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MANUFACTURING

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ABSTRACT

A model of industry speed of price adjustment is derived from firm pricing behaviours. The model is applied to quarterly two-digit Australian manufacturing data for the period 1985:3 to 2002:3. The results suggest that the industry speed of price adjustment is positively related to the average size of large firms within the industry and is negatively related to industry concentration. We also find that import share has a role in attenuating the effects of industry concentration and that growth in a moving average of real GDP reduces the speed of price adjustment. Calculated industry speeds of price adjustment are stable across the period of examination and are also small, suggesting that manufacturing prices are sticky.

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1. Introduction

An outstanding recent feature of many economies has been the dramatic slow down in inflation. To underline the importance of this development, Taylor (2000, p1389) says “examining the causes and consequences of the period of price stability that began in many countries in the 1980s or early 1990s is as useful for future policy making as examining the Great Depression of the 1930s or the Great Inflation of the late 1960s and 1970s”. Australia has been no exception to this change in pricing behaviour, with average annual total manufacturing price change at: 10.4 percent for the period 1970-80; 7.0 percent for the period 1981-90; and 1.8 percent in the period 1991-2001 (Australian Bureau of Statistics, 6437.0, Table 10).

There are various explanations for the recent benign inflationary outcome, which range from structural change to changes in firm expectations to favourable shocks to prices (see Dwyer and Leong, 2001; Taylor, 2000). An approach taken by Dwyer and Leong (2001) is to determine whether evidence of a structural break in the speed of price adjustment for Australian consumer prices exists. Based on recursive point estimates, they suggest that the speed of price adjustment has been slow and fairly stable through the 1990s.

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 prices are sticky.

In this paper, we are concerned with the speed and stability of price adjustment to its long-run value. Our approach is to estimate the determinants of the speed of price adjustment in Australian manufacturing industries and see whether these have shifted over time. This contrasts with the common practice in previous studies of estimating the industry speed of price adjustment as a function of structural variables that are deemed not to vary across time, most particularly industry concentration (for example see Domberger (1983); Dixon (1983); Bedrossian and Moschos (1988); Kardasz and Stollery (1988); Weiss (1993); Shannan and Feinberg (1995). An exception is Kraft (1995) who estimates the industry speed of price adjustment as a function of a dichotomous cycle variable, *inter alia*.

Martin (1993) proposes a pricing model that incorporates quadratic price adjustment costs into the profit equation and derives the speed of price adjustment as a negative function of market power. Implicit in Martin's formulation is that the speed of price adjustment is also a function of firm size, although he does not explore this issue. Indeed, there are only isolated references in the literature that posit a relationship between firm size and the speed of price adjustment. In a recent study of Swedish firms, Apel et al (2005) find that the number of price changes per year is positively related to firm size but negatively related to industry concentration. It is also suggested by Domberger (1983) that large firms have large profit cushions and, as a result, are less risk averse, leading to faster speeds of price adjustment. A natural extension of this argument is that firms have economies of scale over certain implicit adjustment costs.

In Section 2, we present a model of firm behaviour in which the speed of price adjustment is a function of market power and firm size. Averaging across all firms in an industry transforms the pricing 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.

In Section 3, the model is applied to quarterly two-digit Australian manufacturing data for the period 1985:3 to 2002:3. Initially, the industry speed of price adjustment is treated as a function of structural cross-section variables. The results suggest that the industry speed of price adjustment is positively related to the average size of large firms within the industry and is negatively related to industry concentration. With the inclusion of time variables into the final estimating equation, we find that import share has a role in attenuating the effects of industry concentration and that growth in a moving average of real GDP reduces the speed of price adjustment. A dummy variable for the 1990s is included in the final estimating equation and is found to be significant. However, its magnitude is too small to be of economic importance. Calculated industry speeds of price adjustment are stable across the period of examination and are also small, suggesting that manufacturing prices are sticky.

In Section 4, we discuss the results and conclude with some comments on their implications for economic policy.

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:

$$\pi(P_{it}) = (P_{it} - MC_{it})Q_{it} - \alpha_i \left(\frac{P_{it} - P_{it-1}}{P_{it-1}} \right)^2 (Q_{it}^*)^S \quad (1)$$

where i and t represent firm and time subscripts, respectively, and P_{it} , Q_{it} , Q_{it}^* , MC_{it} , α_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.² 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).

