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Product differentiation and cost pass-through: industry-wide versus firm-specific cost shocks*

Thomas Bittmann, Jens-Peter Loy and Sven Anders†

This paper investigates the impact of product differentiation on firm-specific and industry-wide cost pass-through in grocery retailing. We use attribute distance measures to model product differentiation based on a unique set of retail scanner data for ready-to-eat soup products in the Canadian market. Results from a panel error correction model suggest that product differentiation explains a significant share of the variation in the rate of cost pass-through across products. More differentiated products are associated with lower rates of cost pass-through of industry-wide and higher pass-through of firm-specific costs shocks. The findings validate an oligopolistic model of product differentiation, where firms use differentiation as a non-price competitive factor in strategic pricing decisions.

Key words: cost pass-through, non-price competition, product differentiation, industry-wide costs, firm-specific costs, food retail markets, ready-to-eat soups, Canada.

1. Introduction

The impact of competition on firms' pass-through of cost shocks is well documented in the literature. Product differentiation is an instrument for targeting consumer segments and generating price premiums. Given the high degree of price dispersion for fast-moving consumer goods, the role of product differentiation in explaining retail price dynamics has yet to be explicitly analysed in more detail.

Cost pass-through describes the effect of a change in input costs (prices) on firms' output prices (Walters *et al.* 2014).[1] 'Understanding the (cost) pass-through …. is of paramount importance not only for …. policy decision-making but also for portfolio risk management and optimal hedging issues' (Atil *et al.* 2014, p. 567). The extent of cost pass-through provides insights into the division between producer and consumer surplus and can provide vital information on the level of competition in markets (Walters *et al.* 2014).

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[1] Vertical price transmission is an alternative term which more specifically refers to the adjustment of output prices to changes in input prices by single firms or aggregates over firms. Bonnet and Villas-Boas 2016 use the two terms interchangeably. Following Loy and Weiss (2019), we exclusively use the term cost pass-through.

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In this context, Lloyd (2017, p. 4) refers to Hayek's observation 'that the adjustment of prices is the principal mechanism by which information about changes in demand and costs is communicated and argues that (vertical) price transmission can be used as a yardstick to measure the functioning of markets'.

More recently, economists have started to consider additional variables to improve the explanatory power of empirical models of cost pass-through, such as measures of market power, consumer search and menu costs. The majority of previous studies investigated the symmetry of short-run cost pass-through processes (Richards *et al.* 2014; Loy *et al.* 2015, 2016, Bonnet and Villas-Boas 2016). Durevall (2018) estimates long-run cost pass-through for the Swedish retail coffee market and finds significant differences related to market shares, retailer-owned brands and other product characteristics. Kim and Cotterill (2008, p. 46) use the term (product) differentiation in the context of estimating cost pass-through and state that 'there is a gap in the literature, as analysts have paid little attention to cost pass-through in differentiated product markets'. They estimate a structural model to derive product-specific cost pass-through elasticities and compare these with estimates of a reduced-form model. Kim and Cotterill (2008) find that the estimates of a reduced-form model for pooled data fall between the cost pass-through rates from a structural model under collusion or Nash–Bertrand. While Kim and Cotterill (2008) see this as a potential bias of reduced-form estimates, their results could also stem from ignoring the heterogeneity across panel members in such models.

While the theoretical foundation of structural models is appealing, their underlying assumptions have implications for the interpretation of empirical results. Empirical observations dominate the outcome of reduced-form models and present a better picture of actual relationships. However, endogeneity remains an issue of both approaches. The richness of the data available in this paper favours a reduced-form modelling approach also recommended by Leamer (1983) and Angrist and Pischke (2010).

Most studies that employ reduced-form models prioritise estimating short-run relationships between input costs and output prices. From a theoretical perspective, however, long-run relationships are more relevant. Hypotheses about long-run cost pass-through are well grounded in economic theory; the theoretical foundation regarding the dynamics of short-run pass-through rates, however, is still missing. In this paper, we estimate both long-run and short-run cost pass-through processes. We then compare our results to the literature and test relevant theoretical hypotheses.

Although there is a rich literature on differentiated product markets, studies that use explicit measures of product differentiation to explore their effects on cost pass-through are lacking (Loy and Weiss 2019). A recent review of existing price transmission theories and empirical findings by Bakucs *et al.* (2014) does not mention product differentiation at all. This is surprising, as the linkage between product differentiation and market power

is well established. For instance, Baker and Bresnahan (2008, p. 18) argue that 'the source of single-firm market power is differentiation'. Forty years ago, Shubik and Levitan (1980) developed an oligopolistic model based on Cournot's and Bertrand's assumptions to analyse the effects of product differentiation with a focus on the price–cost equilibrium when different numbers of firms are competing. Our models build on Shubik and Levitan (1980) to identify and estimate the direction of cost pass-through for industry-wide and firm-specific cost shocks.

Loy and Weiss (2019) were the first to empirically estimate the impact of a product differentiation measure in a reduced-form cost pass-through model for a European retail market. They show that German yoghurt prices depend on the level of differentiation in product characteristics. In particular, they find that industry-wide cost pass-through changes for more differentiated products. We follow the study by Loy and Weiss (2019) in estimating different measures of product differentiation and their impact on industry-wide cost pass-through rates.

We contribute to the literature by studying the effects of product differentiation on industry-wide and firm-specific cost pass-through.[2] Industry-wide costs apply to every producer; firm-specific costs apply only to one firm or one product. We use the oligopolistic model by Shubik and Levitan (1980) and Wang and Zhao (2007) to derive hypotheses on the cost pass-through. Data on firm-specific costs are hard to come by. We use a comprehensive retail scanner data set which combines retail prices at the UPC (Universal Product Code) level with the respective purchase prices (retail costs) of the same item to test the theory.

Our results support the findings by Loy and Weiss (2019). The elasticity of industry-wide cost pass-through is lower for more differentiated products, implying that firms differentiate products to generate additional market power. More differentiated products have higher prices compared to more similar products. Moreover, firm-specific cost shocks pass on as well; however, more differentiated products indicate higher firm-specific cost pass-through elasticities. For the most part, the theoretical model predicts this well.

The paper outline is as follows. In section 2, we explain the theoretical framework to derive the effects of product differentiation on cost pass-through. We develop the concept of measuring product differentiation in section 3. Section 4 describes the data and the estimation of industry-wide and firm-specific cost shocks. Following in section 5, we develop the model specifications and show the main testable hypotheses. In section 6, we present and discuss the results of the model estimations. In the end, we summarise our findings and point out some areas for future research.

