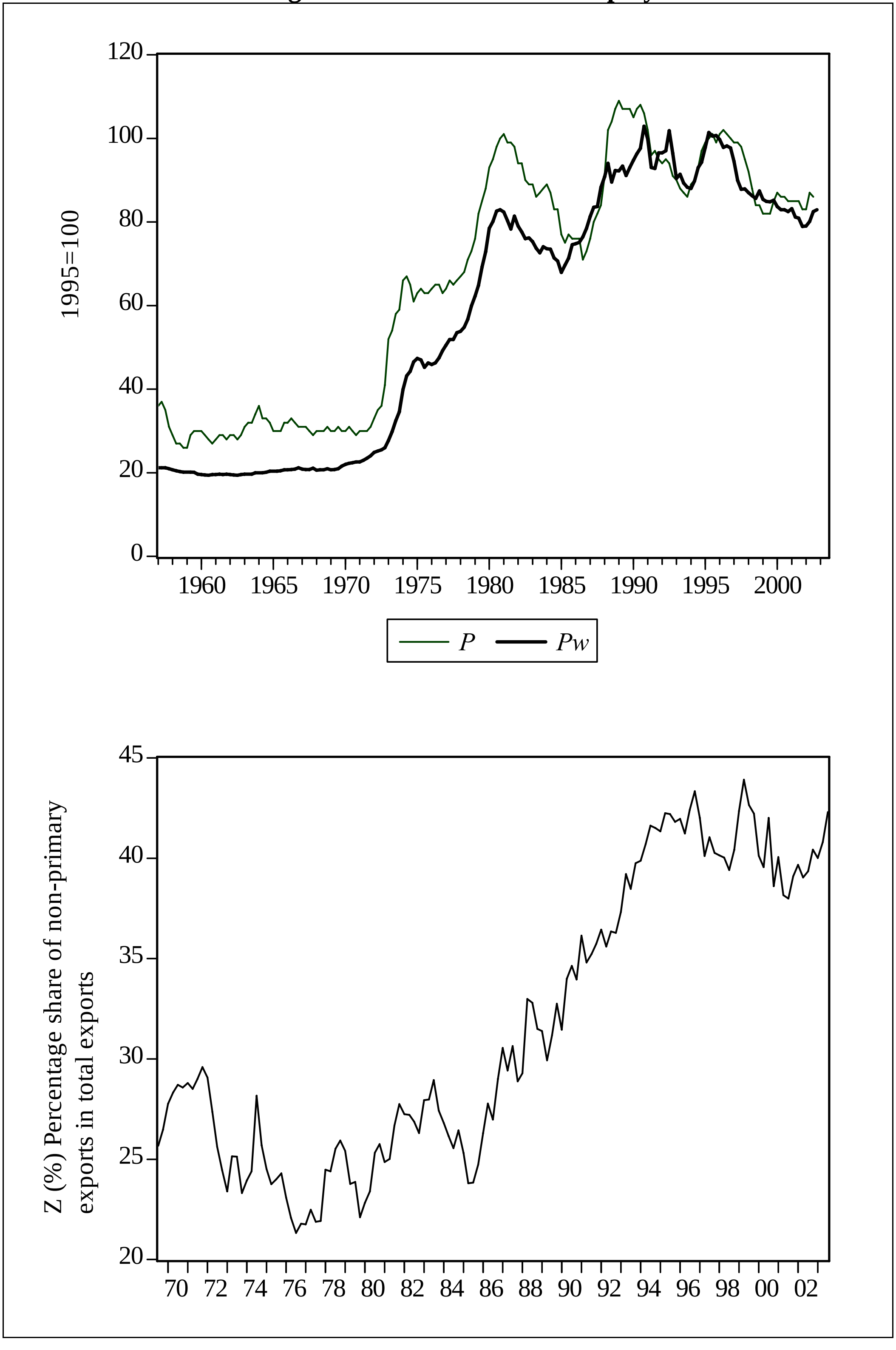
Variables	ADF	Optimal lag	Elliott-Rothenberg-Stock DF-GLS test statistic	Optimal lag
$ln(P_t)$	-1.91	1	-1.58	3
$\Delta ln(P_t)$	-7.36^*	0	-7.10^*	0
$ln(P_t^w)$	-1.43	1	-0.49	1
$\Delta ln(P_t^w)$	-7.45^*	0	-6.86^*	0
$ln(Z_t)$	-2.38	0	-1.97	0
$\Delta ln(Z_t)$	-11.98^*	0	-11.20^*	0

Note: * indicates that the corresponding null hypothesis is rejected at the 5% significance level.

