

SCALE ECONOMIES WITH REGARD TO PRICE ADJUSTMENT COSTS AND THE SPEED OF PRICE ADJUSTMENT IN AUSTRALIAN MANUFACTURING

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ABSTRACT

The standard quadratic price adjustment cost function makes no allowance for firm size or for scale economies. Incorporating quadratic price adjustment costs into the profit function, a firm's speed of price adjustment is both shown to be a positive/negative function of its size when firms have scale economies/diseconomies with regard to these costs and to be a negative function of market power. The intuitive explanation is that large firms that can defray this type of cost have less reason to slow price adjustment, while firms with market power are better able to offset price adjustment costs by slowing their speed of price adjustment. These results are used to derive an industry error correction model of pricing where the speed of price adjustment is a weighted average of the firm effects. Estimation of the model is carried out on data obtained from nine two-digit Australian manufacturing industries during the period 1994:3 to 2002:2. The empirical results suggest that the speed of price adjustment is positively related to the size of firms within an industry and negatively related to industry concentration. Given that these variables do not change rapidly over time, they are likely to have a steadying influence on the speed of price adjustment at an aggregate level in the face of changes to monetary and fiscal policy.

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1. Introduction

A common method of introducing price dynamics into a pricing equation is to add the costs of price adjustment directly into a firm's profit function. As part of a project to extend the microeconomic foundations of macroeconomics, Rotemberg (1982a, 1982b) models these costs as a quadratic function. This innovation has been particularly influential in the New Keynesian literature (see Roberts, 1995, 1992).

Martin (1993) starts from a profit equation that incorporates a quadratic price adjustment cost function and takes the theoretical analysis a step further by deriving the speed of price adjustment as a function of market power. Implicit in Martin's formulation is that the speed of price adjustment is also a function of firm size, although this point remains unexplored in his paper.

There are few references in the literature that posit a relationship between firm size and the speed of price adjustment. Domberger (1983) does suggest that large firms have large profit cushions and, as a result, are less risk averse, leading to faster speeds of price adjustment. However, his focus is on the role of industry concentration and the length of the production period, so he neither models nor tests this proposition.

A natural corollary to Domberger's suggestion is that firms have economies of scale over certain adjustment costs. This also happens to be a hidden assumption underlying quadratic price adjustment cost functions. In Section 2 of this paper, a model is derived that makes explicit the positive relationship between economies of scale in regard to quadratic price adjustment costs and the speed of price adjustment. The logic of the model is that firms that are able to defray these adjustment costs over a large output have less reason to slow their pace of price adjustment.

A second feature of the model is the negative relationship between market power and the speed of price adjustment, as per Martin (1983). This proposition has some support in the empirical literature, as there are a number of studies that use industry concentration as a proxy for market power and find a negative relationship with the speed of price adjustment (for example, Dixon (1983), Bedrossian and Moschos (1988), Weiss (1993) and Shaanan and Feinberg (1995) for Australian, Greek, Austrian and U.S. manufacturing, respectively). Finally, averaging across the industry transforms the model into an error correction form, which places fewer restrictions on the short-run dynamics of the estimating equation when compared to a partial adjustment model.

A review of the empirical literature indicates that the length of the production period is also considered an important determinant of the speed of price adjustment. Domberger (1983) suggests that inventories being valued at historical cost, rather than opportunity cost, sets up a disequilibrium wedge. This results in firms with short production periods placing a greater weight on the costs of disequilibrium, leading to greater speeds of price adjustment. In support, Dixon (1983) finds a negative relationship between the speed of price adjustment and the production lag for Australian manufacturing industries.

Australia, along with many other countries, has entered a period of relative price stability since the beginning of the 1990s. In spite of this, Dwyer and Leong (2001) do not find a shift in the speed of price adjustment at the aggregate level in the period after 1990 when compared to the 1985-1990 period. However, at a disaggregated level, some of the structural features thought to be important determinants of the speed of price adjustment have changed. For example, computerisation has allowed new approaches to inventory management, such as just-

in-time systems, to become widespread. Also, increased exposure to international markets has had an impact on the market power of local firms. For this reason alone, it is worthwhile taking a fresh look at the structural determinants of the speed of price adjustment at the industry level.

In Section 3, empirical testing of the model is carried out on quarterly two-digit Australian manufacturing data for the period 1994:3 to 2002:2. The results suggest that the speed of price adjustment at the industry level is positively related to the average size of sales for large firms within the industry and is negatively related to industry concentration. In this post-1990 period, the influence of the production lag on the speed of price adjustment appears to be negative for some statistical model specifications, but only at a low level of significance.

Section 4 discusses the implications of the theoretical and empirical results as they relate to particular macroeconomic and microeconomic policy areas.