[2] Some authors use the term common for industry-wide (costs) and the term idiosyncratic for firm-specific (costs). As we find industry-wide and firm-specific to be more intuitive, we use these terms throughout the paper.

2. A model of product differentiation and cost pass-through

We follow Shubik and Levitan (1980), Zimmermann and Carlson (2010) and Wang and Zhao (2007), starting with a market of i symmetric non-cooperative competitors offering n differentiated products, where $\phi \in [-1,0]$ symbolises the degree of product differentiation. Zero indicates the maximum level of product differentiation at which products are completely independent. When all products are perfect substitutes, $\phi = -1$.[3] Hence, an increase in ϕ represents an increase in average product differentiation or a corresponding reduction in substitutability. q_i is firm i's output. Equation (1) shows an inverse firm-specific linear demand function (Zimmerman and Carlson 2010).[4]

$$p_i = \alpha - q_i + \phi \sum_{j \neq i} q_j \tag{1}$$

We derive the profit-maximising quantity of firm i under a symmetric Cournot–Nash equilibrium (sCN), where q_j is a vector of outputs excluding firm i, by solving the reaction function for marginal costs c assuming symmetric firms. Equations (2) and (3) show the optimal output and the equilibrium price–cost relationship.

$$q_i^{sCN} = \frac{\alpha - c}{2 - \phi(n-1)} \tag{2}$$

$$p_i^{sCN} = \frac{\alpha + (1 - \phi(n-1))c}{2 - \phi(n-1)} \tag{3}$$

The resulting cost pass-through under Cournot–Nash price–cost equilibrium is positive yet smaller than one (equation (4)).

$$\frac{\partial p_i^{sCN}}{\partial c} = \frac{1 - \phi(n-1)}{2 - \phi(n-1)} > 0 \tag{4}$$

For a given number of firms, perfectly differentiated products ($\phi = 0$) produce a pass-through rate of 50 per cent and firms engage in monopolistic pricing. This result is similar to the monopoly case in Bulow and Pfleiderer (1983).

When all competitors offer homogeneous products ($\phi = -1$), cost pass-through depends on the number of firms ($n/\, n + 1$). This matches the pass-

[3] For ease of comparison and interpretation of the results of the theoretical and the empirical model, we use identical definitions of product differentiation. The standard definition of product differentiation in Zimmerman and Carlson's (2010) theoretical model is reversed ($-\phi \in [0,1]$), and therefore, 1 indicates perfect substitutes.

[4] The assumption of linear demand is important in this setting. As Weyl and Fabinger (2013) show, for non-linear demand functions and market power, cost pass-through rates may be higher compared with perfect competition.

through of an industry-wide cost shock in a Cournot oligopoly with homogeneous products (Ten Kate and Niels 2005).

To investigate the impact of product differentiation on firms' equilibrium prices, we differentiate equation (3) with respect to the product differentiation parameter ϕ, yielding.

$$\frac{\partial p_i^{sCN}}{\partial \phi} = \frac{(n-1)(\alpha - c)}{(2-\phi(n-1))^2} > 0 \tag{5}$$

In equation (5), prices increase with product differentiation. To determine the effect of product differentiation on the degree of cost pass-through, we partially differentiate equation (4) with respect to ϕ, resulting in equation (6).

$$\frac{\partial^2 p_i^{sCN}}{\partial c \partial \phi} = -\frac{(n-1)}{(2-\phi(n-1))^2} < 0 \tag{6}$$

Product differentiation reduces the cost pass-through, or in other words, closer substitutive relationships between products suggest a higher cost pass-through. Zimmermann and Carlson's (2010) analysis focuses on the role of competition, namely the number of firms in the market. Cost pass-through increases with the number of firms and converges to complete (one to one) cost pass-through under perfect competition ($n \to \infty$). Therefore, product differentiation and a declining number of firms have a similar impact on cost pass-through, making the Zimmermann and Carlson (2010) model of product differentiation a special case of Dixit's (1979) model.[5] While Shubik and Levitan (1980, p. 69) used a slightly different utility function, the results above remain unchanged and hold for both Cournot and Bertrand competition with respect to the signs of derivatives. The results are also robust for alternative specifications of the Dixit (1979) model (Loy and Weiss 2019). We present the results for Shubik's and Levitan's (1980) model in the Appendix.

Firms not only differ in their product portfolios but also in their levels of marginal costs, namely firm-specific costs. Retail purchase prices may therefore differ not only due to product differentiation but also due to firm's cost structures and the volumes produced by food processors. In a differentiated product market, the assumption of symmetric firms appears to be rather restrictive. For the data under study, we show later that retail purchase prices vary significantly across products (see Table 3).

We introduce firm asymmetry into the model following Wang and Zhao's (2007, p. 174) extensions of Shubik's and Levitan's (1980) model.

[5] Dixit (1979) uses the following utility function $U(\mathbf{q}, q_0) = \sum_{i=1}^{n} \alpha q_i - \frac{1}{2}\left(\sum_{i=1}^{n} \beta q_i^2 - 2\phi \sum_{j=1}^{n} q_i q_j + q_0\right)$ *for* $i \neq j$. In the Appendix, we document the utility function used by Shubik (1980).

Equation (7) shows the non-symmetric Cournot–Nash (*ncCN*) price–cost equilibrium relationship (Wang and Zhao 2007, p. 178). Again, for the ease of interpretation we transform the product differentiation measure, which ranges from minus infinity to zero ($\phi \in [-\infty, 0]$) (Zimmerman and Carlson 2010, p. 8). However, the interpretation of ϕ remains unchanged as higher parameter values indicate more differentiated products.

$$p_i^{nsCN} = \frac{(n-\phi)\alpha}{2n-(n+1)\phi} - \frac{n\phi\bar{c}(n-\phi)}{(2n-\phi)(2n-(n+1)\phi)} + \frac{nc_i}{2n-\phi} \tag{7}$$

$c_i, \bar{c} \equiv \sum c_i/n$ are firm-specific and average marginal costs, respectively. To determine the impact of changes in firm-specific marginal costs on cost pass-through, we differentiate equation (7) with respect to c_i.

$$\frac{\partial p_i^{nsCN}}{\partial c_i} = \frac{2n^2 - n(n+2)\phi + \phi^2}{(2n-\phi)(2n-(n+1)\phi)} > 0 \tag{8}$$

As for the case of symmetric firms, when $n = 1$ (monopoly) and products are independent ($\phi = 0$), the cost pass-through rate is 0.5. When the number of firms increases to infinity ($n \to \infty$), the cost pass-through remains at 0.5, regardless of the level of product differentiation. Zimmermann and Carlson (2010, p. 11) simulate the cost pass-through for different numbers of asymmetric firms. For a single firm (monopoly), the cost pass-through rate is 0.5 and declines with an increasing number of firms. After reaching a minimum, the rate increases and converges back to the monopoly cost pass-through rate of 0.5. Firms have low incentives to adjust prices to changes in firm-specific cost as consumers quickly switch to other firms, which face no cost changes. Firm-specific costs are passed on at lower rates than for the monopoly case. This effect becomes stronger with an increasing number of firms and with a lower degree of product differentiation. The number of firms, however, also flattens the reaction function in the Cournot model. This leads to a more uniform firm behaviour and to an increasing cost pass-through rate as the number of firms increases until the monopolistic cost pass-through rates are reached (Zimmerman and Carlson 2010).

