How much do import price shocks matter for consumer prices?

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November 2001

Discussion Paper Series

DP2001/06

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Abstract[1]

New Zealand and Australia have recently had large import price increases on the back of currency depreciations. Despite these large import price shocks, consumer prices have not risen as strongly as would be suggested by previously accepted relationships. The recent prolonged period of general US dollar strength has also raised the profile of exchange rate and import price pass-through in a number of other countries.

We re-estimate short- and long-run import price pass-through coefficients for New Zealand consumer prices. The data confirms that the estimated coefficients are smaller than previously thought: a 10 per cent increase in import prices will lead to a 0.5 per cent increase in consumer prices within the first quarter, and to a 1.5 per cent increase over the long-run.

This result suggests that we should not necessarily expect the recent exchange rate depreciation to result in further consumer price increases. Similarly, if the exchange rate were to appreciate in the future, we should not expect large decreases in consumer prices as a direct result of the exchange rate appreciation.

1   Introduction

The degree and timing of pass-through is important for forecasting inflation and for setting monetary policy in response to inflation shocks. The term *long-run pass-through* typically refers to the *total* change in consumption prices generated by a persistent or permanent import price shock, once consumer prices have stopped adjusting to the initial import price shock. It has been suggested that the decline in the Australasian currencies over 1999 and 2000 may represent a permanent or structural down-shift, rather than just a cyclical phenomenon. If this is the case, and the exchange rate never recovers to previous cyclical highs, then understanding long-run pass-through becomes very relevant.

Irrespective of the long-run relationship, it takes time for domestic consumer prices to adjust to any new equilibrium, and understanding the short-run pass-through dynamics is important for two reasons. First, the short-run dynamics explain the adjustment process to achieving the long-run pass-through for a given permanent shock to import prices. Second, understanding the short-run pass-through dynamics helps to tie down the impact of temporary import price shocks on consumer prices.

The price of imports over the docks increased by around 25 per cent in New Zealand and by 11 per cent in Australia in the two years to the end of 2000. The import price increases were driven predominantly by the large depreciation in the Australasian exchange rates over that period – with import prices broadly matching the exchange rate movements one-for-one within one quarter. Over the same period, consumer prices in New Zealand increased by just 5 per cent. This increase in consumer prices was not as large or as quick as previously-accepted relationships would have suggested (RBNZ (2001)).

The recent period of prolonged general US dollar strength has increased the profile of exchange rate and import price pass-through in a number of countries outside of Australia and New Zealand.

The objective of this research is to identify the long-run relationship between import prices and consumer prices, and to identify the short-

run dynamics involved. The theory behind import price pass-through and a summary of some of the earlier empirical literature is discussed in section 2. Section 3 presents the empirical models used in this paper, along with the results. Section 4 concludes.

2   Theory and empirical literature

A typical starting point for identifying the theoretical long-run impact of import prices on consumer prices is to assume that the consumption price elasticity with respect to import prices (the pass-through coefficient) is equal to the equilibrium proportion of consumption goods that are imported. In the case of New Zealand, imports account for approximately 33 per cent of consumption goods. Therefore, based on the above assumption, a 10 per cent increase in import prices would lead to a 3.3 per cent increase in consumer prices – in the long-run.[2]

This is an appropriate approximation provided that it is only the prices of imported consumption goods that are affected by the import price shock, and not the prices of domestically produced consumption goods. If other prices are affected by the import price shock then there are at least two – potentially offsetting – reasons why the pass-through coefficient may not be equal to the import share of consumption.

First, a movement in import prices may lead to a change in consumer prices that is larger than that suggested by the import share of consumption. For example, a positive shock to oil prices, and therefore petrol prices, may lead to higher wage claims if some wages are indexed to consumer prices, which in turn may feed back into higher domestic inflation. This phenomenon, whereby higher import prices lead to higher prices for domestically produced goods is usually labelled *indirect* or *second round* pass-through. If second round pass-through occurs, then the total pass-through from import price changes to consumption prices could be greater than the import share of consumption. Second round pass-through is more likely to

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[2]   These examples can be similarly applied to declines in import prices. However, the examples provided in this paper are typically illustrations of the consequences of an increase in import prices.

play a significant role when the central bank is too slow to respond to import price shocks to prevent inflation persistence.