Another reason given for adjustment costs is slow adjustment of input quantities. Until input quantities are at their long-run equilibrium levels, the short-run profit-maximising price deviates from the corresponding long-run price. Sluggish adjustment of labour is pointed to by many authors (see Kraft, 1995; Kasa, 1998; and Lindbeck and Snower, 2001). 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.

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

² Quadratic price adjustment costs are often contrasted with menu costs. Where the former are affected by the extent of a price change, the latter are fixed regardless of the size of the price change.

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 large multinational company as for a local artisan (given the same α_i). This only makes sense if there are extreme economies of scale. In order to allow for varying scale effects, we posit that 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:

$$Q_{it}^* + (P_{it}^* - MC_{it}) \left(\frac{dQ_{it}^*}{dP_{it}^*} \right) = 0 \quad (2)$$

where * indicates the static 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 e.g. Roberts, 1992; Martin, 1993; Blinder et al, 1998, p226). Given that the actual price and the target price differ, firm output can be approximated using the following first-order Taylor series:

$$Q_{it} \approx Q_{it}^* + \left(\frac{dQ_{it}^*}{dP_{it}^*} \right) (P_{it} - P_{it}^*) \quad (3)$$

Substituting (3) into (1) explicitly expresses profit as a function of price. After deriving the first-order profit maximising condition from (1) and imposing (2), the firm chooses to change prices according to the following model:

$$\Delta P_{it} = \lambda_{it} (P_{it}^* - P_{it-1}) \quad (4)$$

$$\lambda_{it} = [1 - (\frac{\alpha_i \beta_{it}}{\eta_{it} P_{it}^* Q_{it}^{*1-S}})]^{-1} \quad (5)$$

$$\eta_{it} = \frac{P_{it}^*}{Q_{it}^*} \frac{dQ_{it}^*}{dP_{it}^*} \quad (6)$$

where $\Delta P_{it} = P_{it} - P_{it-1}$, λ_{it} is the speed of price adjustment, η_{it} is the elasticity of demand and $\beta_{it} = (P_{it}^* / P_{it-1})^2$. It is readily apparent that the range of λ_{it} is from zero to one and that (4) is a 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. Also, the speed of price adjustment increases with firm revenue when S is zero. Finally, as demand becomes more/less elastic the firm's speed of price adjustment increases/decreases.

It should be noted from (4) that the determinants of the speed of price adjustment are approximately the same for prices expressed in terms of levels or natural logarithms when price changes are relatively small. This is because

$$\Delta P_{it} / P_{it-1} \approx \ln P_{it} - \ln P_{it-1} \text{ and } (P_{it}^* - P_{it-1}) / P_{it-1} \approx \ln P_{it}^* - \ln P_{it-1} \text{ for small price}$$

changes. An examination of the data to be used in the empirical section suggests that this is a reasonable approximation. Dividing both sides of (4) by P_{it-1} and expressing the firm's pricing equation in natural logarithms gives:

$$\Delta p_{it} = \lambda_{it} (p_{it}^* - p_{it-1}) \quad (7)$$

where the lower case letters indicate natural logarithms.

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:

$$\Delta p_{dt} = \gamma_{dt} \Delta p_{dt}^* - \lambda_{dt} (p_{dt-1} - \delta_{dt} p_{dt-1}^*) \quad (8)$$

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

$$p_{dt-1} = \sum w_i p_{it-1}, p_{dt-1}^* = \sum w_i p_{it-1}^*, \gamma_{dt} = \left(\frac{\sum w_i p_{it}^* \lambda_{it}}{\sum w_i p_{it}^*} \right), \lambda_{dt} = \left(\frac{\sum w_i p_{it-1} \lambda_{it}}{\sum w_i p_{it-1}} \right) \text{ and}$$

$\delta_{dt} = \gamma_{dt} / \lambda_{dt}$. Here, w_i represents the i^{th} firm's share of the value of industry shipments at a point in time, so that the industry prices and target prices given in (8) are share-weighted averages. The error-correction form of the model comes about because γ_{dt} and the industry speed of price adjustment (λ_{dt}) are differently weighted averages of each firm's speed of price adjustment.³ If all firms in the industry have the same speed of price adjustment, then δ_{dt} is equal to one and the industry model reverts to the partial adjustment form.