To derive the impact of product differentiation on firms' equilibrium prices, we differentiate equation (7) with respect to product differentiation (ϕ) (equation (9)). Equation (9) shows that firms increase prices of more differentiated products.

$$\frac{\partial p_i^{nsCN}}{\partial \phi} = \frac{(n-1)n\left(4n^2\alpha + n(c(\phi-4))\phi + (2c+\alpha)\phi^2 - 4\alpha\right)}{(\phi-2n)^2(n(\phi-2)+\phi)^2} > 0, \text{ for } n > 1 \tag{9}$$

In equation (10), we derive the cost pass-through (equation (8)) with respect to the measure of product differentiation (ϕ). Contrary to the finding for the

symmetric case, firms offering more differentiated products increase their cost pass-through. The upper limit for the firm-specific cost pass-through is the monopoly rate of 0.5; the lower limit is zero.

$$\frac{\partial^2 p_i^{nsCN}}{\partial c_i \partial \phi} = \frac{(n-1)n\phi(n(\phi-4)+2\phi)}{(\phi-2n)^2(n(\phi-2)+\phi)^2} > 0, \text{ for } n>1 \tag{10}$$

To assess the robustness of the Cournot–Nash model under asymmetric cost structures, Zimmermann and Carlson (2010) also derive the Bertrand price–cost equilibrium relationship. They confirm the Cournot–Nash results. The individual cost pass-through rates have the same upper limit. For competition, cost pass-through rates are always higher, and the rates are monotonically decreasing with the number of firms. The result with respect to product differentiation is the same; cost pass-through of firm-specific costs increases for more differentiated products. Contrary to the Cournot–Nash model outcome, the presence of product differentiation reinforces the effect of intensifying competition; an increase in ϕ positively affects a firm's response to rivals' output decisions. Consequently, cost pass-through rates strictly decline in the number of competing firms (Zimmerman and Carlson 2010).

In conclusion, we expect prices of more differentiated products to be *ceteris paribus* higher and cost pass-through of industry-wide cost shocks to be lower. Cost pass-through is decreasing with product differentiation. For homogenous products and a large (infinite) number of firms, the cost pass-through rate is one. The minimum cost pass-through rate of industry-wide cost shocks is 0.5, as it is in the monopolistic case, implying that 50 per cent of a cost shock is passed on to the retail price. For individual cost shocks, pass-through rates increase with product differentiation, starting at a cost pass-through rate of zero for homogenous products. Independent (perfectly differentiated) products show a pass-through rate of 0.5 for firm-specific cost shocks. For homogenous products, individual costs cannot be passed on to consumers as demand reacts perfectly elastic. For differentiated products, firms can pass on costs. For independent products, pass-through rates converge towards the monopolistic rate of 0.5.

The theoretical models discussed so far do not present any information on the dynamics of the cost pass-through process. Borenstein and Shepard (2002) argue that market power slows down the cost pass-through process. Baker and Bresnahan (2008) point out that product differentiation is a means to generate market power. Thus, we assume that more differentiated products indicate a more sluggish cost pass-through process. Many empirical studies test for asymmetric dynamic responses, namely the 'rocket and feathers' phenomenon (Tappata 2009). Retailers may exert market power by differentiating their product portfolio to raise prices more rapidly in response to a cost increase than to cut prices in response to a cost decrease. Borenstein *et al.* (1997, p. 324) argue that 'prices are sticky downward because when input

prices fall the old output price offers a natural focal point for oligopolistic sellers'. Thus, for more differentiated products, we expect retailers to adjust prices more asymmetrically.

3. Measures of product differentiation

We use UPC-specific product attribute information that covers the ingredients of (canned) soup products, brand names, flavours and package sizes. The data are obtained from Mintel's Global New Products Database (Mintel 2015). Product ingredient data include calories, fat, cholesterol, sodium, carbohydrates, fibres, sugar and protein content, of which two characteristics, brand names and flavours, are discrete variables. In total, the measurement of product differentiation is based on eleven product attributes. The sample includes four brands and seven flavours, for which we create dummy variables. Table 1 provides descriptive statistics of product attributes.

We assume that each soup product is uniquely placed within the multidimensional attribute space represented by a vector z of coordinates of the k attributes $(z = z(z_1, z_2 \ldots\ldots\ldots, z_k))$. This modelling approach allows us to conceptualise product differentiation in terms of the underlying attribute data. The representation of the price of a product as the sum of the values of its attributes is at the centre of hedonic pricing analysis (Lancaster 1966; Rosen 1974). Structural models use the assumption that consumers optimise utility based upon product attributes (Kim and Cotterill 2008). Discrete choice settings employ an attribute space model of product differentiation (e.g. McFadden 1974). Consumers are assumed to choose the product that maximises utility, which again is a function of product attributes. Pinske *et al.* (2002) and Pinske and Slade (2004) incorporate a multidimensional measure of product differentiation in their analysis to reduce the number of model variables. Similarly, Rojas and Peterson (2008) and Bonanno (2012) make use of a distance metric to introduce product differentiation to an Almost Ideal Demand System, and Rojas (2008) analyses advertising patterns.

The estimation of multidimensional differences is similar in nature to the analysis of multidimensional similarities, such as in cluster analysis. This allows to use cluster analysis methods to measure the degree of product differentiation (Kaufman and Rousseeuw 2009, Chapter 2). A common practice is to standardise all variables that build the product attribute space. However, aggregation can be difficult due to different attribute scales, for example continuous and discrete. Further, a variety of distance measures is available, such as Euclidean, Manhattan City Block or maximum distances (Cha 2007). In this paper, we follow the measurement of product differentiation in Loy and Weiss (2019) and develop an alternative measure based on recommendations by Gower (1971), and Gower and Warrens (2017), which considers aggregation over mixed variables. Krzanowski (1993) recommends Gower's coefficient to cope with mixed variables.