2. The model

Consider an imperfectly competitive industry that consists of N firms, each producing a differentiated product. Let the short-run profit function of the i^{th} firm be:

$$\mathbf{p}(p_{it}) = (p_{it} - mc_{it})q_{it} - \mathbf{a}_i \left(\frac{p_{it} - p_{it-1}}{p_{it-1}} \right)^2 (q_{it}^*)^S \quad (1)$$

where i and t represent firm and time subscripts, respectively, and p_{it} , q_{it} , q_{it}^* , mc_{it} , \mathbf{a}_i and S indicate price, output, target output, constant marginal cost (excluding adjustment costs), a cost of adjustment parameter and an economies of scale parameter, respectively. It can be seen that the first term on the right-hand side of (1)

is revenue minus non-adjustment related costs, while the second term on this side is the cost of price adjustment.¹

When S is zero, the cost of price adjustment in (1) is the standard quadratic price adjustment cost function. This implies larger imposts on the firm for larger percentage price changes. Rotemberg (1982a, 1982b) cites unfavourable customer reaction to higher prices as an example of this type of cost. Presumably, the firm imputes a value to the loss of current and future goodwill when prices are raised to levels above expectations or when prices are increased well in advance of competitor prices. In a similar but alternative scenario, firms uncertain about market conditions may be unsure *ex ante* that a given target price is optimal and so impute a cost to rapid price change (for a discussion, see Domberger; 1983, pp 54-59).

Adjustment costs can also arise in input markets, with many authors pointing to turnover costs in relation to labour (see Kraft, 1995; Kasa, 1998; and Lindbeck and Snower, 2001). Given the rationing role of prices, it seems reasonable to model these adjustment costs in the form of quadratic price adjustments under certain conditions. For example, during a demand slowdown firms often hoard labour rather than face the costs associated with retrenching employees and then rehiring during the next upturn. A way of achieving this outcome and limiting losses in labour productivity is to maintain output levels through smaller price changes. In this paper, the quadratic price adjustment cost function is interpreted as representing an amalgam of implicit costs that can arise from adjustments in both product and input markets.

With the standard quadratic price adjustment cost function, the implicit cost to the firm of a given proportional price change remains the same regardless of firm size. Therefore, the absolute value of the cost of price adjustment would be the same for a

¹ Zero fixed costs are assumed for simplicity. This does not affect the analysis.

multinational company as for a local artisan (given the same \mathbf{a}_i). This only makes sense if there are extreme economies of scale. In order to allow for varying scale effects, the price adjustment cost is also a function of the firm's target output level. For a given price adjustment, it can be seen from (1) that the average cost of price adjustment declines with target output (economies of scale) when S is less than one; that it increases with target output (diseconomies of scale) when S is greater than one; and that it is constant when S is equal to one.

In the absence of adjustment costs, the first-order condition for profit maximisation is as follows:

$$q_{it}^* + (p_{it}^* - mc_{it})(dq_{it}^*/dp_{it}^*) = 0 \quad (2)$$

where * indicates the equilibrium values of price, output and the slope of the demand function. When adjustment costs are taken into consideration, q_{it}^* and p_{it}^* become the firm's target output and target price, respectively (this assumption is standard in the literature). Given that the actual price and the target price differ, firm output can be approximated using the following first-order Taylor series:

$$q_{it} \approx q_{it}^* + (dq_{it}^*/dp_{it}^*)(p_{it} - p_{it}^*) \quad (3)$$

Substituting (3) into (1) explicitly expresses profit as a function of price. After calculating the first-order profit maximising condition and incorporating (2) into the analysis, it can be shown that the firm chooses to change prices according to the following model:

$$\Delta p_{it} = \mathbf{I}_{it} (p_{it}^* - p_{it-1}) \quad (4)$$

$$\mathbf{I}_{it} = [1 - (\frac{\mathbf{a}_i \mathbf{b}_{it}}{\mathbf{h}_{it} p_{it}^* q_{it}^{*1-S}})]^{-1} \quad (5)$$

$$\mathbf{h}_{it} = \frac{p_{it}^*}{q_{it}^*} \frac{dq_{it}^*}{dp_{it}^*} \quad (6)$$

where $\Delta p_{it} = p_{it} - p_{it-1}$, \mathbf{I}_{it} is the speed of price adjustment, \mathbf{h}_{it} is the elasticity of demand and $\mathbf{b}_{it} = (p_{it}^* / p_{it-1})^2$. It is readily apparent that the range of \mathbf{I}_{it} is from zero to one and that (4) is just the partial adjustment model. Holding other things constant, it can be seen from (5) that the firm's speed of price adjustment increases/decreases with target output when the firm has economies/diseconomies of scale with respect to the costs of price adjustment; that firm revenue is positively correlated with the speed of price adjustment when S is zero; and that as demand becomes more/less elastic the firm's speed of price adjustment increases/decreases.