3. Data and empirical model

In this section we examine the determinants and stability of the industry speed of

³ This method for obtaining an error correction model could be contrasted with those outlined by Nickell (1985).

price adjustment for nine Australian manufacturing industries at the two-digit level during the period 1985:3 to 2002:3. Two-digit industries are examined because many quarterly data series employed in this study do not exist at a finer level of aggregation. Further, the Australian Bureau of Statistics (ABS) switched to the Australian and New Zealand Standard Industry Classification (ANZSIC) in the early 1990s, making it difficult to extend particular data series beyond this time period. For this reason, total manufacturing average weekly earnings is employed as a proxy for industry average weekly earnings in what follows. See the Data Appendix for a full description of the data series and sources.

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 significantly greater than zero 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 output is influential in highly concentrated industries with low exposure to foreign competition.

These studies suggest that industry price is likely to be a function of both demand and cost variables. In this study, industry target price is modelled as a natural log-linear function of industry materials price (mp_{dt}), average weekly earnings (w_t), import price (imp_{dt}), aggregate manufacturing price (pm_t) and a moving average of real GDP (y_t) [the d subscript is dropped where the variable is common to all industry target prices].

Given (5) and (8), the industry speed of price adjustment is modelled as a function of average firm size and variables that are likely to affect the industry elasticity of demand. In the first instance, we look at the influence on the industry speed of price adjustment of structural variables that change across industry, but not (rapidly) across time.

Numerous studies in the structure-conduct-performance tradition find statistically positive relationships between industry concentration and price-cost margins (for a review, see Lipczynski et al, 2005). 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. Following this reasoning, it is expected that an increase in industry concentration reduces the industry speed of price adjustment.

Previous studies suggest that Australian industry concentration changes only slowly over time. Dixon (1987) shows that the four-firm concentration ratio averages a 4.3 percent change across 101 Australian manufacturing industries between 1968 and 1982, while Bhattacharya and Bloch (2000) find that industry concentration adjusts by 10 percent per year toward its long-run equilibrium value. The four-firm concentration ratio is only available from ABS in selected years. With slow adjustment, however, using industry values for one year (1993-94) should not greatly affect the empirical analysis.

Given economies of scale with regard to price adjustment costs, it is expected that average firm output positively influences industry speed of price adjustment. Here, turnover averaged across the four largest firms within an industry for 1993-94 is used to represent average firm output. This seems reasonable, as the largest firms are likely to have the greatest weights in (8). Firm size is also treated as a cross-sectional

variable due to the lack of data availability. However, as large firm size is generally highly correlated with concentration within an industry, but not necessarily across industries, this seems a reasonable approximation.⁴

Given the error-correction model developed in Section 2, the basic empirical pricing equation for each industry is of the form:

$$\Delta p_{dt} = \theta_{d1} + \theta_{d2} \Delta mp_{dt} + \theta_{d3} \Delta w_t + \theta_{d4} \Delta imp_{dt} + \theta_{d5} \Delta pm_t + \theta_{d6} \Delta y_t - \lambda_d ECM_{dt-1} + \varepsilon_{dt} \quad (9)$$

$$ECM_{dt-1} = p_{dt-1} - \phi_{d2} mp_{dt-1} - \phi_{d3} w_{t-1} - \phi_{d4} imp_{dt-1} - \phi_{d5} pm_{t-1} - \phi_{d6} y_{t-1} \quad (10)$$

$$\lambda_d = (\ell_1 + \ell_2 CR4_d + \ell_3 SIZ4_d) \quad (11)$$

where Δ indicates first difference, θ_{d1} to θ_{d6} and ϕ_{d2} to ϕ_{d6} and ℓ_1 to ℓ_3 are parameters, ECM_{dt-1} is the error-correction mechanism, and ε_{dt} is an error term.⁵

Initially, the industry speed of price adjustment (λ_d) is modelled as a function of the four firm concentration ratio ($CR4_d$) and the average firm turnover of the four largest firms within each industry ($SIZ4_d$). The parameters ℓ_1 to ℓ_3 are constrained to be the same for each industry, so that λ_d varies across industries but not across time.