Table 1 Summary statistics of product attributes

Variables	Mean	Standard Deviation	Minimum	Maximum
Brand name (1 to 4)	1.74	1.01	1	4
Flavour (1 to 7)	4.01	2.17	1	7
Calories (per cup)	158.10	48.40	30	250
Fat (g/cup)	4.76	4.75	0	15
Cholesterol (mg/cup)	8.33	7.80	0	20
Sodium (mg/cup)	956.90	426.70	360	1700
Carbohydrate (g/cup)	22.52	6.90	4	34
Fibre (g/cup)	2.74	2.24	0	10
Sugar (g/cup)	6.97	7.02	0	20
Protein (g/cup)	5.71	2.03	2	9
Package size (ml)	422.4	181.4	284	796

Source: Own calculation based on Mintel (2015).

Loy and Weiss (2019) first create separate distance matrices for each product attribute. They measure the similarity between products by employing the inverse Euclidean distance as in Rojas and Peterson (2008) and then reverse it to result in a measure of product differentiation. We deviate here by directly measuring distances. For discrete variables, we create matrices of zeros and ones, where one indicates different products and zero indicate similarity in a product attribute. We then calculate row averages of the matrices. The entries of these vectors for discrete attributes represent shares of different products, for example a share of 0.3 means that a specific UPC-coded product is of the same brand in 70 per cent of all products or differs from 30 per cent of all products. For continuous attribute variables, we calculate the squared distances between all product pairs and fill the distance matrices accordingly. We again take row averages and repeat this procedure for all attributes. We standardise all vectors to result zero means and standard deviations of one. Then, we sum over all standardised vectors. Finally, we shift the variable to a minimum of zero and divide it by the standard deviation (Table 2).

Loy and Weiss (2019) use a second procedure to aggregate mixed variables by transforming all continuous attributes into binary variables. As this procedure would eliminate information in our data and may generate some arbitrary differentiations (Krzanowski 1993), we choose an alternative second measure following Milligan and Cooper (1988). We standardise all continuous characteristics by using the range to balance the effect of the different scales. The measure is capable of aggregating over discrete and continuous data (Gower 1971). The standard element of Gower's dissimilarity coefficient is the average of all attributes' differences. For discrete variables, the distance is either 0 if the products are of the same brand or flavour or 1 if not. For continuous variables, we calculate absolute differences in attribute values between products using a Manhattan or city block distance normalised by the range of the attribute across all products. The range standardises and renders the metric scale-invariant; thereby, entries of the distance matrix range from zero to one. Gower's dissimilarity coefficient is obtained as the average over

Table 2 Alternative measures of product differentiation and correlations

Variable	Distance metric	Mean	SD	Min	Max
ϕ^G	Gower	1.29	1	0	3.46
ϕ^{LW}	Loy and Weiss	1.30	1	0	3.85
ϕ^E	Euclidian	0.90	1	0	4.76
ϕ^{SD}	Squared Differ.	0.88	1	0	5.22
ϕ^{BD}	C. Block Distance	1.15	1	0	4.51

	ϕ^G	ϕ^{LW}	ϕ^E	ϕ^{SD}	ϕ^{BD}
ϕ^G	1				
ϕ^{LW}	0.96	1			
ϕ^E	0.87	0.84	1		
ϕ^{SD}	0.85	0.82	0.99	1	
ϕ^{BD}	0.95	0.92	0.98	0.97	1

Source: Own calculation based on Mintel (2015).

all continuous and discrete distance matrices. Finally, we take the row sum of the matrix. The resulting variable is divided by its standard deviation and shifted to a minimum at zero. Table 2 shows descriptive statistics for the two measures.

We calculate three additional measures of product differentiation. We use the same matrices for discrete variables as before. For continuous variables, we standardise all attribute vectors (zero mean and standard deviation of one), before we fill the distance matrices for each attribute and sum over all matrices; for the resulting matrix, we calculate row sums. The vector of row sums is also standardised to a minimum of zero and variance of 1. Rows 3 to 5 in Table 2 show the results for the additional measures.

We use Euclidian, squared and city block distances. Euclidean and squared distances make larger differences more significant; as a consequence, the weight of continuous variables increases over discrete variables, especially if the discrete distance matrices contain many zeros. Both measures deviate in their mean and maximum from the Gower coefficient. The city block distance is closer to the Gower coefficient in terms of its mean and maximum value. The Gower coefficient and the distance measure based on Loy and Weiss (2019) are almost identical. Correlations between the product differentiation measures are high, in particular the Gower and Loy and Weiss (2019) coefficients are very close.

4. Data and descriptive statistics

The analysis in this paper uses a set of weekly store-level retail scanner data with 178 (T) observations over time (2004 to 2007) for 568 cross-sectional units (N) (SIEPR-Giannini Data Center 2012).[6] The retail stores in the

[6] The data were obtained from SIEPR-Giannini Data Center via a research agreement between the retail chain and co-author Sven Anders at the University of Alberta. The data are not available due to third-party restrictions.

Table 3 Summary statistics of retail and purchase prices in CAD$ (2004-2007)

Variable	Mean	SD	Min	Max	Obs.
Retail prices	1.92	2.19	0.35	13.92	101104
Average retail prices#	1.92	0.13	1.59	2.30	178
Logarithms of retail prices (p_{it}^r)	0.38	0.63	−1.03	2.63	101104
Purchase prices	1.48	1.96	0.22	10.09	101104
Average purchase prices#	1.48	0.11	1.22	1.77	178
Logarithms of purchase prices (p_{it}^w)	0.02	0.75	−1.51	2.31	101104

Note:: p_{it}^r and p_{it}^w indicate logarithms of prices in the following. # cross-sectional averages.
Source: Own calculation based on SIEPR (2012).

sample belong to a major North American retail chain. The data include information on UPC-level retail prices and respective purchase prices revealing variable retail costs.

The canned soup category represents a perfect opportunity to investigate strategic product placement and its implications for cost pass-through. The soup market is a saturated consumer market with an oligopolistic market structure with significant non-price competition as indicated by the high degree of product differentiation in ingredients and nutritional factors. The leading brand has a market share of 70 per cent. The sample consists of 20 products sold in 70 stores. Table 3 presents descriptive statistics of retail and purchase prices per unit in CA$. The average unit size is about 420 ml, with products ranging from 284 to 796 ml. We use a logarithmic transformation of prices to account for the differences in unit sizes and for testing and model estimation purposes. Over the sample period, the average retail price of canned soup is 1.92 CA$ per unit, and the average purchase price is 1.48 CA$ per unit. The average markup is 44 cent or about 30 per cent.