In order to give further direction to the empirical analysis in this paper, it is necessary to aggregate firm effects across the industry. Taking a weighted average of (4) across all firms in the industry and manipulating gives the following error correction model:

$$\Delta p_{dt} = \mathbf{g}_{dt} \Delta p_{dt}^* - \mathbf{I}_{dt} (p_{dt-1} - \mathbf{d}_{dt} p_{dt-1}^*) \quad (7)$$

where d is an industry subscript and $\Delta p_{dt} = \sum w_i \Delta p_{it}$, $\Delta p_{dt}^* = \sum w_i \Delta p_{it}^*$,

$$p_{dt-1} = \sum w_i p_{it-1}, p_{dt-1}^* = \sum w_i p_{it-1}^*, \mathbf{g}_{dt} = \left(\frac{\sum w_i p_{it}^* \mathbf{I}_{it}}{\sum w_i p_{it}^*} \right), \mathbf{I}_{dt} = \left(\frac{\sum w_i p_{it-1} \mathbf{I}_{it}}{\sum w_i p_{it-1}} \right) \text{ and}$$

$d_{dt} = g_{dt} / I_{dt}$. Following Bloch (1992), w_i represents the i^{th} firm's share of the value of industry shipments at a point in time. Therefore, the industry prices and target prices given in (7) are share-weighted averages. The error correction form of the model comes about because g_{dt} and the industry speed of price adjustment (I_{dt}) are differently weighted averages of each firm's speed of price adjustment.² If all firms in the industry have the same speeds of price adjustment, then the industry model reverts to the partial adjustment form.

In order to further inform the empirical analysis, the industry target price is derived when firms have log-linear and linear demand functions. The workings are shown in Appendix 1. In the former case, the elasticity of demand is exogenous and the industry target price is a linear function of the weighted average of each firm's marginal cost. In the latter case, the industry target price is a linear function of the weighted averages of each firm's marginal cost and demand shift variables.

Generally, pricing equations will be a function of cost and demand shift variables, except when the demand function is iso-elastic and moves proportionally (see Bloch, 1992; and Olive, 2002).

3. Data, the empirical model and estimation

Uncovering the structural determinants of industry speed of price adjustment for nine Australian manufacturing industries at the two-digit level during the period 1994:3 to 2002:2 provides the focus for this section. At the beginning of this period, the Australian Bureau of Statistics (ABS) switched to the Australian and New Zealand Standard Industry Classification (ANZSIC), thus making it difficult to extend particular data series beyond this date. Also, two-digit industries are examined

² This method for obtaining an error correction model could be contrasted with those outlined by Nickell (1985).

because many quarterly data series employed in this study do not exist at a lower level of aggregation. See the Data Appendix for a full description of the data series and sources.

3.1 Industry speed of price adjustment

Numerous studies find statistically positive relationships between industry concentration and price-cost margins. With regard to heterogeneous goods, Sawyer (1982) suggests that industry concentration may act on firm price conjectures, *inter alia*, to make demand less elastic and increase margins. This causal relationship is further discussed and applied to Australian manufacturing by Bloch (1992) and Bloch and Olive (1999). Following this reasoning, it is expected that an increase in industry concentration will reduce the industry speed of price adjustment. However, Katics and Petersen (1994) and Ghosal (2000) find that increasing import share weakens the market power of firms in highly concentrated U.S. manufacturing industries.

Therefore, the measure of industry concentration employed in this study is the ratio of sales from the four largest firms divided by the value of sales in the domestic market, which takes into account the increased volume of imported product that domestic firms in Australia have had to compete with through the 1990s.

Given sufficient economies of scale with regard to price adjustment costs, it is expected from (5) and (7) that average firm revenues will be positively correlated with the industry speed of price adjustment. In the empirical model, sales averaged across the four largest firms are used to represent average firm revenue. This seems reasonable as the largest firms are likely to have the greatest weights in (7).

Although it is not immediately obvious from the theoretical model how the length of production period should impact on the speed of price adjustment, it is

included in the empirical model because of its prominence in the literature. Dixon (1983) calculates the number of production lags dependent on the time it takes to progress from raw materials to work in progress to finished goods. Times are calculated from the value of stocks at each stage, the usage and sales rates, and assumptions about the continuous and progressive nature of the production process. A series for the length of production period in each industry is constructed using this method. From these calculations, it is clear that the number of production lags (in quarters) have decreased in most industries since the mid 1980s, but particularly in Printing Publishing and Recorded Media (24), Petroleum, Coal, Chemical and Associated Product Manufacturing (25) and Machinery and Equipment Manufacturing (28).