⁴ The authors correlate the real turnovers for the largest four firms with the four-firm concentration ratio for 12 two-digit ASIC industries between 1968-69 and 1991-92. In nine industries the correlation value is over 0.75 and for seven industries it is over 0.90. However, this contrasts with the cross-sectional correlation for 1993-94 for two-digit ANSIC data, which is only 0.42.

⁵ The error term in (9) can be justified in a number of ways. At the level of the firm, a stochastic error term could be added to (3) to suggest some uncertainty in demand once the price is set. Alternatively, the change in profit with respect to a marginal change in firm price could be stochastic with an expected value of zero if firms are uncertain as to the profit maximising price. The outcome in both of these cases is that the term $v_{it} f(\alpha_i, Q_{it}^*, P_{it-1}, dQ_{it}^* / dP_{it}^*)$ is added to (7), where v_{it} is an independent random variable that has an expected value of zero. This suggests that the error term in (9) has an expected value of zero and is independent of other right hand side variables but that heterogeneity exists.

In order to determine whether our model is statistically adequate, 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 the natural logarithm of domestic industry price, industry materials price and import price each have a unit root in levels but are first-difference stationary, while the weighted symmetric tau test (Pantula et al, 1994) for single time-series indicates that the natural logarithm of average weekly earnings, manufacturing price and the moving average of

Table 1

Tests for non-stationarity of series in natural logarithm form

Variable	Level	First Difference	Test Type
p_{dt}	1.19	-6.61**	IPS
mp_{dt}	-2.28 *	-11.28**	IPS
imp_{dt}	-1.16	-15.25**	IPS
w_t	0.81 [7]	-5.68** [2]	WS
pm_t	0.46 [5]	-2.59* [3]	WS
y_t	-0.01 [5]	-2.89* [5]	WS
cointegration	-2.52**		Group ADF

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 in single time series.

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 a maximum eight lags and no time trend. WS tests have no time trend and the lag lengths are determined by the Akaike Information Criteria, with the lag lengths shown in brackets.

** indicates significant at the 1 percent level for a one-tailed test.

* indicates significant at the 5 percent level for a one-tailed test.

GDP also have unit roots in levels but are first-difference stationary.⁶ Using Pedroni's (1999) group ADF test for panel data, the null of no cointegration is rejected at the 1 percent level of significance.⁷ These results suggest that the error-correction mechanism and the error term from (8) are both stationary, so that inferences resulting from estimation of the model are not spurious.

4. Results

The model is estimated using non-linear seemingly unrelated regression (SUR), while constraining the coefficients in the industry speed of price adjustment to be the same for all industries. Although this is a one-stage procedure, pre-testing is carried out in order to determine the appropriate short-run dynamics for the regression. When the short-run dynamics are excluded, Patterson (2000) shows that long-run parameter estimates may be biased for finite series and Kremers et al (1992) show that hypothesis tests of the speed of adjustment are likely to have low power. Therefore, the error-correction model is estimated individually for each industry by non-linear least squares, allowing for up to two lags in the short-run difference variables. These lags are successively eliminated when the estimated coefficients are insignificant based on t and F tests, until the final short-run dynamic structure is obtained for each industry.⁸

⁶ The IPS test of industry materials price in levels rejects a unit root at the 5 percent level of significance. However, Levin and Lin tests do not reject the existence of a unit root in levels and the IPS test of industry materials price in first differences strongly rejects a unit root. Therefore, industry materials price is taken to be first-difference stationary.

⁷ Note that this test is tentative as it does not include the cross-industry constraints of later estimation.

⁸ The details can be obtained from the authors.

Table 2 presents the regression results for (11).⁹ For comparison, an industry and time invariant speed of price adjustment is also estimated. The t-statistics presented in this table are computed from heteroscedastic-consistent standard errors. It can be seen that the estimate of the speed of price adjustment in column (1) is 0.17 and significantly different from zero. Kremers et al (1992) find that the distribution for the speed of adjustment in a single equation is somewhere between normal and Dickey-Fuller. Given the large t-statistic (8.16), the inference is the same for both distributions (the asymptotic Dickey-Fuller critical value at the 5 percent level is 4.98).