The standard deviation or the range over retail and purchase prices indicates significant price dispersion with coefficients of variation above 100 per cent. Taking averages over the cross section largely diminishes this variation. The average retail prices show a coefficient of variation of about 7 per cent, which is the same as for the purchase price. Though smaller than the variation over the cross section, average prices indicate significant changes over time (see Figure 1).

Panels of price data for products within the same category often show a high level of cross-correlation. Applying Pesaran's (2004) test, we find evidence of cross-sectional dependence for retail and purchase prices, respectively. Purchase prices represent retailing costs, and cross-correlation implies that some movements of costs over time are similar to all products. We therefore separate purchase prices into industry-wide and firm-specific cost components. We employ two different approaches. First, Pesaran (2006) approximates unobserved factors in terms of cross-sectional averages to eliminate the differential effects of unobserved common factors. We follow this approach and use the average purchase prices over all products (firms) to

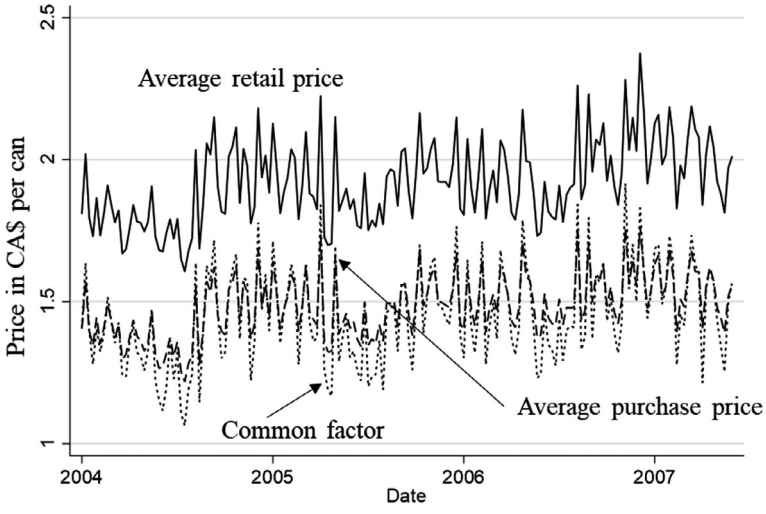

Figure 1 Average retail prices and measures of industry-wide costs. *Note*: Correlation between the first principal component and the average purchase price is 0.97. Weekly average retail and purchase prices from 2004 to 2007 in CA$. The first principal component is extracted from an approximate factor model on weekly wholesale prices. Source: Own calculations based on SIEPR (2012).

estimate an industry-wide cost component. The differences between the individual and the average purchase prices represent the firm-specific cost component.

For the second approach, we follow Bai and Ng (2002) who estimate the number of common factors by the method of asymptotic principal components. Determining the number of common factors is a model selection problem. In particular, for panel data with large N and T, the Bayesian information criterion (BIC) is valid and consistent in determining common factors (Bai and Ng 2002). BIC indicates that purchase prices share one common factor. Figure 1 illustrates the movement of the average retail prices and the two measures of industry-wide costs, the average purchase price and the estimated common factor. Cross-sectional averages of retail and purchase prices strongly co-move, indicating that industry-wide costs appear to be a significant driver of average retail prices.

The variation of the industry-wide cost component only makes up a small amount of the overall variation in purchase prices (see Table 3). Most of the variation comes from differences between average product purchase prices (cross-sectional variation). To estimate the variation in firm-specific costs, we calculate the average of the standard deviation of the individual series of firm-specific costs. The results show that firm-specific costs vary more than industry-wide costs over time. While the standard deviation of the industry-wide component is 0.11, the firm-specific component standard deviation is 0.66. The central limit theorem explains why variations of firm-specific costs exceed the variation of industry-wide costs.

Table 4 Phillips–Perron and KPSS tests for average logarithmic prices

Variable	Test statistic	Lags	Null hypothesis
Phillips–Perron Test			
Retail price $\left(p_{it}^{\overline{r}}\right)$	−2.05 [a)]	3	Non-stationary
Purchase price $\left(p_{it}^{\overline{w}}\right)$	−2.14 [a)]	6	Non-stationary
KPSS test			
Retail price $\left(p_{it}^{\overline{r}}\right)$	1.51*** [b)]	10	Stationary
Purchase price $\left(p_{it}^{\overline{w}}\right)$	1.31*** [b)]	9	Stationary

Note: a) Optimal lag length is chosen according to Ng and Perron (1995). Critical values 10%: −3.14; 5%: −3.44; 1%: −4.02. Automatic bandwidth selection is according to Newey and West (1994). b) Critical values: 10%: 0.35; 5%: 0.46; 1%: 0.74. ***, ** and * denote 1%, 5% and 10% level of significance. For the ADF test, optimal lags are chosen according to Ng and Perron (1995). In case of KPSS tests, automatic bandwidth selection is chosen according to Newey and West (1994).
Source: Own calculations based on SIEPR (2012).

Non-stationarity is a common feature of (retail) price data (Richards *et al.* 2014; Loy *et al.* 2015). For panel structures, specific (panel) unit root and co-integration tests have been developed. Many panels of price data such as the data under study show strong cross-correlations. First-generation panel unit root tests ignore this cross-sectional dependence (Breitung and Das 2008). Additionally, the likelihood of rejecting the null hypothesis increases with the number of panel members (Hanck 2008). Second-generation panel unit root tests consider cross-sectional dependence. However, Breitung and Das (2008) show that second-generation tests are invalid when common factors are non-stationary. Therefore, we subtract cross-sectional means from every time series before applying ADF and KPSS stationarity tests to each of the individual panel series and the cross-sectional means separately. The ADF test result rejects the null hypothesis of non-stationarity for individual panel series from retail (purchase) prices in 84 (64) per cent of all series at 5 per cent significance. The KPSS test rejects the null hypothesis of trend stationarity for the individual retail (purchase) price series at the 5 per cent significance level in 37 (65) per cent of all series. We follow the KPSS test results, which for the majority of firm-specific costs rejects stationarity. Table 4 shows the results of unit root tests for the cross-sectional averages of retail and purchase prices. The results indicate that both series are non-stationary. A panel with a factor structure is non-stationary if a common factor is non-stationary and/or the individual component is non-stationary (Breitung and Das 2008). Thus, we assume retail prices and purchase prices to be non-stationary. First differences of all individual and average price series are tested stationary. Thus, we assume all series to be integrated of order 1.