In Equation (9) below, the speed of price adjustment is modelled as fixed for any one industry over the period of examination, but as a linear function of the structural variables across industries. This cross-sectional approach is necessary because industry concentration is only available for particular years. However, taking industry values at a point in time may also better approximate the long-run values of the structural variables. Given reasonably small differences between lag period prices and target prices means that the cross-industry variation in the average industry b_{it} should be relatively small when compared to the cross-industry variation in the structural variables, and consequently, this influence is not incorporated into the empirical model.

3.2 Target price and the empirical model

O'Regan and Wilkinson (1997) estimate the long-run elasticity of domestic industry price with respect to import price for 30 Australian manufacturing industries over the

period 1983:2 to 1995:2 and find it significant in nearly every case. As the authors point out, cost and other demand shift variables are not included in the model, so the results may be overstated. Bloch and Olive (1999) estimate a pricing equation for 89 Australian manufacturing industries over the period 1971/72 to 1984/85. While they generally find that unit cost is the dominant influence on industry price, competing foreign price and manufacturing price are influential in highly concentrated industries exposed to foreign competition and aggregate income is influential in highly concentrated industries with low exposure to foreign competition.

Using the results from Appendix 1, industry target price is initially modelled as a linear function of average variable cost (ac_{dt}), import price (imp_{dt}), manufacturing price (pm_t) and aggregate income (y_t). These variables are positively related to industry target price if imports and product from other industries are largely substitutes for the domestic industry product and if manufacturing industries produce normal goods. Applying this to Equation (7), the empirical industry pricing equation is given as:

$$\Delta p_{dt} = \mathbf{q}_{d1} + \mathbf{q}_{d2} \Delta ac_{dt} + \mathbf{q}_{d3} \Delta imp_{dt} + \mathbf{q}_{d4} \Delta pm_t + \mathbf{q}_{d5} \Delta y_t - \mathbf{I}_d ECM_{dt-1} + \mathbf{e}_{dt} \quad (8)$$

$$ECM_{dt-1} = p_{dt-1} - \mathbf{f}_{d2} ac_{dt-1} - \mathbf{f}_{d3} imp_{dt-1} - \mathbf{f}_{d4} pm_{t-1} - \mathbf{f}_{d5} y_{t-1} \quad (9)$$

$$\mathbf{I}_d = (\ell_1 - \ell_2 CR4_d + \ell_3 SIZ4_d - \ell_4 PLAG_d) \quad (10)$$

where Δ indicates first difference, \mathbf{q}_{d1} to \mathbf{q}_{d5} and \mathbf{f}_{d2} to \mathbf{f}_{d5} are positive constants that vary across industry, ℓ_1 to ℓ_4 are positive constants that do not vary across industry, ECM_{dt-1} is the error correction mechanism, \mathbf{I}_d is the industry speed of price adjustment, $CR4_d$ is the domestic market four-firm concentration ratio, $SIZ4_d$ is the

average sales across the largest four domestic firms in the industry and $PLAG_d$ is the length of production period in quarters. The expected signs are shown in (8), (9) and (10). A feature of the empirical pricing equation is that the error correction mechanism and speed of price adjustment differ across industries, although the latter only differs to the degree that the structural variables are different in each industry. The error term (e_{dt}) in (8) is taken to be stationary, however the possibility of correlated errors across industries is left open as shocks that are unaccounted for in the model (for example, variability in consumer behaviour) may simultaneously influence numbers of manufacturing industries.

A consequence of non-stationary errors might be a spurious OLS regression that over-rejects the null hypothesis. In order to test whether the error term is indeed stationary, the time series properties of the data are investigated, with the results presented in Table 1. The Im, Pesaran and Shin (2003) test for unit roots in panel data indicates that indices of domestic industry price (p_{dt}), average variable cost (ac_{dt}) and import price (imp_{dt}) each have a unit root in levels but are first-difference stationary, while the weighted symmetric tau test (Pantula et al, 1994) indicates that an index of manufacturing price (pm_t) and a chain volume measure of GDP (y_t) also have unit roots in levels but are first-difference stationary. If these variables in levels are cointegrated, then the error correction mechanism and the error term are both stationary. Using Pedroni's (1999) Group ADF test for panel data, the null of no cointegration is rejected when p_{dt} , ac_{dt} , imp_{dt} and pm_t are included in the test, but is not rejected when y_t is added to this list. It should be noted that this test is only indicative, as it does not account for cross-sectional correlation in the error terms. In light of these results, the empirical model is estimated both with and without y_t .

Table 1
Tests for non-stationarity

Variable	Level	First Difference	Test Type
p_{dt}	3.75	-6.14**	IPS
ac_{dt}	-0.96	-10.53**	IPS
imp_{dt}	-0.71	-6.15**	IPS
pm_t	-0.07 [2]	-2.89* [6]	WS
y_t	0.08 [5]	-10.48** [0]	WS
cointegration excluding y_t	-2.94**		Group ADF
cointegration including y_t	-1.40		Group ADF

Panel data tests carried out using Pedroni program for RATS and individual data tests carried out using TSP.