Conversely, the change in consumer prices generated by the import price shock may be *smaller* than that suggested by the import share of consumption. It is possible that the import price shock may result in a relative price change and not necessarily a levels shift in the overall consumer price index. For example, if the import price increase occurred without a compensating change in nominal income (via higher export prices), the higher price of imports may simply be offset by lower prices for other goods as demand declines – a relative price shift. Because of these relative price shifts, the effect of an import price shock on consumer prices may be smaller than the simple import share of consumption relationship suggests. This income effect is likely to be more prevalent in an economy with a credible inflation target. For example, if the price of the non-imported goods did not decline to compensate for the rising import prices, the central bank would be expected to tighten monetary policy in response to a positive import price shock.

Also, if the higher price of imported goods results in decreased consumption of imported consumption goods, then the imported share of consumption might also decline. Hence, the increase in the total consumer prices from the initial import price shock may again be smaller than the original import share of consumption, because of this substitution effect away from imported consumption goods. This substitution effect should be compounded by the competing sectors' expansion.

The offsetting forces outlined above mean that identifying the exact magnitude of the pass-through coefficient is an empirical question. Dwyer and Lam (1994) show that the sizeable domestic costs involved in the distribution and sale of imports imply that the long-run import price pass-through coefficient in Australia will be lower than the share of consumption goods that is imported. However, in a more recent study, Gruen and Leong (2001) find that the estimated long-run pass-through from import prices to domestic consumption prices in Australia is between 0.2 and 0.4. This is comparable to the import share of consumption in Australia, which is around 30 per

cent. Hence Gruen and Leong (2001) could not conclude which of the above effects dominates.

Consumer prices do not adjust to their new long-run value instantaneously following an import price shock – these things take time. Firms will usually alter their margins and absorb some of the import price shock – at least temporarily – until they get a better understanding of the nature of the import price shock or until market conditions change. This partial pass-through reflects the facts that it is typically costly for firms to adjust prices (due to menu costs) and that they are generally unsure whether import price movements are going to be permanent or just cyclical. Firms therefore delay some, or all, of their price adjustment until they determine the duration of the shock – thus varying their margins through the cycle. A desire to maintain market share will also encourage firms to delay increasing their price in response to higher import prices until their competitors have committed to a similar price rise, again varying their margins.

Debelle and Wilkinson (2000) plot the estimated short-run pass-through coefficient from a time series of 10 year rolling regressions of a difference equation for the period 1983Q1 to 2000Q2. They show that the coefficient for New Zealand has declined from around 0.05 (a 10 per cent rise in import prices would lead to a 0.5 per cent rise in consumption prices within one quarter) for the 10 years to 1992, to around 0.02 for the 10 years to 2000. After allowing for feedback from past inflation to current inflation, Debelle and Wilkinson's full-sample estimates imply that a 10 per cent movement in New Zealand import prices leads to a 1.5 per cent movement in domestic consumer inflation within 12 months, and a 2.3 per cent movement after two years. This compares to their estimates for Australia of 0.6 and 1.0 per cent respectively, 0.2 and 0.3 per cent for Canada, and 0.8 and 0.9 per cent for the United Kingdom. Using a vector autoregression approach, McCarthy (2000) estimated the coefficients for a number of countries, and found a much larger value for the United Kingdom – around 5 per cent.[3] This suggests that the estimates of import price pass-through

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[3] McCarthy (2000) estimated the elasticity from the exchange rate to consumer prices rather than from import prices to consumer prices. However, even allowing for this difference, their estimates still suggest much larger import price pass-through than do those of Debelle and Wilkinson (2000).

may be sensitive to the estimation technique used. Debelle and Wilkinson's results suggest that the extent of short-run import price pass-through has declined in most countries since the beginning of the 1980s. However, they do not conclude whether this decline in short-run pass-through is a result of a *reduction* in long-run pass-through, or whether the pass-through has just become more *protracted*.

Taylor (2000) puts forth the view that the reported decline in pass-through or pricing power around the world is due to the low inflation environment that has been achieved in many countries. First, a microeconomic model of price setting is used to show that lower pass-through is caused by lower perceived persistence of cost changes. Evidence is then presented showing that inflation is positively correlated with inflation persistence, suggesting that the low inflation itself has caused the low pass-through.

