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Exchange Rate Pass-Through for Australian Manufactured Imports: Estimates from the Johansen Maximum-Likelihood Procedure

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ABSTRACT

This paper estimates exchange rate pass-through for Australian manufactured imports by applying an econometric procedure which avoids the pit-falls in previous studies to a carefully assembled data set. For the first time, we provide estimates of pass-through based on the Johansen (1988) ML procedure. Our finding of incomplete pass-through has important implications for policy and the macroeconomy. Incomplete pass-through brings into question the validity of the "small" country assmption and the exogeneity of the terms of trade with respect to exchange rate changes. We may also have to reconsider the extent of the apparent inflationary (deflationary) consequences of exchange rate depreciations (appreciations), and the effects of exchange rate variability on international trade flows.

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1. Introduction

The remarkable resilience that trade balances of major trading nations have shown to the wild fluctuations in floating exchange rates has been prominent in recent policy debates. The conventional explanations couched in terms of elasticity pessimism have little to offer in solving this "adjustment puzzle"; there is now a vast empirical literature that points to Marshall-Lerner conditions being easily satisfied in most countries [see Goldstein and Khan (1985)]. In this context, a number of authors have been motivated to step back and examine more closely the underlying relationship between exchange rates and prices of internationally traded goods, now popularly known as the *exchange rate pass-through* relationship. Exchange rate pass-through refers to the degree to which exchange rate changes are reflected in the destination currency prices of traded goods. If exchange rate changes are not fully or substantially reflected in the selling prices of traded goods, then the anticipated quantity adjustment will be retarded even if the degree of demand elasticity is sufficiently large. In this respect, knowledge of pass-through could be useful in accounting for the persistence in trade imbalances world-wide.

Reflecting the importance of the issues, studies on the pass-through relationship have thrived recently. The survey by Menon (1992a, Chapter 3) identifies 48 such studies, with the majority of them conducted during the past five years. These studies suffer from a number of short-comings, however. First, most researchers have ignored the time series properties of the data in conducting their estimations. Given that the data used to estimate pass-through is usually trended, it is likely that previous estimates of pass-through may have been biased as a result of the non-stationarity of the data. Second, most previous studies are subject to limitations imposed by inadequate data. In particular, there are reasons to suspect that the use of defective data with respect to both import and competitor "prices" would have biased the reported pass-through estimates. Third, much of the empirical work on pass-through has concentrated on the experience of "large country" cases, especially that of the US and Japan. Generalisations based on the experience of these countries for the smaller and more trade-dependent economies such as Australia would be highly questionable.

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This paper provides estimates of exchange rate pass-through for Australian imports of manufactures covering the period 1981q3 to 1992q2. This is done by applying an econometric procedure which avoids the pit-falls in previous studies to a carefully assembled data set. For the first time in the literature, we provide estimates of pass-through based on Johansen's (1988) Maximum Likelihood (ML) estimation of cointegration vectors. We use actual import prices, and overcome problems associated with the "world" price index by employing a foreign cost of production index instead. It is expected that evidence for Australia will add variety to a literature deficient in its coverage of small open economies.

The paper is organised in 4 sections. The model and data used to estimate pass-through is the subject of Section 2. Section 3 discusses the econometric methodology and presents the results. A final section considers the policy implications of our findings.

2. Model and Data

Previous studies in this area have employed import price equations derived within the mark-up framework popularised by Hooper and Mann (1989) to measure pass-through. Given that manufactured goods are typically differentiated and frequently sold in markets that are imperfectly competitive, the mark-up model seems appropriate in our case.

We begin by assuming that foreigners set their foreign currency export price (*PX*) as a mark-up (π) on their production cost in foreign currency (*CP*):

$$PX = \pi \, CP \tag{1}$$

The mark-up is expressed in ratio form (i.e., $\pi = (1 + \lambda)$, where λ is the profit margin). The Australian dollar import price (PM) is therefore given by:

$$PM = PX \, ER = (\pi \, CP) \, ER \tag{2}$$

The profit mark-up is hypothesised to depend on competitive pressures in the domestic market, and the exchange rate. The gap between the price of import-competing goods (*PD*) and the exporter's production cost is used to proxy the competitive pressure. The influence of domestic demand conditions on the import pricing decision is also captured by *PD*. The profit mark-up is thus modelled as:

$$\pi = \{PD \, / \, (CP \, ER)\}^{\alpha} \tag{3}$$

Substituting (3) into (2) we obtain:

$$PM \quad = \quad \{(PD \,/\, CP\, ER)^{\alpha}\}\; CP\, ER \tag{4}$$

Denoting logarithms of the variables as lower-case letters, and after some manipulation, we have:

$$pm \quad = \quad \alpha\, pd \;+\; (1\text{-}\alpha)\, cp \;+\; (1\text{-}\alpha)\, er \tag{5}$$

The model as specified above implies a rate of pass-through that is equal in magnitude for changes in foreign costs and the exchange rate. The coefficient of *pd* is α, which is the residual of the coefficients of *cp* and *er*. The cross-coefficient restrictions implied by this model suggests that *pd* has no (full) effect in determining import prices if pass-through is complete (zero). We will check these cross-coefficient restrictions when we estimate the model.