To check the causal structure of the model, we apply Granger causality tests to the individual time series of retail and purchase prices and to the average prices. After fitting a pooled reduced-form panel VAR on individual series, we use Wald tests with the null hypothesis that the coefficients on all the lags of the endogenous variable are jointly equal to zero (Love and

Table 5 Westerlund panel co-integration tests

H_o: No co-integration	Statistic	*P*-value
H_{a1} : Group mean statistics		
Group_τ	−5.93***	0.00
Group_α	−99.15***	0.00
H_{a2} : Panel statistics		
Panel_τ	−166.27***	0.00
Panel_α	−146.92***	0.00

Note:: Panel co-integration tests for logarithms of retail prices $\left(p_{it}^{r}\right)$, industry-wide $\left(p_{it}^{\overline{w}}\right)$ and firm-specific costs $\left(p_{it}^{w*}\right)$. H_0: no co-integration; bootstrapping critical values with 500 draws. Lag length selected between 1 and 7 according to AIC. Average selected lag length is between 2 and 3. ***, ** and * denote 1%, 5% and 10% level of significance.
Source: Own calculations based on SIEPR (2012).

Zicchino 2006). The results indicate at the 5 per cent level of significance that purchase prices Granger-cause retail prices and not vice versa.

Panel VARs of non-stationary processes with large numbers of cross-sectional units have a standard limiting distribution under general conditions (Holtz-Eakin *et al*. 1988). In contrast, Wald tests on Granger causality for non-stationary cross-sectional means may deliver spurious results, because the test statistic does not follow a chi-square distribution under the null (Dolado and Lütkepohl 1996). For Granger causality tests of cross-sectional means, we estimate bivariate VAR specifications with an additional surplus lag that is not used in the hypothesis test (Dolado and Lütkepohl 1996, Toda and Yamamoto 1995). The results show that average purchase prices cause average retail prices and not vice versa. This result holds at the 10 per cent significance level.[7]

Westerlund (2007) developed four panel co-integration tests for a null hypothesis of no panel co-integration against the alternative $(H_{\alpha 1})$ that at least one member of the panel is co-integrated. He shows that each statistic converges to a standard normal distribution and that the tests have sufficient power even in small samples. We use bootstrapped confidence intervals that are robust to cross-sectional dependence. Table 5 reports the test results.

All tests reject the null hypothesis of no panel co-integration at the 1 per cent level of significance. This provides solid evidence of existence of a long-run equilibrium relationship between retail and industry-wide (average purchase) and firm-specific costs.

In the next chapter, we develop empirical model specifications to estimate the long-run price–cost equilibrium and the dynamics of the cost pass-through process.

[7] Detailed results of Granger causality test can be obtained from the authors upon request.

5. Empirical model specifications

We estimate a reduced-form panel cost pass-through model for $i = 1, ..., N$ price series of differentiated products over $t = 1, ..., T$ time periods as in Richards *et al.* (2014) and Loy and Weiss (2019). To estimate the impact of product differentiation on retail prices, we introduce the measure of product differentiation as a right-hand-side variable. To estimate the effect on the long-run cost pass-through, we introduce a multiplicative term of product differentiation times the industry-wide cost component. To estimate the impact of firm-specific costs, we multiply the product differentiation measure with the firm-specific cost component. These three variables allow us to estimate the effects of product differentiation on retail prices and on the cost pass-through rates. Equation (11) shows the final panel model specification:[8]

$$p_{it}^{r} = \beta^{0} + \beta^{\phi}\phi_i + \beta^{\overline{w}} p_{it}^{\overline{w}} + \beta^{\overline{w}\phi} p_{it}^{\overline{w}}\phi_i + \beta^{w*} p_{it}^{w*} + \beta^{w*\phi} p_{it}^{w*}\phi_i + \nu_i + u_{it}, \tag{11}$$

where p_{it}^{r} denotes the logarithm of retail prices, $p_{it}^{\overline{w}}$ is the average of the logarithm of purchase prices (industry-wide costs), and p_{it}^{w*} stands for firm-specific costs calculated by the differences between the logarithm of purchase prices and the average of the logarithm of purchase prices. β^{0} is the deterministic intercept. The intercept v_{i} is random with zero mean, and u_{it} is the error term.

The model specification includes the measure of product differentiation (ϕ_i) and two multiplicative interaction terms $\left(p_{it}^{\overline{w}}\phi_i, p_{it}^{w*}\phi_i\right)$. These allow us to directly test for the impact of product differentiation on industry-wide and firm-specific cost pass-through. $\beta^{\overline{w}}$ and β^{w*} define the reference levels of cost pass-through for homogenous products $(\phi_i = 0)$. If $\beta^{\phi} > 0$, more differentiated products show higher retail prices. A negative value for $\beta^{\overline{w}\phi}$ implies that cost pass-through for industry-wide costs decreases with the level of product differentiation. A positive value for $\beta^{w*\phi}$ implies that cost pass-through for firm-specific costs increases with the level of product differentiation. If $\beta^{\phi} = \beta^{\overline{w}\phi} = \beta^{w*\phi} = 0$, product differentiation has no effect on the long-run price–cost equilibrium.

In a second stage, we investigate the short-term dynamics of the cost pass-through process by estimating a panel error correction model. To test whether the degree of product differentiation influences the speed of cost pass-through adjustment, we introduce a multiplicative interaction term between the error correction term and the product differentiation measure. The error correction term (ect_{it}) is estimated by the error term in the first stage (equation (11). Equation (12) shows the panel error correction specification.

[8] Note that identical products may be sold at multiple stores. For the ease of reading, we suppress store indices.

$$\Delta p^r_{it} = \alpha^0 + \delta ect_{it-1} + \delta^{\phi} ect_{it-1}\phi_i + \sum \alpha^{wj} \Delta p^w_{it-j} + \sum \alpha^{rj} \Delta p^r_{it-j} + \gamma_i + \varepsilon_{it} \quad (12)$$

δ indicates the speed of adjustment back to the price–cost equilibrium relationship. We expect δ to be negative and larger than minus 1. The closer δ is to minus 1, the faster is the adjustment. If $\delta^{\phi} \neq 0$, the speed of adjustment changes with the level of product differentiation. A positive coefficient δ^{ϕ} indicates that retail prices of more differentiated products adjust more slowly to cost shocks, similar to the effect of reduced competition (Borenstein and Shepard 2002, Loy and Weiss 2019).