Dwyer and Leong (2001) support the view that price stability has had a significant impact on reducing second round pass-through in Australia in the sense that inflation expectations are better anchored so unit labour costs are less affected by exchange rate shifts. However, they note that *first round* pass-through has also fallen since the advent of price stability.

Finally, import price pass-through in Australia appears to be asymmetric: the pass-through of landed import prices is slower during appreciations than during depreciations (Dwyer, Kent and Pease (1993)). In the event of depreciation it takes two years before half of the estimated pass-through is realised. In the event of appreciation, it takes even longer.

3   Empirical models and results

The theoretical basis for this paper is a mark-up framework, whereby consumer prices are represented as a mark-up over the costs of production. Assuming that the economy can be divided into traded and non-traded sectors, and that exportables are not consumed, the relevant costs are unit labour costs in the *non-traded* sector and

import prices.[4] In the long-run, therefore, we should have (up to a constant):

$$p_t^c = \gamma ulc_t^{NT} + (1-\gamma)pm'_t \qquad (1)$$

where p_t^c is logged consumer prices, ulc_t^{NT} is logged unit labour costs in the non-traded sector, pm'_t is logged import prices plus tariffs, and γ is the non-traded proportion of the economy. No official time series exists for New Zealand unit labour costs for the non-traded sector only, and hence equation (1) cannot be estimated directly. We therefore make the following data transformation to correct for this.[5]

We make the common assumption that the trend rate of labour productivity growth in the traded sector is higher than in the economy as a whole. As a consequence, unit labour costs rise more slowly in the traded sector than in the economy as a whole,

$$ulc_t^T = ulc_t - x*t \qquad (2)$$

where ulc_t is logged economy-wide unit labour costs, ulc_t^T is unit labour costs in the traded sector, t is a linear time trend, and x is the assumed difference between traded and economy-wide rates of labour productivity growth. Treating aggregate unit labour costs as a geometric average of unit labour costs in the traded and non-traded sectors yields:

$$ulc_t = \gamma ulc_t^{NT} + (1-\gamma)ulc_t^T \qquad (3)$$

and consequently, equations (2) and (3) can be combined to show that

$$\gamma ulc_t^{NT} = \gamma ulc_t + (1-\gamma)x*t \qquad (4)$$

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[4] Unit labour costs can also be thought of as a proxy for price expectations and non-imported marginal costs.

[5] This section is drawn extensively from Beechey et al. (2000).

which may be substituted into equation (1) to give

$$p_t^c = \gamma ulc_t + (1-\gamma)(pm'_t + x*t) \qquad (5)$$

or

$$p_t^c = \gamma ulc_t + (1-\gamma)p_t^m \qquad (6)$$

where

$$p_t^m = pm'_t + x*t$$

To calculate the value of x, the difference between traded and economy-wide rates of labour productivity growth, we required estimates of labour productivity growth for each of the sectors over the sample period. The tradable sector was characterised as encompassing agriculture, hunting, forestry, fishing, manufacturing, mining, and quarrying. The level of productivity for each sector was calculated as the seasonally adjusted level of employment in the sector (as measured by the Household Labour Force Survey) divided by seasonally adjusted real GDP for that sector. x was calculated as the average difference in the quarterly growth rates of economy wide and traded sector labour productivity for the period 1989Q1 to 2001Q2.

As noted in section 2, there are a number of reasons why the estimated value of γ may not equal the non-traded proportion of the economy – even in the long-run. We therefore estimate the following unrestricted equation:

$$p_t^c = \alpha + \gamma_1 p_t^m + \gamma_2 ulc_t + \delta_1 GAP_{t-3} + \delta_2 \sum_{i=0}^{2}\Delta GAP_{t-i} + \varepsilon_t \qquad (7)$$

which is very similar to the model developed in de Brouwer and Ericsson (1995). Failure to make the above adjustment to the import price series would mean that the estimated value of γ_2 would overstate its true value. The added inclusion of the output gap (GAP) and an accelerator term on the output gap in equation (7) is

consistent with the result that output gaps are useful in explaining consumer price inflation (Coe and McDermott (1997)).

A few other potential problems with estimating import price pass-through using a mark-up equation are worth noting at this point. The first is that consumer prices are not just a mark-up over production costs, but they also include indirect taxes. Second, firms' mark-ups may vary over the long-run – biasing the long-run coefficient estimates. At the end of section 3.2 we make an allowance for this potential bias, and find that it does not have a significant effect on the results. Third, there are potential problems with assuming an economy-wide mark-up, rather than trying to allow for different mark-ups at a more disaggregated level. Finally, mark-up equations are only reduced forms of the behaviour of prices in the economy. In particular, the monetary authority and inflation expectations appear only implicitly under this specification.