With respect to data, we have been fortunate to have gained access to actual import prices recently made available by the Australian Bureau of Statistics. Thus, we avoid the many problems associated with using price proxies such as import unit values, which most previous researchers have been forced to rely on [see Menon (1992a, pp. 139-41)]. Alterman (1991), for instance, compares the pass-through estimates obtained using import prices and import unit values, and concludes that the discrepancy is large enough to warrant concern over the reliability of estimates obtained using price proxies.

Most previous studies have employed a "world" price variable (in the form of import-weighted export prices) to capture changes in competitiveness of export sales to destination markets. The problem with this index is that it represents the pricing decision on exports to *all* markets, and is particularly inappropriate for "small" countries, and when "pricing to market" and incomplete pass-through behaviour is common-place for manufactures [see Menon (1992b)]. The findings of the studies overseas would imply that the "world" price variable *already* incorporates the incomplete pass-through on sales to other markets. To avoid these problems, we use an import-weighted[2] foreign cost of production index that is unaffected by the "pricing to market" problem. Unlike export prices, the cost of production index in any one country does not depend on the export market being targetted [see Knetter (1989)].

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(2) The 1985 import shares of the five major trading partner countries are used to construct the foreign cost of production and exchange rate index. The cost of production index for *each* trading partner is an input-output weighted average of wage and material costs. A detailed data appendix describing the method of construction and data sources is available on request.

3. Econometric Procedure and Results

We begin by testing for the presence of unit roots in the variables, using the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests, and the Johansen test for cointegration in one variable. The results from these tests are summarised in Table 1, and clearly indicate that all series are $I(1)$. In light of this, we proceeded to check if the level variables are able to form a cointegrating vector. There are two different approaches to testing for cointegration. They are the Engle-Granger (1987) two-step procedure and the Johansen (1988) ML procedure.

Table 1
Results of Unit Root Tests: DF/ADF and Johansen[1]

	DF/ADF	JOHANSEN
pm	-0.9047	3.2704
Δpm	-6.1596***	25.0155**
pd	-1.4967	4.2091
Δpd	-5.1916***	18.1587**
er	-1.0037	3.0409
Δer	-4.3118***	23.7492**
cp	-0.4950	0.1367
Δcp	-3.5915***	10.3020**

Notes:
(1) Δ is the (first) difference operator. For the DF and ADF tests, the significance levels were determined using the critical values reported in Mackinnon (1991). Critical values (sample size = 40): 10% = -2.60 (*), 5% = -2.93 (**), 1% = -3.58 (***). Critical values for the Johansen statistic is the likelihood ratio (LR) test statistic for cointegration in one variable based on maximum eigenvalue of the stochastic matrix. Critical values (sample size = 40) are: 10% = 6.5030 (*), 5% = 8.1760 (**).

The Engle-Granger procedure has been frequently employed in the literature, but suffers from a number of problems. First, should a cointegrating relationship be identified, the assumption is made that the cointegrating vector is unique. This need not be true in the multivariate case; if we denote the number of variables as n, then there can be up to n - 1 cointegrating vectors. If there is more than one cointegrating vector, the estimates from the Engle-Granger will be invalid. Second, there are concerns about the considerable small-sample bias in estimates from the Engle-Granger procedure. Stock (1987) shows that the bias in finite-samples will be in the order of $1/T$, where T is the sample size. Banerjee *et al* (1986) investigate this potential bias further, and show that it is related to $(1-R^2)$, and that this bias may decline much more slowly than the theoretical rate. Finally, the Engle-Granger procedure, unlike the Johansen procedure, is unable to accommodate dynamics in the cointegrating regression. Allowing short-run dynamics helps reduce biases and improve efficiency in the estimated cointegrating relationships.

For these reasons, we employ the Johansen ML procedure as the preferred test of cointegration and estimator. A formal exposition of the Johansen procedure can be found in Johansen (1988). Here we provide a brief and intuitive account of this procedure. The Johansen procedure starts from a general vector autoregression (VAR) model which is parameterised as a system error correction model so that the VAR consists mostly of lagged first difference terms and a set of lagged levels terms. If we denote the length of the lag chosen in the VAR as k, and the number of variables as n, then the VAR will contain (n x k-1) difference terms, and n levels terms each of which is lagged by k periods. It is clear that OLS could be applied to this system to provide a consistent estimate of the long run parameters for each equation, as demonstrated by Stock (1987). The problem, however, lies in the fact that these estimates may simply represent complex linear combinations of all the cointegrating vectors which link these variables together. If this happens we cannot interpret the resulting equation in an economically meaningful way.