We also test for short-run asymmetry in the speed of adjustment of the dynamic process by the following asymmetric panel error correction specification (equation (13)).

$$\begin{aligned}\Delta p^r_{it} = \; & \alpha^0 + \delta^{+} ect^{+}_{it-1} + \delta^{-} ect^{-}_{it-1} + \delta^{+\phi} ect^{+}_{it-1}\phi_i + \delta^{-\phi} ect^{-}_{it-1}\phi_i + \sum \alpha^{wj} \Delta p^w_{it-j} \\ & + \sum \alpha^{rj} \Delta p^r_{it-j} + \gamma_i + \varepsilon_{it}\end{aligned} \quad (13)$$

Here, the error correction term (ect_{it}) is split into positive and negative observations $\left(ect_{it} = ect^{+}_{it} + ect^{-}_{it}\right)$. The 'rocket and feather' phenomenon implies that negative deviations are reduced faster than positive ones (e.g. Tappata 2009). If product differentiation enforces this behaviour, positive deviation will be reduced even more slowly and/or negative ones will be reduced even faster $\left(\delta^{+\phi} > 0 \text{ or } \delta^{-\phi} < 0\right)$.

In accordance with the theory in section 2, we estimate the model specification in equations (11), (12) and (13) to test the following set of hypotheses:

1: $\beta^{\phi} > 0$, more differentiated products show higher retail prices (equation (11)).

1: $\beta^{\overline{w}\phi} < 0$, more differentiated products have a lower cost pass-through elasticity of industry-wide cost shocks (equation (11)).

1: $\beta^{w*\phi} > 0$, more differentiated products have a higher cost pass-through elasticity of product-specific cost shocks (equation (11)).

1: $\delta^{\phi} > 0$, more differentiated products return more slowly to the long-run price–cost equilibrium (equation (12)).

1: $\delta^{+\phi}>0$ and/or $\delta^{-\phi}<0$, product differentiation reinforces the asymmetric cost pass-through (equation (13)).

6. Estimation results

We estimate the long-run cost pass-through model (equation 11) using both random-effects and fixed-effects estimators. The random-effects model allows us to estimate the effects of time-invariant variables (product differentiation). The fixed-effects model is estimated to check the robustness. The Hausman test favours the fixed-effects model; however, the differences between the model estimates are rather small (Hausman 1978). For the random-effects model, we use two measures of product differentiation, which we discussed above, the Gower coefficient and the measure used in Loy and Weiss (2019). Thereby, we test the robustness of the product differentiation measure.

To capture the short-run dynamics of the cost pass-through process, we estimate the second-stage panel error correction model using the regression residuals from the first-stage random-effects model. In line with previous retail cost pass-through studies, we assume purchase prices to be exogenous (e.g. Richards *et al.* 2014; Loy *et al.* 2015, 2016; Loy and Weiss 2019). The Granger causality tests reported above support this assumption. Table 6 shows the results of the first-stage regression.

The estimation results are very similar for all three models. In the following, we mainly discuss the favoured model (1), the random-effects model using the Gower coefficient.

The results of model (1) provide evidence that product differentiation does exert a significant influence on cost pass-through. The null hypothesis of $\beta^{\phi}=\beta^{\overline{w}\phi}=\beta^{w*\phi}=0$ is tested with a chi-square test and is rejected at one per cent.[9] A high R^2 of 0.85 indicates an overall goodness of model fit. All coefficients show the expected signs and support the theory outlined in section 2. Higher retail prices are charged for more differentiated products $\left(\text{H1}:\beta^{\phi}>0\right)$. The most differentiated soup product $(\max(\phi)=3.46)$ is about 21 per cent more expensive than the least differentiated product $(\min(\phi)=0)$. A unit increase in product differentiation leads to an expected increase in the retail price by 6 per cent.

Product differentiation leads to a reduced pass-through of industry-wide cost shocks $\left(\text{H3}:\beta^{\overline{w}\phi}<0\right)$. The retail price of the most substitutable product increases on average by 0.95 per cent in response to a one per cent change of industry-wide costs. For the most differentiated product, the pass-through of industry-wide cost falls considerably to 71 per cent (0.71 = 0.95–0.07 * 3.46). The above results are in line with Loy and Weiss (2019) who investigated the effect of product differentiation on industry-wide cost pass-through the

[9] Table A1 in the appendix shows the estimation results based on the alternative distance measures (Table 2). Overall, the results are robust to alternative specifications of the product differentiation measure.

Table 6 Fixed- and random-effects panel model of cost pass-through (equation (11))

Dependent variable: retail prices	Model 1 Random effects Gower	Model 2 Random effects Loy & Weiss	Model 3 Fixed effect Gower
Independent variables, Coeff.	$\phi = \phi^{G}$	$\phi = \phi^{LW}$	$\phi = \phi^{G}$
Product differentiation (PD),β^{ϕ}	0.06*** (0.01)	0.04*** (0.01)	-
Industry-wide costs,$\beta^{\overline{w}}$	0.95*** (0.05)	0.97*** (0.05)	0.93*** (0.05)
Industry-wide c. * PD,$\beta^{\overline{w}\phi}$	−0.07*** (0.03)	−0.09*** (0.03)	−0.06** (0.03)
Firm-specific costs,β^{w*}	0.89*** (0.02)	0.84*** (0.02)	0.82*** (0.02)
Firm-specific c. * PD,$\beta^{w*\phi}$	0.05** (0.02)	0.09*** (0.02)	0.13*** (0.02)
Constant,β^{0}	0.28*** (0.01)	0.31*** (0.01)	0.37*** (0.00)
R^2, within	0.84	0.84	0.84
R^2, between	0.91	0.90	0.89
R^2, overall	0.85	0.84	0.83
Observations	101,104	101,104	101,104

Note:: Robust standard errors are reported in parentheses. ***, ** and * denote 1%, 5% and 10% level of significance.
Source: Own calculation based on SIEPR (2012) and Mintel (2015).

German retail yoghurt market. Their estimates of cost pass-through rates for the most differentiated and least differentiated products range from 36 to 86 per cent.

For firm-specific cost shocks, we find a positive effect of product differentiation on retail prices $\left(\mathrm{H3}: \beta^{w*\phi} > 0\right)$. Accordingly, 89 per cent of a firm-specific cost shock passes on to the retail price of the least differentiated soup product. Cost pass-through of individual cost shocks is smaller than 100 per cent, thus incomplete. In contrast, the most differentiated product features a pass-through of firm-specific costs of 106 per cent ($1.06 = 0.89 + 0.05 * 3.46$).