Also, the results provide support for the theoretical model when firms have economies of scale with regard to price adjustment costs. It is shown in Table 2 that $SIZ4_d$ is significantly positive at the 10 percent level, while $CR4_d$ is significantly negative at the 1 percent level. This latter result conforms to the findings in Dixon (1983), Bedrossian and Moschos (1988), Weiss (1993) and Shaanan and Feinberg (1995).

To get a feel for the impact of firm size and industry concentration on the industry speed of price adjustment we can multiply actual values by the coefficient estimates. Metal Product Manufacturing (28) has the largest average turnover for the biggest four firms at \$3.1 billion in 1993-94 and this increases the industry speed of price adjustment by 0.12. Non-Metallic Mineral Product Manufacturing (26) has the highest industry concentration at 0.53 and this reduces the speed of price adjustment by 0.22.

⁹ The concise presentation of the estimates of the parameters in (9) and (10) and the diagnostics from each industry equation is not possible and, therefore, these are omitted from the paper. However, these results can be obtained from the authors upon request. It may be noted that the Durbin-Watson test for

Table 2

SUR estimation results for the industry speed of price adjustment (λ_d) when it is treated as a constant and as a function of cross-sectional variables

Variable	(1)	(2)
Constant	0.17** (8.16)	0.21** (4.47)
$CR4_d$		-0.42** (-2.66)
$SIZ4_d$		0.04# (1.91)

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

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

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

indicates significant at the 10 percent level for a two-tailed t test.

Thus far we have assumed that the industry speed of price adjustment is a function of cross-sectional variables. However, structural variables that change across industry and time, and cyclical variables that change across time, may also affect the industry speed of price adjustment through their influence on the elasticity of demand.

We treat import share (M_{dt}) as a structural variable that varies across time.¹⁰ A number of studies have estimated the impact of import share on the speed of price adjustment (for example, see Dixon, 1983; Kraft, 1995; Shannan and Feinberg, 1995). Dixon (1983) suggests that domestic industry concentration needs to be modified with the inclusion of import share in the estimating equation in order to represent the true degree of concentration with foreign competitors in the market. In a formal model

each of the nine industries is between 1.63 and 2.41 in the final formulation of the empirical model (shown in Table 3).

¹⁰ The recent opening up of the Australian economy has seen rapid import penetration in a number of manufacturing industries. In the 10 years from 1990 to 2000, import penetration increased at an average annual rate of 4.1 percent for All Manufacturing, 5.1 percent for Textile, Clothing, Footwear and Leather, and 7.2 percent for Metal Products (Productivity Commission, 2003).

with small country assumptions, Bloch (1994) shows that higher import shares reduce the elasticity of price conjectures, thus making demand more elastic. In light of our model, we therefore expect higher import share to increase the speed of price adjustment.

The influence of the business cycle on the elasticity of demand at an industry level is ambiguous. In boom times, increased levels of demand may increase market power for individual firms. However, anticipation of larger profits may lead to a reduction in co-operation amongst existing firms and to other firms entering the market, resulting in more elastic demand. Therefore, the impact of the cycle on industry speed of price adjustment is to be empirically determined. The cycle variable employed here is the growth in the moving average of real GDP (TG_t)¹¹.

The impact of import share and the cycle on the industry speed of price adjustment are estimated in turn and the results are shown in Tables 3. In order to indicate if there has been a structural change in the industry speed of price adjustment emanating from an unknown source, the impact of a dummy variable (*Dummy*) that is zero from 1985:3 to 1990:4 and one from 1991:1 to 2002:3 is also estimated. In regressions (1), (2) and (3), *Dummy*, M_{dt} and TG_t are each insignificantly different from zero. However, when these variables are all included together in regression (4), *Dummy* is significantly positive at the one percent level and TG_t is significantly negative at the five percent level, while M_{dt} is insignificantly different from zero. These results suggest that there is an unexplained speeding up of industry price adjustment in the 1990s via the dummy variable, but that the magnitude of the change is too small (0.003) to be economically relevant. Also, the negative sign on TG_t is

¹¹ TG_t is calculated as $y_t - y_{t-1}$. Rosenbaum and Sukharomana (2001) find that a cycle dummy based on a deterministic trend is negatively related to market power in the US Portland cement industry.

consistent with higher growth levels leading to greater market power for firms and this slowing industry speed of price adjustment.