Figure 2 shows the plots of the quarterly growth rate of export prices for Australia and the world as a whole during the period 1957q2-2002q3. An informal inspection of this graph supports the fact that Australia's export prices are more volatile than that of the world, particularly until the early 1990s. Furthermore, a similar conclusion emerges using standard deviation as a measure of volatility. During the same period, the standard deviations of quarterly growth rates of export prices in Australia and the World were 0.0406 and 0.0274, respectively.

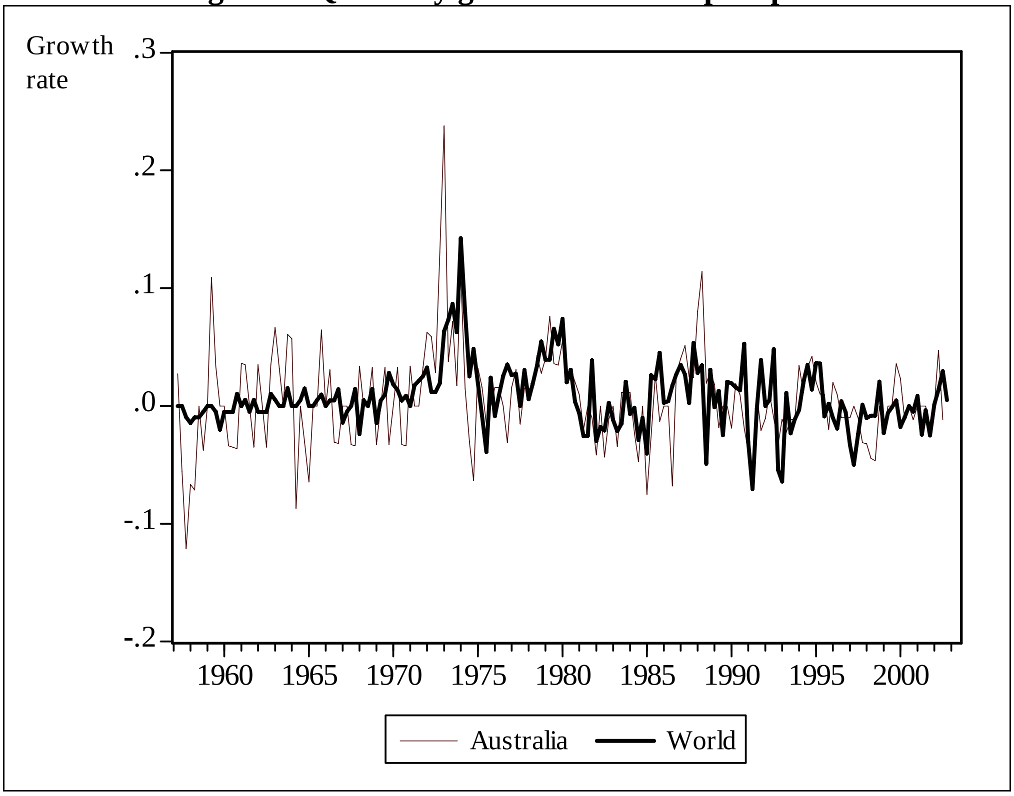
According to Table 2, not only did Australia's coefficient of variation (CV) increase in the post-1975 period (from 683% to 900%) but its international rank also increased from the 5th highest to the 2nd highest (only Finland had a higher CV). In both absolute and relative terms then the volatility of Australia's export prices has increased markedly between the two sub-periods. These results therefore seem to lend some *prima facie* support to those who argue Australia's export prices are relatively "too" volatile.

Figure 1: Plots of the data employed



Sources: (1) http://ifs.apdi.net/imf/logon.aspx; and (2) ABS (2004, Table 55).

Figure 2: Quarterly growth rates of export price



Sources: http://ifs.apdi.net/imf/logon.aspx

Given the fact that the volatility of Australia's export prices has increased through time, it would be useful to measure its effect on the growth rate of prices. This involves the use of an ARCH-in-mean model to test whether or not the error variance has affected Australia's export prices in a significant manner. It should be recognized that conditional variances can be interpreted as temporary increases or decreases in uncertainty and as such there is a possibility that quarterly price changes react to the changes in uncertainties in international markets. Therefore, an ARCH-in mean model will be specified in the next section to substantiate the effect (if any) of the time varying nature of volatility on Australia's export prices.