IPS indicates Im, Pesaran and Shin (2003) test for unit roots in panel data.

WS indicates Pantula et al (1994) weighted symmetric tau test for unit roots.

Group ADF indicates Pedroni (1999) test for cointegration in panel data.

For each test the null hypothesis is non-stationarity. The panel data test statistics are z distributed under the null and all panel tests have four lags and no time trend.

WS tests have no time trend and seasonal dummy variables. Lag lengths are determined by the Akaike Information Criteria and shown in brackets.

** indicates significant at the 1% level.

* indicates significant at the 5% level.

3.3 Estimation

Equations (8), (9) and (10) are estimated in a two-stage procedure analogous to the single equation error-correction-based test of cointegration suggested by Kremers et al (1992). First, a third-order autoregressive distributive lag model is separately estimated for each industry using OLS. The estimates from this model are then used to numerically calculate the long-run solutions, namely the estimates of \mathbf{f}_{d2} , \mathbf{f}_{d3} , \mathbf{f}_{d4}

and f_{d5} shown in (9).³ The second stage involves estimating Equations (8) and (10) by the method of seemingly unrelated regressions (SUR) while constraining the coefficients in the error correction mechanism to their long run values.

Table 2

SUR results for the industry speed of price adjustment ($?_d$)

Variable	Time series variables in levels		Time series variables in natural logs	
	y_t excluded	y_t included	y_t excluded	y_t included
Constant	0.37** (3.51)	-0.09 (-0.51)	0.39** (3.80)	0.02 (0.12)
$CR4_d$	-0.68** (-5.62)	-0.53** (-3.22)	-0.74** (-6.01)	-0.59** (-3.38)
$SIZ4_d$	0.12** (5.87)	0.17** (5.00)	0.12** (5.84)	0.16** (4.36)
$PLAG_d$	-0.23* (-2.05)	0.22 (1.38)	-0.23* (-2.14)	0.14 (0.85)

Estimation carried out using TSP.

t statistics computed from heteroscedastic-consistent standard errors are in parentheses.

** indicates significant at the 1% level for a two-tailed t test.

* indicates significant at the 5% level for a two-tailed t test.

Table 2 presents the results for (10) when the coefficient estimates for the speed of price adjustment are constrained to be the same across industries and the other coefficients from (8) are unconstrained. The t-statistics presented in this table are computed from heteroscedastic-consistent standard errors. The four sets of results are for the cases when the model includes and excludes y_t , and when the time series

³ An ARDL(3, 3) model is chosen in order to capture lags that could occur for up to one year. Although an ARDL(1, 1) is more likely to be under-parameterised, it is closer in spirit to the economic model outlined in this paper. When the latter model is estimated the results are qualitatively similar to those presented in Table 2.

variables are in levels and in natural logs.⁴ Natural logs are alternatively employed in the model as a way of testing the sensitivity of the results to changes in the functional form of demand.

The regression results give support to the theoretical model when firms have economies of scale with regard to price adjustment costs. In the four sets of results, $SIZ4_d$ is positive and significant at the 1% level, suggesting that speed of price adjustment increases as the average firm sales increase. This outcome is consistent with the view that the impact of price adjustment costs decline as firm scale increases. In each of the estimate sets, $CR4_d$ is negative and significant at the 1% level, which conforms to the findings in Dixon (1983), Bedrossian and Moschos (1988), Weiss (1993) and Shaanan and Feinberg (1995). It suggests that firms in concentrated industries are in a better position to take price adjustment costs into account by slowing their speed of price adjustment.

By comparison, the results for the production lag are relatively weak with a negatively significant effect at the 5% level occurring only when y_t is excluded from the regression.⁵ One explanation for this outcome could be the result of changes in the way that inventories are managed in an age of computerisation, thus changing what behavioural impact that production lags have had on firm pricing policy.

⁴ As the series used in this study are in index form, not all of the regression estimates for (8), (9) and (10) are easy to interpret when series are in levels. Characterising an index as an unknown number multiplied by its true value, parameter estimates will generally be the true estimates multiplied by unknown constants. The exceptions are the speed of price adjustment estimates for (10) which will be the true estimates only when industry prices (and, therefore, lagged prices and price differences) are multiplied by the same unknown number. In contrast, if the time series variables are transformed into natural logs, the slope coefficient estimates can be directly understood as the unknown numbers fall into the estimates of the constant coefficients (i.e. the q_{d1}). Given the similarity of the results shown in Table 2, using series in levels does not seem to be a serious problem in the estimation of (10).