Table 3

SUR estimation results for the industry speed of price adjustment when it is a function of cross-sectional and time variables.

Variable	(1)	(2)	(3)	(4)	(5)
Constant	0.21** (4.57)	0.26** (5.31)	0.24** (5.35)	0.26** (6.85)	0.29** (7.49)
$CR4_d$	-0.46** (-4.52)	-0.49** (-3.27)	-0.44** (-4.60)	-0.53** (-10.17)	-0.59** (-9.44)
$SIZ4_d$	0.05* (2.28)	0.03 (1.32)	0.04 (1.95)	0.03* (2.01)	0.03* (2.07)
<i>Dummy</i>	-0.001 (-0.91)			0.003** (4.44)	0.003** (3.92)
M_{dt}		0.02 (1.63)		-0.001 (-0.28)	-0.10** (-2.86)
TG_t			-2.36 (-1.59)	-1.86* (-2.30)	-2.05* (-2.40)
$M_{dt} * CR4_d$					0.25** (2.74)

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

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

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

indicates significant at the 10 percent level for a two-tailed t test.

In order to capture the degree to which industry concentration is modified by openness to import competition, industry concentration multiplied by import share ($M_{dt} * CR4_d$) and is added to regression (5). While all other coefficient estimates are of a similar value and significance to regression (4), M_{dt} is negative and significantly different to zero and $M_{dt} * CR4_d$ is positive and significantly different to zero. These results imply that import share reduces the impact of industry concentration on the

industry speed of price adjustment when the level of industry concentration is higher than 0.4. However, import share slows the industry speed of price adjustment when industry concentration is lower than 0.4. One explanation for this result is that import competition may disrupt firm price adjustment toward their target levels in low concentration industries (see Shannan and Feinberg, 1995).

It can be seen from Table 3 that the inclusion of time variables only has a minor effect on $SIZ4_d$ (ranging from 0.03 to 0.05), $CR4_d$ (ranging from -0.44 to -0.59) and the constant (ranging from 0.21 to 0.29), although $CR4_d$ and the constant are clearly higher for regression (5). These results confirm the finding from Table 2 that industry concentration and the size of firms within an industry are important determinants of the industry speed of price adjustment.

The implied speed of price adjustment for an industry in any one time period is obtained by multiplying the coefficient estimates for regression (5) in Table 3 by the relevant industry variable values and summing. Table 4 shows the implied speed of price adjustment for each industry averaged over the whole period, averaged over the period 1985:3 to 1990:4 and averaged over the period 1991:1 to 2002:3. It is obvious that the average speed of price adjustment in 1985:3 to 1990:4 is virtually the same as in 1991:1 to 2002:3 for every industry. This supports Dwyer and Leong's (2001) finding at the aggregate level, that the 1990s slow down in inflation has not been caused by a shift in the speed at which firms adjust their prices.

Table 4 also shows the maximum and minimum implied speed of price adjustment for each industry. The range of values from minimum to maximum is either 0.03 or 0.04. Such a small range suggests that the industry speed of price adjustment is stable over the sample period. Greater variation in the industry speed of price adjustment is found across industries, with the maximum and minimum average

values across the period being, respectively, 0.23 for (21) Food, Beverage and Tobacco and 0.002 for (26) Non-Metallic Mineral Products.

Table 4

Implied speed of price adjustment for nine ANZSIC two-digit industries calculated from the regression (5) results presented in Table 3.

Industry	Industry Speed of Price Adjustment				
	Average (1985:3- 2002:3	Average (1985:3- 1990:4	Average (1991:1- 2002:3	Maximum	Minimum
(21) Food, Beverage and Tobacco	0.23	0.23	0.23	0.26	0.22
(22) Textile, Clothing, Footwear and Leather	0.17	0.18	0.17	0.20	0.16
(23) Wood and Paper Products	0.09	0.09	0.09	0.11	0.07
(24) Printing, Publishing and Recorded Media	0.14	0.14	0.14	0.17	0.13
(25) Petroleum, Coal, Chemical and Associated Products	0.16	0.16	0.16	0.18	0.14
(26) Non-Metallic Mineral Products	0.002	0.001	0.002	0.02	-0.02
(27) Metal Products	0.16	0.16	0.16	0.18	0.14
(28) Machinery and Equipment	0.18	0.18	0.18	0.20	0.17
(29) Other Manufacturing	0.21	0.21	0.21	0.23	0.19