With these points noted, we proceed to use variants of the above mark-up model to estimate the pass-through coefficient for New Zealand import prices. First we test for unit roots in the series, before proceeding to estimate the long-run relationships.

3.1   Unit root tests

Table 1 shows the results of Augmented Dickey-Fuller unit root tests on p^c, p^m, ulc, and GAP for the period 1989Q1 to 2001Q2.[6]

The results suggest that import prices and unit labour costs are integrated of order 1 (I(1)), while they reject the null hypothesis that the output gap is non-stationary. Although the test rejects the null hypothesis of a unit root in consumer prices, a plot of the series and the correlogram suggests that the series exhibits a unit root.[7]

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[6]   This period was chosen because import price data including tariffs was not available prior to this date – see the Data Appendix for details of the tariff time series. The model was also estimated for the period 1992Q2 to 2001Q2 – the period since inflation was stabilised in New Zealand – but this did not have a significant effect on any of the results.

[7]   The null hypothesis of each of the series being I(2) was rejected for all of the series. The results are not reported here.

Taking these results in conjunction with previously held theoretical perceptions of the data properties, it was decided to treat consumer prices, import prices, and unit labour costs as I(1) series, and the output gap as a stationary series.

3.2   Cointegration levels model

This subsection uses econometric analysis to try to identify the long-run import price pass-through coefficient. Assuming that import prices, consumer prices, and unit labour costs are all I(1), it is appropriate to estimate the long-run relationship represented by equation (7) using cointegration analysis. Stock and Watson (1993) show that the cointegrating vector can be estimated by a simple ordinary-least-squares (OLS) regression augmented by lead and lag differences. Equation (7) then becomes:[8]

$$p_t^c = \alpha + \gamma_1 p_t^m + \gamma_2 ulc_t + \delta_1 GAP_{t-3} + \delta_2 \sum_{i=0}^{2} \Delta GAP_{t-i} + \sum_{s=-1}^{1} \beta_{1s} \Delta p_{t-s}^m + \sum_{s=-1}^{1} \beta_{2s} \Delta ulc_{t-s} + \varepsilon_t \qquad (8)$$

Stock and Watson (1993) explain that the advantage of estimating equation (8), as opposed to equation (7), is that the inclusion of the first difference variables helps correct for potential endogeneity from consumer prices back into unit labour costs and import prices. This dynamic OLS estimator performs well, in a small sample sense, relative to other asymptotically efficient estimators. Stock and Watson (1993) do not explicitly examine the case where exogenous stationary inputs are included, in this case the output gap terms, but such inputs can be accommodated (see Phillips (1991)).

Hamilton (1994, pp 608-612) shows that an OLS regression of equation (8) will produce incorrect standard errors, and summarises the methodology for estimating a consistent estimate of the standard error of the regression. This consistent estimate is denoted as λ.

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[8]   Additional lags and leads were included in the original regression, but were dropped due to their lack of statistical significance. The estimates for γ_1 and γ_2 were not affected by the inclusion of the additional lags.

This approach can be applied to a regression that is estimated using any asymptotically efficient estimator. Equation (8) satisfies this condition. Hamilton (1994, pp 608-612) shows that the standard OLS F statistics can be transformed into correct test statistics by multiplying them by the ratio (s^2/λ^2), where s is the unadjusted OLS standard error of the regression. Similarly, the OLS standard errors can be transformed into consistent standard errors by dividing them by the ratio $(s^2/\lambda^2)^{1/2}$. If the dynamics of ε_t (from equation (8) above) were to be approximated on the basis of q autocovariances, the Newey-West estimator can be used to obtain a consistent estimate of λ:

$$\lambda^2 = c_o + 2\sum_{j=1}^{q}[1 - j/(q+1)]c_j \qquad (9)$$

where

$$c_j = T^{-1}\sum_{t=j+1}^{T}\varepsilon_t\varepsilon_{t-j}$$

and ε_t are the residuals from equation (8) above.[9]