III. THEORETICAL FRAMEWORK

We hypothesise that a country's export price variation consists of two components: the first component pertains to overall global macroeconomic factors and the second component associated with more localised factors affecting that particular country – called, say, country-specific volatility. To enable the factoring out of global from country-specific volatility, Australia's export price growth series is regressed against the export price growth rate of some appropriate proxy for the global export portfolio. A higher coefficient of variation of such an equation means that the fortunes of the country are closely tied to internationally common macroeconomic factors.

Given that both price indices are integrated of order one, the dynamic least squares (DLS) technique is used to generate optimal multivariate estimators of the cointegrating parameters in the following manner:

$$\ln(P_t) = \delta_0 + \delta_1 \ln(P_t^w) + \sum_{i=-4}^{k=4} \pi_i \Delta \ln(P_{t-i}^w) + e_t \qquad (1)$$

It is argued that OLS can be used to estimate this equation and this DLS technique provides a consistent estimate of the cointegrating parameters (δ_0 and δ_1). The lags and leads of the first difference of the independent variable augment a standard OLS regression to remove the effects of regressor endogeneity on the distribution of the OLS estimator. The DLS estimators will be consistent in spite of the fact the residual term in equation (1) could be correlated with the right hand side variables. It is worth noting that "OLS estimators of the cointegrating parameters are "superconsistent", converging to the true parameter values at a rate proportional to the sample size T rather than proportional to $\sqrt{T}$ as in ordinary applications" (Lettau and Ludvigson, 2001, p.823). For a more detailed account of the DLS see Stock and Watson (1993).[i]

Table 2: Volatility statistics of the growth of export unit prices for selected OECD countries

Country	Pre-1975 Period					Post-1975 Period				
	Mean	SD	CV %	SD Rank	CV Rank	Mean	SD	CV %	SD Rank	CV Rank
Australia	0.020	0.135	683	13	11	0.009	0.084	900	8	14
Canada	0.037	0.075	200	3	2	0.024	0.067	285	2	3
Finland	0.030	0.270	905	15	14	0.010	0.110	1070	14	15
Germany	0.054	0.068	127	2	1	0.016	0.106	657	13	11
Ireland	0.027	0.084	309	6	6	0.022	0.076	349	5	8
Italy	0.016	0.086	546	8	9	0.027	0.092	342	10	6
Japan	0.011	0.078	728	4	12	0.03	0.074	247	4	2
Netherlands	0.022	0.102	463	9	8	0.018	0.080	448	6	9
Norway	0.029	0.119	416	11	7	0.02	0.116	581	15	10
NZ	0.020	0.123	606	12	10	0.025	0.083	327	7	5
Spain	0.009	0.110	1191	10	15	0.014	0.102	754	12	13
Sweden	0.019	0.158	837	14	13	0.014	0.102	752	11	12
UK	0.029	0.082	284	5	5	0.029	0.088	304	9	4
US	0.032	0.064	201	1	3	0.025	0.040	163	1	1
World	0.035	0.084	244	7	4	0.020	0.070	342	3	7

Sources: (1) International Monetary Fund (2003) on-line IFS database. (2) Layton and Valadkhani (2004).
Notes: A higher rank means that the corresponding statistic for the country in question is higher compared with the other countries in the set (rank goes from 1 to 15). SD=standard deviation and CV=coefficient of variation.

Starting with a maximum number of four lags and four leads (*i.e.* $k=\pm 4$), the general-to-specific methodology is now used to omit the insignificant π_i coefficients on the right hand side of equation (1) and this has been achieved on the basis of a battery of maximum likelihood tests. The estimation results for the parsimonious specific model capturing the long-run function are presented below (for compactness the coefficients estimates on the first lagged and lead differences are not shown below but they are available from the authors upon request):

$$\ln(P_t) = 1.08427 + 0.771\ln(P_t^w)$$

t: (31.7) (88.6) (2)

$\bar{R}^2 = 0.974$ ADF(residuals)=-4.5

The optimal long-run coefficients are seen to be of consistent sign and order of magnitude and are highly significant. This equation performs very well in terms of goodness-of-fit (adjusted $R^2 = 0.974$) and it generates white noise residuals. It is clear that Australia's export price index is well explained by common global macroeconomic factors, captured by the world export prices. The above slope coefficient may be interpreted as a measure of Australia's price sensitivity to common global fluctuations. A larger δ_l implies a country is relatively more sensitive to systematic global factors. We have also tested the null of δ_l=1, and the Wald test results, *i.e.* $F(1,174)=696.2$ [p-value=0.000], indicate that this hypothesis is rejected at any conventional significance level.