⁵ Kremers et al (1992) indicate that the speed of adjustment in the single equation has an asymptotic distribution that ranges from normal to Dickey-Fuller under the null of no cointegration. The exact nature of the distribution is determined by the short-run dynamics and the signal-to-noise ratio. In the absence of an exact analytical distribution for a panel under the null, the t-distribution is assumed. However, that Student's t ratio might over-reject the null is a further reason to describe the results for the production lag as weak.

Comparing across the results from Table 2 shows that the constant term is significantly positive when y_t is excluded and insignificantly different from zero when y_t is included. Also, when y_t is included, the absolute values of the coefficient estimates for average firm sales increase while (in absolute value) those for domestic market four-firm concentration decrease.

Implied speeds of price adjustment for each industry are obtained by multiplying the coefficient estimates by the industry structural variable values and summing. Table 3 presents the estimated speeds of price adjustment when the time series variables are in natural logs and y_t is included. Industries with the lowest speeds of price adjustment are Printing, Publishing and Recorded Media (22) and Other Manufacturing (26) with values of 0.08 and 0.11, respectively, while industries with the highest speeds of price adjustment are Petroleum, Coal, Chemical and Associated Product Manufacturing (25) and Machinery and Equipment Manufacturing (28) with values of 0.43 and 0.51, respectively. Clearly firm size is a dominant factor in the speed of price adjustment as the two latter industries have much larger firms on average than do the two former industries. Using the mean-lag formula $(1 - I_d / I_d)$, the maximum number of quarterly lags it takes for half the impact of a change in price to be felt for an industry is 11.5 lags when the speed of price adjustment is 0.08 and 1.0 lags when the speed of price adjustment is 0.51.

Apart from the speed of price adjustment and its determinants, the long-run elasticities of industry price with respect to average variable cost, import price, manufacturing price and aggregate income are also of interest, as they can be compared to previous studies. When the time series variables are expressed as natural

Table 3

Long-run elasticity estimates and implied speed of price adjustment when the time series variables are in natural logs

Industry	ac_{dt-1}	imp_{dt-1}	pm_{t-1}	y_{t-1}	$?_d$
(21) Food, Beverage and Tobacco Manufacturing	0.31** (13.30)	-0.16* (-2.00)	0.22** (3.89)	0.43** (8.62)	0.33** (12.07)
(22) Textile, Clothing, Footwear and Leather Manufacturing	0.27** (7.18)	-0.01 (-0.52)	-0.15* (-2.09)	0.41** (10.36)	0.19** (3.02)
(23) Wood and Paper Product Manufacturing	-0.12 (-0.20)	0.12 (0.69)	0.51 (0.73)	0.11 (0.71)	0.18** (6.60)
(24) Printing, Publishing and Recorded Media	0.30** (19.94)	-0.18** (-14.35)	-0.22** (-15.25)	0.96** (44.25)	0.08 (1.63)
(25) Petroleum, Coal, Chemical and Assoc. Prod. Manufacturing	0.73** (6.13)	-0.32** (-3.29)	1.01* (2.34)	-0.23* (-2.13)	0.43** (11.80)
(26) Non-Metallic Mineral Product Manufacturing	0.05 (0.90)	-0.04 (-0.69)	-0.02 (-0.41)	0.19** (3.34)	0.12** (3.17)
(27) Metal Product Manufacturing	-0.17 (-0.43)	0.12 (1.05)	0.64* (2.29)	-0.13 (-0.96)	0.28** (8.26)
(28) Machinery and Equipment Manufacturing	0.37** (3.04)	-0.02 (-0.49)	0.31** (3.74)	0.07 (1.66)	0.51** (10.27)
(29) Other Manufacturing	-0.04 (-0.74)	0.02 (0.60)	0.58** (18.07)	0.17** (7.63)	0.11* (2.29)

Estimation carried out using TSP.

** indicates significant at the 1% level for a two-tailed t test.

* indicates significant at the 5% level for a two-tailed t test.

logs, these elasticities are given by the parameters f_{d2} to f_{d5} in Equation (9).⁶ It can be seen from the results in Table 3 that average variable cost, manufacturing price and aggregate income each have a significant affect on price in more than half of the industries, suggesting the importance of including demand shift variables , as well as cost variables, into industry pricing equations. In contrast to O'Regan and Wilkinson (1997), import price is only significant in a small number of industries. One reason for this finding is the inclusion of cost and other demand shift variables into the pricing equation. However, the results are likely to underestimate the influence of import price as there will be an affect on average variable cost through the use of imported intermediate inputs in domestic production.