A desirable property of mark-up equations is static homogeneity. Static homogeneity states that an equal percentage change in unit labour costs and import prices will cause a one-for-one percentage change in consumer prices in the long-run. Static homogeneity therefore ensures that monetary policy cannot result in long-run relative price changes (eg between consumer goods and import prices). The restriction of static homogeneity prevents monetary policy from having a long-run compositional effect on the real economy via relative prices. In equation (8), static homogeneity is imposed by setting $\gamma_1 + \gamma_2 = 1$. A Wald test on this restriction was conducted, the results of which are presented in table 2. The restriction was accepted by the data, and hence the remainder of the results are for equation (8) with static homogeneity imposed. If the import price series was not adjusted for the cross-sectoral productivity differences, as described at the beginning of section 3, static homogeneity was rejected.

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9   The value of q was set equal to the number of statistically significant autocorrelation coefficients from the residuals obtained from equation (8).

Engle-Granger cointegration tests were also performed to test whether a long-run relationship exists between the level of consumer prices, import prices, and unit labour costs (in the long-run, the output gap is assumed to have a mean of zero).

Hypothesis test results, the coefficient estimates, and the adjusted standard errors from equation (8) with static homogeneity imposed are all presented in table 2. The coefficient results for the dynamic variables are not included in the table because they were only included in the regression to solve for potential endogeneity problems. However, the coefficient estimates on the output gap variables are included in the table because they are potentially of economic interest.

The results suggest that the import price pass-through coefficient is 0.14. This implies that if import prices increase by 10 per cent, then consumer prices will increase by around 1.4 per cent in the long-run.[10] This is around half of the value of the import share of consumption for New Zealand, suggesting that there is typically a lot of relative price shifts following an import price shock, and relatively little second round pass-through. This is consistent with the Reserve Bank of New Zealand operating within a credible inflation-targeting environment.

The Engle-Granger test on the residuals from equation (8) rejected the null hypothesis of no cointegration, therefore suggesting that a long-run relationship does exist between import prices, consumer prices, and unit labour costs. This finding validates the above coefficient estimates.

If firms' profit margins changed significantly over the estimation period, then excluding some measure of profit margins from the above analysis could bias the results. Hence equation (8) was re-estimated with a measure of margins (calculated from the national accounts) in levels. The coefficient on the margins variable was negative − a result that was inconsistent with the idea that firms'

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10   This methodology assumes that the impact of import price movements on consumer prices is symmetric.

profit margins and their selling price are positively correlated. Nevertheless, its inclusion in equation (8) did not have a significant impact on the other coefficients. Therefore, the margins variable was excluded from the analysis.

3.3   Error-correction mechanism

The next objective was to determine the short-run adjustment process required to achieve the long-run relationship. Having identified a stable cointegrating relationship, it is convenient to re-estimate the above relationship using an error-correction mechanism (ECM):

$$\Delta p_t^c = \mu(p_{t-1}^c - \alpha - \gamma_1 p_{t-1}^m - \gamma_2 ulc_{t-1}) + \delta_1 GAP_{t-3} +$$
$$\delta_2 \sum_{i=0}^{2} \Delta GAP_{t-i} + \sum_{s=-1}^{1} \beta_{1s} \Delta p_{t-s}^m + \sum_{s=-1}^{1} \beta_{2s} \Delta ulc_{t-s} + \varepsilon_t \qquad (10)$$

Re-estimating the above relationship using an alternative specification, such as an ECM, also helps to test the robustness of the earlier results. As in equation (8), the restriction of static homogeneity $\gamma_1 + \gamma_2 = 1$, was tested.[11] The results of this test are presented in table 3, and again the restriction was accepted. As a result, the remainder of the results are from equation (10) with the restriction of static homogeneity imposed. The regression results are shown in table 3.

The estimated coefficient on the error-correction term, μ, is around -0.37, implying that around 37 per cent of a disequilibrium in the long-run relationship is corrected each quarter. This coefficient is statistically significant, supporting the earlier result that a long-run relationship exists between import prices, consumer prices, and unit labour costs. The μ coefficient can be used to approximate how long it would take for half of any deviation from the long-run equilibrium to be eliminated. This *half-life* is calculated using:

$$halflife = \ln(0.5) / \ln(1 + \mu) \qquad (11)$$

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[11]   Additional lags and leads were included in the original regression, but were dropped due to their lack of statistical significance. The estimates for γ_1 and γ_2 were not affected by the inclusion of the additional lags.