We can now calculate the error correction residuals from equation (2) as follows:

$$ECM = \ln(P_t) - \left[1.08427 + 0.771\ln(P_t^w)\right] \tag{3}$$

Initially a conventional short-term error correction model was estimated but the correlogram of squared residuals of such a model exhibited significant ARCH (Autoregressive Conditional Heteroskedasticity) effects (see Figure 3). Therefore, in order to capture any possible ARCH and GARCH (Generalised Autoregressive Conditional Heteroskedasticity) effects, the following ARCH-in mean (Engle, Lilien and Robins, 1987; Zakoïan, 1994; and Bollerslev, 1986, 2001) will be put into test in this paper:

$$\Delta\ln(P_t) = \alpha + \sum_{i=0}^{k=4}\theta_i\Delta\ln(P_{t-i}^w) + \sum_{i=1}^{k=4}\eta_i\Delta\ln(P_{t-i}) + \phi ECM_{t-1} + \gamma\sqrt{h_t} + \omega D83 + u_t \tag{4}$$

$$u_t = \varepsilon_t\left(\lambda_0 + \sum_{i=1}^{q}\alpha_i u_{t-i}^2 + \sum_{j=1}^{p}\beta_j h_{t-j}^2 + \lambda_1\Delta\ln(Z_t)\right)^{1/2} \tag{5}$$

$$h_t = \lambda_0 + \sum_{i=1}^{q}\alpha_i u_{t-i}^2 + \sum_{j=1}^{p}\beta_j h_{t-j}^2 + \lambda_1\Delta\ln(Z_t) \tag{6}$$

where α and λ_0 are the corresponding intercept terms in the mean and variance equations, respectively, θ_i shows the responsiveness of the growth of Australia's export prices to the current and lagged growth rates of world export prices, η_i up to four quarters are added to the dynamic model to ensure the resulting residuals are white noise, ϕ captures the error correction mechanism derived from the estimated equation (3), the estimated coefficient γ is referred to as a measure of the risk-return tradeoff in financial econometrics but in this paper this term indicates that the conditional mean of ΔlnP depends on the conditional standard deviation obtained from equation (6), α_i and β_j are the ARCH and GARCH coefficients, respectively, q is the order of the moving average ARCH terms, p is the order of the autoregressive GARCH terms, and the estimated coefficient on λ_1 captures the effect of Australia's export diversification measure (Z_t) on price volatility. These types of models are usually employed in financial econometrics to test the effect of the expected asset risk on the expected return on an asset.

In equation (4), a sustained dummy variable (*D83*) has also been inserted to capture the effect of Australian dollar being floated in December 1983. This dummy variable takes the value of 1 in and after the third quarter of 1983 and zero elsewhere.

As seen from Figure 4, the average growth rate of Australia's export price (defined as quarter-by-quarter log differences) was +0.008324 during the pre-floating period, whereas this rate declined to almost zero (*i.e.* -0.000152) in the post-floating exchange rate regime. Based on this observation it is expected that ω<0.