4. Implications and conclusions

Macroeconomic models regularly emphasise the importance of the speed of price adjustment in determining real and nominal aggregates. New Keynesian models suggest that strong quantity adjustments and slow price adjustments in response to demand shocks are key to understanding business cycles. Depending on the interactions with other sectors of the economy, the speed of price adjustment may additionally impact on the long-run inflation rate. It is also suggested that exchange rates can overshoot if domestic interest rates are out of alignment with world interest rates and if prices are sticky (for example, Dornbusch et al, 2002).

These possibilities have quite rightly become the concern of Australian policy makers, thus making the determinants of the speed of price adjustment more than just an academic question. Evidence of exchange rate overshooting occurred in 2001-02

⁶ Pesaran and Shin (1999) indicate that valid inferences on the long-run parameters can be made using standard normal asymptotic theory when the variables are I(1) and the model is of the single equation ARDL type.

when the nominal exchange rate fell sharply to 0.5 United States dollars per Australian dollar, only to rise and steady off at around 0.75 United States dollars per Australian dollar in late 2003. This happened at a time of a fall in the relative Australian/United States interest rates, suggesting an outflow of capital. A lack of adjustment in relative prices meant that real exchange rate movements followed the pattern of the nominal exchange rate, which probably contributed to the surge in exports up to and through 2001 (see Reserve Bank of Australia Chart Pack). In the context of this paper, it could be asked whether the structural features of the Australian economy played a role in restricting price adjustments and causing a larger depreciation than might have otherwise occurred. Or looked at another way, did slow price adjustment allow Australia to improve its balance of trade position for a period with only minor inflationary consequences?

The theory developed in this paper and the supportive results suggest that the speed of price adjustment is determined in part by the mean size of firms within an industry and industry concentration. Given that these variables do not change rapidly, a stable speed of price adjustment at the aggregate level is to be expected in the face of changes in monetary and fiscal policy. This expectation concords with the observations of Dwyer and Leong (2001) as discussed in the introduction.

The empirical results indicate that the mean price adjustment lag in some Australian manufacturing industries can be longer than two years, while in other industries with larger firms, such as Petroleum, Coal, Chemical and Associated Product Manufacturing (25) and Machinery and Equipment Manufacturing (28), prices tend to adjust quite quickly. At a policy level, scope for altering the aggregate speed of price adjustment is limited as firm size is largely related to the characteristics of physical capital. However, it is possible that anti-merger action by the Australian

Competition and Consumer Commission (ACCC) could influence both firm size and industry concentration in individual cases.

The explanation of the empirical results, as developed in the theoretical model, is that large firms have scale economies with regard to price adjustment costs and less cause to slow their speed of price adjustment, while firms with market power can offset price adjustment costs by slowing the speed of price adjustment. This contradicts the belief that firm size and market power move the speed of price adjustment in the same direction. Perhaps a future research direction would be to theoretically identify the role of the production lag in a speed of price adjustment model so that the statistical testing of its significance is less ad hoc.

Data Appendix

All manufacturing industry data are at the two-digit ANZSIC level, while all time series data are quarterly and for the period 1994:3 to 2002:2.

$CR4_d$ – Ratio of sales for the four largest domestic firms to the value of the domestic market in the year 1998-99. Obtained by multiplying the four-firm concentration ratio by the ratio of domestic firm sales to the value of the domestic market. These data are sourced to ABS Industry Concentration Statistics (8140.0.55.001; electronic delivery) and ABS Manufacturing Australia (8225.0; Table 4.3, 2002), respectively.

$SIZ4_d$ – An average of sales (\$billions) for the four largest domestic firms in 1998-99. Obtained by multiplying the four-firm concentration ratio by domestic firm sales and dividing by four.

$PLAG_d$ – Length of the industry production period averaged across 1994:3 to 2002:2. Applying Dixon’s (1983) formula, the production lag in quarters at a point in time is given by:

$$PLAG = \left[\frac{SM}{U} \right] + \left[\frac{2(SWIP)}{Sales(1+cd)} \right] + \left[\frac{2(SFG)}{Sales} \right]$$

where SM is the stock of materials, $SWIP$ is the stock of works in progress, SFG is the stock of finished goods, U is the usage rate of materials, $Sales$ is the value of final sales per quarter, c is the share of materials costs in sales revenue and d is the proportion of materials costs that are incurred at the beginning of the period. Stock values are purchased directly from ABS, quarterly sales are sourced to ABS Business Indicators Australia (5676.0; Table 21), c is ‘purchases in’ divided by sales for 1995-96 (sourced to ABS Manufacturing Industry, Australia; 8221.0), U is $Sales$ multiplied by c , and d is 0.66 by assumption.