The average value of the speed of price adjustment across all industries is 0.15 (calculated from Table 4), which is slow when compared to a number of overseas industry studies (for example, see Bedrossian and Moschos (1988) for Greece (0.30 on average) and Kardasz and Stollery (1988) for Canada (0.59 on average)). Using the median lag formula $[\ln(0.5)/\ln(1 - \lambda_d)]$, this equates to half of the impact in industry price from a disturbance to the long-run relationship taking 4.3 quarters to work itself

out on average.¹² However, Dwyer and Leong's (2001) estimate for the speed of price adjustment for Australian consumer prices is 0.07, which suggests an even longer median lag of 9.6 quarters (similar estimates for Australian consumer prices are obtained by Stone et al; 2005). This difference may be due to the level of aggregation and differences between manufacturing and other sectors of the economy.

5. Implications and conclusions

The empirical results presented in this paper suggest that the industry speed of price adjustment to its long run level is quite slow and is determined in the most part by structural variables that do not change rapidly over time. If these findings can be extended to other sectors of the economy, then policy makers can be reasonably confident that the aggregate speed of price adjustment will remain slow and stable in the face of monetary, fiscal and external shocks.

The results also suggest that industry concentration reduces the industry speed of price adjustment, while it is increased by average firm size. An intuitive explanation 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 may hesitate to adjust prices due to the uncertainty regarding the reactions of rivals and customers. Finally, the results also imply that there is little evidence of structural change to industry speeds of price adjustment at the beginning of the 1990s and that while the business cycle has an influence that is measurable, it is small.

¹² This median lag formula is appropriate for the partial adjustment model, so the values derived here are approximations of the true median lag.

Data Appendix

All manufacturing industry data are at the two-digit ANZSIC level, while all time series data are quarterly and for the period 1985:3 to 2002:3. With the exception of TG_t , all lower case variables are in natural logarithms and all upper case variables are in levels.

p_{dt} – Price index of articles produced by domestic manufacturing industries (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.

mp_{dt} – Price index of material inputs by domestic manufacturing industries (ABS Producer Price Indexes; 6427.0, Table 14 [and 6411.0 prior to 1989:3 for industries 23, 24 and 25]). In particular cases, a weighted average of subgroup price indices is taken to create two-digit price indices.

imp_{dt} – Price index of manufactured imports (ABS International Trade Price Indexes, Australia; 6457.0, Table 15 [and 6414.0 prior to 1991:2 for industries 22, 23, 24, 25, 28 and 29]). In particular cases, an import weighted average of subgroup import price indices is taken to create two-digit import price indices.

w_t – Average weekly total earnings for total persons in manufacturing (ABS Average Weekly Earnings; 6302.0, Table 10i).

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

y_t – Trend chain volume measure of GDP (ABS Australian National Accounts; 5206.0, Table 14). This is an ABS calculated moving average of the chain volume measure of GDP.

$CR4_d$ – Ratio of turnover for the four largest domestic firms to the value of total industry turnover in the year 1993-94 (ABS Manufacturing Industry, 8221.0).

$SIZ4_d$ – Average turnover for the four largest domestic firms within the industry (\$billions) in the year 1993-94 (ABS Manufacturing Industry, 8221.0).

M_{dt} – Industry import share of apparent domestic consumption. Imports sourced from ABS 5368.0 (Table 17), exports are sourced from ABS 5368.0 (Table 14) and sales are taken from ABS 5676.0 (Table 21). Prior to 1989, import share is obtained from *Manufacturing industry and international trade data, 1968-69 to 1992-93*, Industry Commission (1995). Annual import share data are converted into a quarterly series by assuming that import share is constant for each year.

TG_t –Growth rate in the moving average of the chain volume measure of GDP (ABS Australian National Accounts; 5206.0, Table 14). This is obtained by first differencing the natural logarithm of the moving average of a chain volume measure of GDP.

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