IV. EMPIRICAL RESULTS

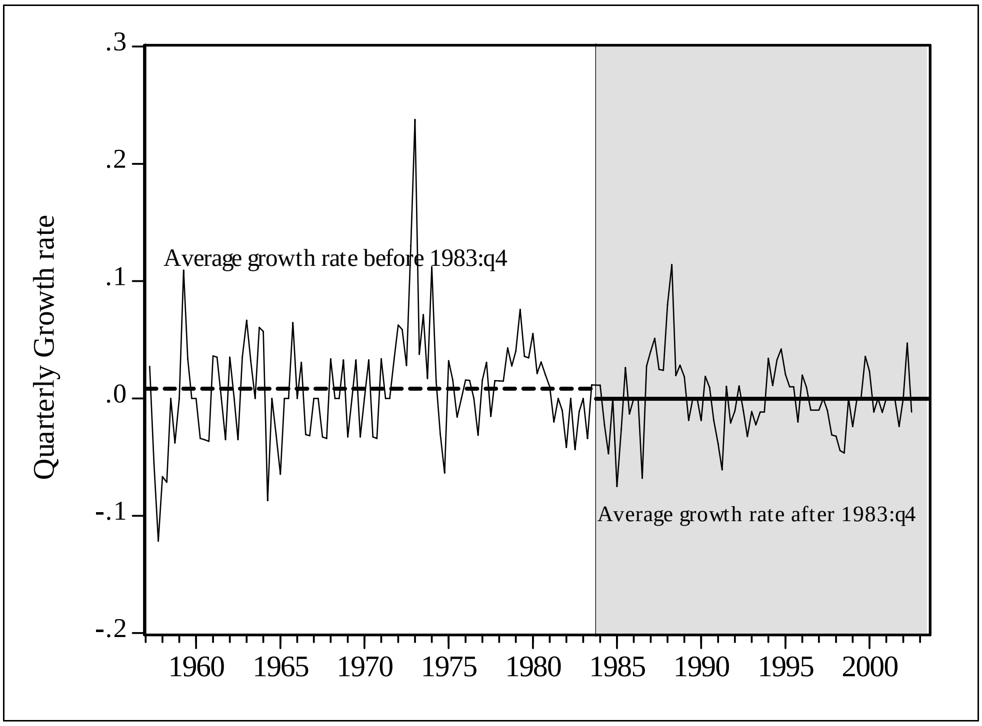
As mentioned earlier, a cursory look the correlograms of residuals for the estimated short-run dynamic, without capturing ARCH and GARCH effects, reported in Figure 3 exhibits volatility clustering. Once the ARCH and GARCH effects or the conditional heteroskedasticity in the residuals are modeled, as described in equations (4) to (6), the correlograms of the resulting residuals appear to be more statistically acceptable (see Figure 5).

Figure 3: Correlogram of squared residuals for the estimated short-run dynamic before capturing ARCH and GARCH effects

Autocorrelation	Partial Correlation		AC	PAC	Q-Stat	Prob
		1	0.202	0.202	7.2394	0.007
		2	-0.023	-0.067	7.3357	0.026
		3	0.069	0.092	8.2055	0.042
		4	0.107	0.075	10.286	0.036
		5	0.026	-0.005	10.410	0.064
		6	0.026	0.028	10.535	0.104
		7	0.188	0.175	17.036	0.017
		8	0.065	-0.019	17.814	0.023
		9	0.052	0.063	18.311	0.032
		10	0.062	0.021	19.025	0.040
		11	-0.065	-0.121	19.832	0.048
		12	-0.016	0.016	19.884	0.069
		13	0.066	0.045	20.717	0.079
		14	-0.047	-0.118	21.150	0.098
		15	-0.007	0.050	21.158	0.132
		16	0.029	-0.007	21.318	0.167
		17	-0.036	-0.078	21.576	0.202
		18	-0.072	0.001	22.591	0.207
		19	-0.034	-0.024	22.826	0.245
		20	0.037	0.023	23.094	0.284
		21	-0.039	0.002	23.402	0.323
		22	-0.024	-0.016	23.514	0.373
		23	-0.054	-0.067	24.118	0.397
		24	-0.059	0.004	24.822	0.415
		25	0.007	0.021	24.833	0.472
		26	0.042	0.054	25.206	0.507
		27	-0.046	-0.042	25.643	0.538
		28	-0.030	-0.004	25.835	0.582
		29	0.067	0.067	26.798	0.583
		30	-0.057	-0.082	27.504	0.597
		31	-0.074	-0.010	28.674	0.586
		32	-0.027	-0.019	28.830	0.628
		33	-0.005	-0.036	28.834	0.675
		34	0.006	0.048	28.842	0.718
		35	0.132	0.144	32.685	0.580
		36	0.038	-0.056	33.014	0.611

Source: authors' calculations.

Figure 4: Quarterly growth rate before and after Australian dollar was floated in December 1983



Source: International Monetary Fund (2003) on-line IFS database.