p_{dt} – Price index of articles produced by domestic manufacturing industries (sourced to ABS Producer Price Indexes; 6427.0; Table 11). In particular cases, a weighted average of subgroup price indices is taken to create two-digit price indices. Weights are the combined turnover shares for subgroups in 1993-94 (sourced to ABS Manufacturing Industry; 8221.0). The relevant subgroups are: 221-222 and 223-226; 231-232 and 233; 251-252 and 253-254 and 255-256; 271-273 and 274-276; 281-282 and 283-286.

c_{dt} – Index of domestic industry average variable cost. This index is not directly available, so it is constructed by taking a weighted average of material costs and unit labour costs. The formula used to construct an index of average variable cost is given by:

$$c_{dt} = SM_{d1} * \left[\frac{pm_{dt}}{pm_{d1}} \right] + SL_{d1} * \left[\frac{ulc_{dt}}{ulc_{d1}} \right]$$

where 1 indicates the first period, SM is the share of material costs in the value of output, SL is the share of labour costs in the value of output, pm is the price of materials and ulc is unit labour cost. In the calculation of c_{dt} , it is necessary to assume that the ratio of material inputs to output is fixed. SM_{d1} and SL_{d1} are constructed from 1995-96 data sourced to ABS Manufacturing Industry (8221.0). The price of material inputs is taken from ABS Producer Price Indexes (6427.0; Table 14). As with p_{dt} , it is necessary to take weighted averages across particular subgroups. These subgroups are: 221-222 and 223-224 and 225 and 226; 231-232 and 233; 251-252 and 253-254 and 255-256; 271-273 and 274-276; 281-282 and 283-286. Quarterly unit labour costs are calculated as 13 multiplied by average weekly total time earnings for each quarter (purchased directly from ABS) multiplied by quarterly total employment (ABS Labour Force Australia; 6291.0; Table 9I) divided by a quarterly chain volume measure of gross value added (ABS Australian National Accounts; 5206.0; Table 57).

imp_{dt} – Price index of manufactured imports (sourced to ABS International Trade Price Indexes, Australia; 6457.0; Table 15). In order to get two-digit import price indices, weighted averages of the following subgroups are taken: 221-222 and 223-226; 271-273 and 274-276; 281-282 and 283-286. Weights are import value shares for 1993-94 taken from the International Economic Database (IEDB), Australian National University.

mp_t – Price index of articles produced by manufacturing industries (sourced to ABS Producer Price Indexes; 6427.0; Table 10).

y_t – Chain volume measure of GDP taken from ABS Australian National Accounts (5206.0; Table 57).

Appendix 1

Deriving the industry target price when firms have log-linear demand functions and linear demand functions. Let the i^{th} firm have the following log-linear demand function:

$$\log q_{it} = A_{it} + \mathbf{h}_i \log p_{it} \quad (\text{A1})$$

where A_{it} is a function of demand shift variables and \mathbf{h}_i is the exogenous elasticity of demand. Rearranging the first-order condition as presented in (2), the firm's target price can be written as:

$$p_{it}^* = (1 + 1/\mathbf{h}_i)^{-1} mc_{it} \quad (\text{A2})$$

It can be seen from (A2) that the target price is a linear function of marginal cost at the firm level. Taking a weighted average across firms, the industry target price can be written as:

$$p_{dt}^* = e_d mc_{dt} \quad (\text{A3})$$

where $p_{dt}^* = \sum w_i p_{it}^*$, $mc_{dt} = \sum w_i mc_{it}^*$ and $e_d = \frac{\sum w_i mc_{it} (1 + 1/\mathbf{h}_i)^{-1}}{\sum w_i mc_{it}}$. Given that

the elasticity of demand and marginal cost are uncorrelated in this case,

$E(e_d) = \sum (1 + 1/\mathbf{h}_i)^{-1} / N$ where E is the expectations operator and N is the number of firms in the industry.

Now let the i^{th} firm have the following linear demand function:

$$q_{it} = A_{it} - b_i p_{it} \quad (\text{A4})$$

where b_i is a parameter that incorporates the firm's price conjectural variations with regard to other firms in the industry (for the impact of price conjectures entering in this manner on the partial adjustment model, see Olive, 2004). The firm's target price is obtained by substituting (A4) into the first-order condition and rearranging to give:

$$p_{it}^* = \frac{A_{it}}{2b_i} + \frac{mc_{it}}{2} \quad (\text{A5})$$

Equation (A5) represents the target price as a linear function of marginal cost and the demand shift variables. Taking a weighted average across firms, the industry target price can be written:

$$p_{dt}^* = \frac{b_d^{-1} A_{dt}}{2} + \frac{mc_{dt}}{2}$$

where $b_d^{-1} = \sum w_i b_i^{-1}$ and $A_{dt} = \frac{\sum w_i b_i^{-1} A_{it}}{\sum w_i b_i^{-1}}$. Therefore, the industry target price is a

linear function of the average influence of demand shift variables and marginal cost on firm target price.

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