The -0.37 μ coefficient obtained from equation (10) is equivalent to saying that the half-life is around 5 months.[12]

The γ_1 and γ_2 from the ECM are very similar to those obtained from equation (8), and suggest that the import price pass-through coefficient is 0.16 (0.14 in equation (8)). Combining the μ and γ_1 coefficients from the ECM, and ignoring the β_i coefficients, a 10 per cent increase in import prices would lead to a 0.6 per cent increase in consumer prices within one quarter.[13]

The long-run import price pass-through coefficients estimated here are a little lower than those calculated by Debelle and Wilkinson (2000), while the speed of adjustment is quicker. One possible explanation for the differences stems from the fact that Debelle and Wilkinson (2000) use headline import prices, while this paper uses import prices adjusted for tariffs and the difference in labour productivity growth between the traded and non-traded sectors. Second, the analysis done by Debelle and Wilkinson (2000) uses a simple difference equation, whereas this paper uses cointegration based estimation techniques.

4   Conclusion

The objective of this paper was to determine the relationship between import prices and consumer prices in New Zealand. Understanding both the long-run impact and the short-run dynamics of import price shocks on consumer prices is crucial for monetary authorities in small open economies. Further, the recent period of prolonged US dollar strength has raised the profile of exchange rate and import price pass-through in a number of other economies besides New Zealand.

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[12]   This approximation of the half-life does not take into account the dynamic adjustment terms represented by the β_i coefficients. However, given that these coefficients are statistically and economically small in this example, the above equation gives a close approximation.

[13]   Again, this estimation does not take into account the β_i coefficients.

The results suggest that the long-run import price pass-through coefficient for New Zealand is around 0.15, with around half of any deviation from the long-run relationship being corrected in around five months. Combining these two results suggests that a 10 per cent increase in import prices will lead to a 0.5 per cent increase in consumer prices within the first quarter, and to a 1.5 per cent increase over the long-run.

Earlier internal research done at the Reserve Bank of New Zealand (using data for the period 1985 to the mid 1990s) suggested that the long-run import price pass-through coefficient was around 0.25 to 0.30. The much lower value of 0.15 estimated in this paper is consistent with the more recent informal analysis of the Reserve Bank, which suggests that the recent depreciation in the exchange rate has not resulted in increases in consumer prices as large as we had expected (RBNZ (2001)).

This result suggests that there may not necessarily be further large increases in consumer prices arising from the recent exchange rate depreciation. Conversely, any future exchange rate appreciation should not be expected to have a large direct impact on consumer prices.

An important caveat to this analysis is that the measured value of the long-run import price pass-through coefficient also takes into account any impact of changes in monetary policy following a shock to import prices. The fact that the pass-through coefficient is significantly lower than the import share of consumption is consistent with the Reserve Bank operating a credible inflation targeting regime. The inflation targeting regime has helped to ensure that an increase in import prices is typically met with relative price shifts in consumer goods, with some domestically produced goods' prices falling to offset the rising price of imported consumer goods, and limited, if any, second round or indirect pass-through. It therefore follows that in a different regime, where the Reserve Bank was not expected to tighten policy in response to import price shocks that threatened to destabilise inflation expectations, then the pass-through coefficient would probably be higher than the 0.15 estimated here.

Table 1: Unit root tests

Test statistics (5 per cent critical values)		Lags	Constant	Trend
pd	-4.75* (-3.50)	0	Yes	Yes
pm	2.02 (-1.95)	0	No	No
ulc	-2.25 (-2.93)	0	Yes	No
GAP	-2.08* (-1.95)	0	No	No

Notes: * indicates significance at the 5 per cent level of significance. The null hypothesis of a unit root was tested using the Augmented Dickey-Fuller unit root test. Starting with eight lags, and sequentially dropping the statistically insignificant ones determined the lag length. The constant and trend variables were excluded if they were not statistically significant. The critical values were obtained from Hamilton (1994).