Three different versions of equation (4) have been estimated and the results presented in Table 3. As can be seen, if the ARCH and GARCH effects are not dealt with, the estimated equation cannot pass the ARCH test with various lag length. Therefore, it is important to capture these effects by a GARCH (*p*,*q*) process in equation (4). Assuming that γ=0 and $\gamma\neq0$, Table 3 presents the econometric results of the two different forms of the estimated equation (4) using maximum likelihood method. One can observe that the estimated γ is insignificant thus rendering the ARCH-in mean model irrelevant. However, the last three columns of Table 3 presents the results of our preferred equation where γ=0. It should be noted that only our preferred GARCH (1,1) model, with the lowest SC and the highest $\bar{R}^2$, passes various ARCH tests reported in Table 3 and its resulting correlogram is well-behaved (see Figure 5). Due to Bollerslev (1986, theorem 1), the preferred equation also satisfy the stationarity of the parsimonious GARCH (*p*=1,*q*=1) process as $\alpha_1 + \beta_1 < 1$. It should be noted that the SC and significant spikes in the relevant correlogram of squared residuals are used to determine *p* and *q*.

When *D83* and ARCH and GARCH effects are excluded from the model, the adjusted R^2 would be around 0.33 (these results have not been reported in Table 3 but they are available from the authors upon request). Thus, one can argue that only about one-third of the short-term variation of Australia's export prices is explained by systematic global factors. In other words, in terms of reducing export price volatility risk, Australia can still benefit from diversifying its export base. Of course, it goes without saying that to accomplish this may necessitate Australia incurring very significant opportunity costs of inefficiently using its scarce resources for producing in areas other than where its natural comparative advantages lie.

Figure 5: Correlogram of squared residuals for the estimated short-run dynamic model after capturing ARCH and GARCH effects

Autocorrelation	Partial Correlation		AC	PAC	Q-Stat	Prob
		1	0.014	0.014	0.0264	0.871
		2	-0.082	-0.083	0.9481	0.622
		3	0.090	0.093	2.0524	0.562
		4	0.097	0.088	3.3576	0.500
		5	-0.078	-0.068	4.2154	0.519
		6	0.101	0.113	5.6391	0.465
		7	0.050	0.018	5.9882	0.541
		8	0.112	0.136	7.7677	0.456
		9	0.112	0.115	9.5780	0.386
		10	0.028	0.015	9.6879	0.468
		11	-0.105	-0.101	11.288	0.419
		12	0.050	0.010	11.663	0.473
		13	0.047	0.017	11.988	0.529
		14	-0.107	-0.108	13.709	0.472
		15	0.082	0.082	14.720	0.472
		16	0.062	-0.015	15.303	0.503
		17	0.021	0.045	15.373	0.569
		18	-0.034	-0.036	15.553	0.624
		19	-0.022	-0.042	15.626	0.682
		20	-0.022	0.010	15.702	0.735
		21	-0.052	-0.081	16.141	0.762
		22	-0.084	-0.076	17.274	0.748
		23	-0.038	-0.055	17.505	0.784
		24	-0.020	-0.037	17.573	0.823
		25	-0.017	-0.037	17.621	0.858
		26	-0.038	-0.005	17.868	0.881
		27	-0.043	-0.015	18.182	0.898
		28	-0.046	-0.030	18.544	0.912
		29	-0.093	-0.057	20.020	0.892
		30	-0.011	0.021	20.042	0.915
		31	-0.064	-0.031	20.749	0.918
		32	-0.056	-0.043	21.310	0.925
		33	-0.061	-0.063	21.976	0.928
		34	-0.062	-0.058	22.679	0.931
		35	0.006	0.044	22.686	0.946
		36	0.015	0.032	22.726	0.958

Source: Authors' calculations.

Table 3: Modeling the short-run dynamics of $\delta ln(p_t)$ using equation (4)