It is not possible to separate out these individual relationships by OLS based methods. In order to achieve this an ML procedure is used and the following two sets of variables are defined: (i) the residuals obtained from regressing each of the contemporaneous difference terms on all of the lagged difference terms, and (ii) the residuals obtained from regressing each of the lagged levels terms on all of the lagged difference terms. The first set of variables can be viewed as the first difference of the data adjusted for the dynamics in the system, where as the second set represents the levels of the variables adjusted for the dynamics of the system. The first set will also be stationary by definition, where as the second set will be integrated at the order of the original data (in our case mostly $I(1)$).

The essence of the Johansen procedure lies with the realisation that the combinations of the levels variables which produce a high correlation with the difference variables are in fact the cointegrating vectors. Furthermore, the standard technique of canonical correlation will provide estimates of all of the distinct cointegrating vectors which may link a set of variables together. Finally, the associated eigenvalues may be used to construct a likelihood ratio test of the number of truly distinct cointegrating vectors which have been found.

The results of the Johansen ML estimation is reported in Table 2. The cointegrating parameters have been normalised on *pm*, and are based on the largest eigenvalues. To determine the lag length in the VAR, we used the likelihood ratio criterion. We started with a four lag system and tested down to the minimum number of significant lags using standard likelihood ratio tests[3]. Using this criterion the optimal lag length proved to be one, although the parameter estimates proved to be qualitatively unchanged for VAR lag lengths of 1 to 4[4].

The likelihood ratio statistics suggest that the hypothesis of one cointegrating vector is preferred. It is clear from the unrestricted cointegrating vector that exchange rate pass-through is incomplete (66.27 percent). We test two linear restrictions on the parameters of the cointegrating vectors implied by our theoretical model in Section 2. The first restricts the pass-through of foreign costs and exchange rate changes to be equal in magnitude. The LR statistic for this restriction cannot be rejected at the 1 percent level. Second, we impose the exclusion restriction on domestic competing prices, and find that this is rejected at the 5 percent level. The rejection of the exclusion restriction can also be viewed as a rejection of the full pass-through hypothesis.

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(3) The degrees of freedom correction proposed by Sims (1980) was utilised. An alternative procedure in determining the lag depth of the VAR is to use the Akaike Information Criterion [see Akaike (1974)]. This procedure involves identifying the lag depth at which the Akaike Information Criterion is minimised and then test down to the minimum number of jointly significant lags without inducing serial correlation in the residuals. We decided against employing this procedure in the light of the evidence of Sawa (1978), who finds that minimising the Akaike Information Criterion may lead to over-parameterisation.

(4) While the parameter estimates from the eigenvectors remain relatively invariant to extensions in the lag length in the VAR, the maximum number of unique cointegrating vectors (*r*) tends to increase for a given finite sample size [see Hall (1991)]. Since the optimal lag length our VAR proved to be 1, this problem did not interfere with our task of identifying an acceptable eigenvector.

Table 2
Johansen ML Estimates of Exchange Rate Pass-through[1]

(a) Tests for cointegration (VAR lag length = 1)

Null hypothesis	Likelihood ratio statistic	5% critical value
Number of cointegrating vectors r		
$r \leq 3$	3.264	3.762
$r \leq 2$	9.361	14.069
$r \leq 1$	17.175	20.967
$r = 0$	48.165	27.067

(b) Estimated cointegrating vector (largest eigenvalue only)

Unrestricted:

$$pm_t = 0.6627er_t + 0.7538cp_t + 0.3748pd_t$$

With equality restriction on pass-through of exchange rate and foreign cost changes:

$$pm_t = 0.6135er_t + 0.6135cp_t + 0.3421pd_t$$

Likelihood ratio statistic: LR(2) = 5.16

With exclusion restriction on domestic competing prices:

$$pm_t = 0.6051er_t + 0.8133cp_t + pd_t$$

Likelihood ratio statistic: LR(2) = 11.38

Notes:
(1) The LR(n) statistics are asymptotically $\chi^2(n)$ variates under the null; the 5 percent critical value for rejection of the null is 10.59.

4. Conclusion and Policy Implications

This paper has analysed the relationship between exchange rates and prices of Australian manufactured imports by applying an econometric procedure which avoids the pit-falls in previous studies to a carefully assembled data set. Our finding of incomplete pass-through has a number of implications for policy and the macroeconomy[5]. First, it rejects the small-country assumption of international price taking behaviour for Australia's imports. Second, it suggests that exchange rate changes are likely to lead to real effects in the economy operating through changes in the terms of trade. Third, it implies that exchange rate policy may be a blunt instrument when used to restore external balance since relative price adjustments may be limited. This might go some way towards explaining why both nominal and real imports have remained stubbornly high following the massive depreciation of the Australian dollar in the mid-1980s, despite a high elasticity of import demand. Finally, the extent of the inflationary (deflationary) effects of exchange rate depreciation (appreciation) operating through changes in the prices of imported goods will be limited.

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(5) For a more detailed discussion on the policy and macroeconomic implications of incomplete pass-through, see Menon (1993).

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