Table 2: Coefficient estimates and hypothesis tests for equation (8)

	Coefficient estimates (standard errors)	
γ_1	Import prices	0.14 (0.01)
γ_2	Unit labour costs	0.86 (0.01)
δ_1	*GAP*	0.41 (0.03)
δ_2	*∑Δ(GAP)*	0.26 (0.03)
Hypothesis tests (5 per cent critical values)		
$\gamma_1 + \gamma_2 = 1$		1.86 (3.84)
No cointegration		-4.98* (-3.74)

Notes: * indicates significance at the 5 per cent level of significance. The standard errors are calculated by the OLS standard errors statistics by an estimate of $(s^2/\lambda^2)^{1/2}$. The *F* statistics are calculated by multiplying the OLS *F* statistics by an estimate of (s^2/λ^2). (Hamilton (1994) p 610). A test statistic larger than the critical values (in absolute terms) rejects the listed hypothesis. All critical values are at the 5 per cent level of significance. The null hypothesis that $\gamma_1 + \gamma_2 = 1$ was tested using a Wald test. The null hypothesis of no cointegration was tested using the Engle-Granger (1987) cointegration test, with the number of lags used in the residual regression determined using sequential *t* tests. The critical values were obtained from Hamilton (1994).

Table 3: Coefficient estimates and hypothesis tests for equation (10)

	Coefficient estimates (standard errors)	
μ	Error-correction	-0.37 (0.12)
γ_1	Import prices	0.16 (0.03)
γ_2	Unit labour costs	0.84 (0.03)
δ_1	*GAP*	0.15 (0.06)
δ_2	*∑Δ(GAP)*	0.11 (0.03)
$\sum\beta_{1s}$	*Δ(Import prices)*	0.04 (0.03)
Hypothesis tests (5 per cent critical values)		
$\gamma_1 + \gamma_2 = 1$		0.10 (3.84)
No serial correlation		5.72 (9.49)
Normally distributed residuals		0.58 (5.99)
No heteroskedasticity		32.39* (18.3)

Notes * indicates significance at the 5 per cent significance level. A test statistic larger than the critical values (in absolute terms) rejects the listed hypothesis. All critical values are at the 5 per cent level of significance. The null hypothesis of no serial correlation was tested using the Breusch-Godfrey test. The lag length for this test was set to four. The null hypothesis of normal residuals was tested using the Jarque-Berra test. The critical values were obtained from Hamilton (1994). The null hypothesis of no heteroskedasticity was tested using the White test. To compensate for the reported heteroskedasticity, the OLS standard errors have been adjusted using the White covariance adjustment according to the following equation:

$$\sum{}_w = \frac{T}{T-k}(X'X)^{-1}\left(\sum_{t=1}^{T} u_t^2 x_t x_t{}'\right)(X'X)^{-1}$$

Data Appendix: Table A1 - Data series details

Table A1 summarises the series used in this analysis. All data are at a quarterly frequency. The sample period was 1989Q1 to 2001Q2.

Series code	Series name	Details
p^m_t	Import prices	Log(Import price index for total merchandise imports)$_t$ + $tariff_t$ + x. See section 3 for explanation and derivation of x
	$tariff_t$	(Customs duty)/(Value of total merchandise imports)
p^c_t	Consumer prices	Log(CPI ex interest and GST)
GAP_t	Output gap	Output gap taken from Reserve Bank of New Zealand August 2001 projection
ulc_t	Unit labour costs	The official unit labour cost series for New Zealand (Labour Cost Index) only begins in 1992Q4. Therefore, a proxy series for the annual growth rate of unit labour costs back to 1989Q1 was constructed by subtracting annual productivity growth from annual nominal wage growth. The estimated annual growth rate series for unit labour costs was then combined with the five most recent observations (2000Q2 to 2001Q2) for the official Labour Cost Index to produce a levels series of Unit Labour Costs (ULC). Then $ulc_t = Log(ULC_t)$
	$ulc_t - ulc_{t-4}$	$(Log(Wage_t) - Log(Wage_{t-4}))$ $- (Log(Productivity_t) - Log(Productivity_{t-4}))$
	$Wage$	Seasonally adjusted average hourly earnings ordinary time private sector – Quarterly Employment Survey
	$Productivity$	(GDP/hours worked) run through a Hodrick-Prescott filter with $\lambda = 1,600$
	GDP	Seasonally adjusted real GDP – total production GDP from National Accounts
	$Hours\ worked$	Seasonally adjusted actual weekly hours worked – Household Labour Force Survey
	$Labour\ Cost\ Index$	Seasonally adjusted total salary and wage rates private sector – Labour Cost Index

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