Variables	Equation without ARCH effects using OLS method			ARCH-in Mean equation♣			GARCH (1,1) equation♣ (preferred model)		
	Coefficient	*z*-Statistic	Probability	Coefficient	*z*-Statistic	Probability	Coefficient	*z*-Statistic	Probability
Intercept	0.0073**	2.09	0.038	0.0070	0.83	0.407	0.0079*	1.84	0.065
$\Delta ln(P_t^w)$	0.4890**	5.07	0.000	0.3683**	3.90	0.000	0.3670**	3.71	0.000
$\Delta ln(P_{t-1})$	0.3082**	4.62	0.000	0.2553**	2.41	0.016	0.2735**	2.52	0.012
$\Delta ln(P_{t-3})$	0.2482**	3.63	0.000	0.2601**	4.16	0.000	0.2466**	3.74	0.000
D83	-0.0118**	-2.28	0.024	-0.0096*	-1.86	0.063	-0.0105*	-1.86	0.063
ECM_{t-1}	-0.1757**	-5.35	0.000	-0.1026**	-3.41	0.001	-0.1033**	-3.10	0.002
$\sqrt{h_t}$	-	-	-	0.0099	0.03	0.977	-	-	-
					Variance Equation			Variance Equation	
Intercept	-	-	-	0.0003**	3.33	0.001	0.0004**	5.30	0.000
u^2_{t-1}	-	-	-	0.3533**	3.18	0.002	0.3839**	3.11	0.002
h^2_{t-1}	-	-	-	0.2320	1.52	0.128	0.1975**	3.31	0.001
$\Delta ln(Z_t)$	-	-	-	-0.0041**	-383.90	0.000	-0.0043**	-3.52	0.000
Adjusted R^2	0.35226			0.366			0.373		
Durbin-Watson statistic	2.0202			1.94			1.97		
Akaike info criterion	-4.050			-4.329			-4.347		
Schwarz criterion	-3.941			-4.089			-4.129		
Overall *F*-statistic	19.9**		0.000	8.6**		0.000	9.7**		0.000
ARCH LM *F* Test:									
1 lag	7.298**		0.008	1.147		0.286	1.121		0.292
2 lag	4.306**		0.015	2.413**		0.101	2.189		0.111
3 lag	3.610**		0.015	1.828		0.145	1.790		0.152
4 lag	2.890**		0.024	1.271		0.285	1.281		0.281

Notes: * and ** indicate that the corresponding null hypothesis is rejected at 10 and 5 per cent significance levels, respectively. ♣ shows that the maximum likelihood (ML) method and the Berndt-Hal-Hall-Hausman optimization algorithm have been used in the estimation process.

Based on the last three columns of Table 3 (the results of our preferred model), the major findings of the paper are summarized below. First, to a large extent, the world export price index determines Australia's export prices both in the long- and short-run. The feed-back coefficient (or ϕ) is as low as -0.103, suggesting that in every quarter 10 per cent of the divergence between short-term price from its long-term path is eliminated. Based on this result, the adjustment appears to be reasonably slow. Similar result was also obtained by Yip and Wang (2002) for the equation for export prices in Hong Kong. Second, as it is expected, λ_1 is negative and highly significant in the variance equation, supporting the view that, *ceteris paribus*, increasing the share of Australia's non-primary exports in its total exports can reduce the volatility of export prices through time. Third, it appears that floating Australian dollar after the third quarter of 1983 has had a rather significant and negative effect on the average growth of Australia's export prices. This is consistent with what we have already observed in Figure 4. Fourth, the insignificant estimated coefficient (γ) on the time varying conditional standard deviation ($\sqrt{h_t}$) in Table 3 indicates that volatility itself has not exerted any impact on the growth rate of Australia's export price.

V. CONCLUDING REMARKS

This paper examines major sources of volatility in Australia's export prices (P) using a parsimonious GARCH (1,1) process augmented with two important variables, namely the world export price index (P^w) and the ratio of the exports of goods and services generated by non-primary industries to total exports (Z). These two variables capture both the global factors and country-specific peculiarities, respectively. A major finding of the study is that Australia's export prices are relatively more volatile in both the pre-1975 and post-1975 periods compared to that of other OECD countries. Furthermore, the empirical evidence reviewed in the paper suggests that during the period 1969q3-2002q3, about one-third of Australia's overall export price growth volatility could be attributed to global macroeconomic factors rather than domestic factors. Therefore, the remainder of the overall volatility of export prices may be regarded as emanating from country-specific volatility which is partly explained by Z. Hence, if policymakers consider that reduction of export price volatility is a desideratum, then this goal is achievable through the promotion of policies for the diversification of the country's export base.

It should be noted that this paper was concerned only with export earnings risk deriving from fluctuations in prices. Another important dimension of volatility relates to production quantities (*e.g.*, drought, in the case of rural exports) which, together with price volatility, will give rise to volatility in export revenues. Although this is certainly a valid point, the present paper focused exclusively on the price dimension since export price volatility has been a matter for grave concern in the contemporary policy debate and will continue to occupy the centre stage in the policy forum in years to come.

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ENDNOTES

[i] Given that Australia is a small open country and its export prices can not significantly affect the world export prices, it is plausible to assume that P^w is totally exogenous with respect to our dependent variable. For this reason, we have found very negligible differences between the magnitudes of the OLS and DLS estimators. In other words, similar results were obtained when the cointegrating vector was estimated using the Engle-Granger two-step procedure.