The estimated coefficients of the cost pass-through of industry-wide costs match the results of Loy and Weiss (2019), the outcome of Shubik's and Levitan's (1980) oligopolistic model, and those of Wang and Zhao (2007). Although the impact of product differentiation on firm-specific costs confirms the theoretical model, the empirical estimates of cost pass-through rates of nearly 100 per cent are incompatible with the theory. This, however, may be due to the theoretical model's strong assumptions regarding the underlying structure of the market as well as the functional form of demand. As Weyl and Fabinger (2013) show, under non-linear demand market power can result in cost pass-through rates well above 100 per cent.

Further, after separating firm-specific from industry-wide costs into independent cost components, firm-specific costs and particularly those for the same firm may still be correlated. Thereby, firm-specific costs may capture industry-wide costs to some degree. We find correlations between individual firm-specific costs of 0.2 to 0.5. Thus, a fraction of firm-specific costs also applies to other products or firms. Therefore, the above estimates may indicate some bias towards industry-wide cost pass-through.

Table 7 Symmetric and asymmetric random-effects panel error correction model

Symmetric model (equation (12))		
Dependent variable: First differences of retail prices	Model 4 Gower	Model 5 Loy & Weiss
Independent variables, coefficient	$\phi = \phi^G$	$\phi = \phi^{LW}$
Error correction term (ect),δ	−0.67*** (0.03)	−0.63*** (0.03)
ect * product differentiation,δ^{ϕ}	0.04*** (0.01)	0.02*** (0.01)
R^2	0.41	0.41
Observations	98,264	98,264
Asymmetric model (equation (13))		
Dependent variable: First differences of retail prices	Model 6 Gower	Model 7 Loy & Weiss
Independent variables, coefficient	$\phi = \phi^G$	$\phi = \phi^{LW}$
ect^+,δ^+	−0.67*** (0.05)	−0.60*** (0.05)
ect^-,δ^-	−0.88*** (0.05)	−0.88*** (0.05)
ect^+ * product differentiation,$\delta^{+\phi}$	0.29*** (0.03)	0.27*** (0.03)
ect^- * product differentiation,$\delta^{-\phi}$	0.03*** (0.01)	0.02*** (0.01)
R^2	0.41	0.41
Observations	98,264	98,264

Note:: Robust standard errors in parentheses. ***, ** and * denote 1%, 5% and 10% significance. The symmetric lag length is determined by AIC. The results of first difference lags of retail and purchase prices are omitted.
Source: Own calculation based on SIEPR (2012) and Mintel (2015).

Another reason for why the theory fails to accurately predict higher pass-through rates of firm-specific costs may lie in the specific structure of the retail market under study. The theoretical model assumes symmetric firms producing one product under Cournot or Bertrand competition. The Canadian retail soup market is dominated by one firm (manufacturer) producing multiple products with a combined market share of 70 per cent. A Stackelberg leader–follower model with multiple products may thus better reflect market structure and outcomes. Further, both prices and quantities can be strategic variables, which are negotiated simultaneously. Haucap *et al.* (2015) argue that traditional models of competition cannot explain the outcomes of complex negotiations between food processors and retailers. Gaudin (2016, 2018) presents based on a similar argument results on the effects of the type of the vertical agreement and the relative bargaining power on cost pass-through. Future research needs to explore more of these and other alternative frameworks.

To analyse the short-run dynamics of the cost pass-through processes, we estimate the two model specifications shown in equations (12) and (13). Equation (12) replicates the model in Loy and Weiss (2019) and estimates a symmetric process. As many studies have investigated the 'rocket and feathers' phenomenon, we also estimate the standard asymmetric model shown in equation (13). Table 7 presents the estimates of the symmetric and the asymmetric retail cost pass-through model.

For the symmetric model, we find that relative deviations from the long-run price–cost equilibrium are adjusted back to the equilibrium at a rate of 63 to 67 per cent per period. Thus, after 4 periods (weeks) 99 per cent of the deviation from equilibrium is reversed. Loy and Weiss (2019) find a much slower adjustment process with δ equal to −0.08 to −0.19, an adjustment rate of 8 to 19 per cent. A reason behind the difference in the speed of adjustment could be the cost measure. Loy and Weiss (2019) use the price of raw milk as the retail cost measure for yoghurt. We use product-specific purchase prices, which are much closer to actual retail costs and therefore produce a better fit and a faster adjustment. Our results are closer to the findings by Richards *et al.* (2014) who report a δ of 57 per cent for ready-to-eat cereal products in the Los Angeles retail market. Bittmann and Anders (2016) find cost pass-through rates between 10 and 49 per cent for fresh apples using scanner data. Finally, Loy *et al.* (2015) find regime-dependent pass-through rates of between 15 and 55 per cent in the German dairy retail market.

δ^{ϕ} measures the effect of product differentiation on the speed of adjustment back to the price–cost equilibrium. If $\delta^{\phi}>0$, product differentiation slows down the speed of adjustment. With significant estimates of 0.02 to 0.04, we confirm the hypothesis H4 for both measures of product differentiation. Retail prices of the most differentiated products move back towards equilibrium at a rate of 53 per cent instead of 67 per cent for their less differentiated counterparts ($-0.53 = -0.67 + 0.04 * 3.46$), a 14 percentage point difference. Instead of 4 weeks, it takes 6 weeks to reduce the deviation by 99 per cent. Loy and Weiss (2019) find a similar result for the minimum and maximum values of product differentiation with a difference of 12.6 percentage points.

The coefficient estimates of the asymmetric model (equation 13) provide evidence of asymmetry in the speed of adjustment back to the price–cost equilibrium. As expected, in the context of market power, the asymmetry reveals a 'rocket and feathers' pattern. Positive deviations from the price–cost equilibrium (*ceteris paribus*, retail prices and margins are above average) are reduced at a lower rate than negative deviations (*ceteris paribus*, retail prices and margins are below average). The results show that the difference in the speed of adjustment between positive (δ^{+}) and negative (δ^{-}) deviations is 11 and 28 percentage points, respectively. Thus, product differentiation enforces asymmetric behaviour; in particular, the reduction of positive deviations is slowing down significantly. For the mean value of the product differentiation measure ($\phi = 1.3$), the rate of adjustment for positive deviation is 29 per cent ($-0.29 = -0.67 + 1.3 * 0.29$). For negative deviations, the rate of adjustment is 84 per cent ($-0.84 = -0.88 + 1.3 * 0.03$). The asymmetry rises from 11 percentage points for the most substitutable product to 55 percentage points for the most differentiated product.

Richards *et al.* (2014) and Loy *et al.* (2016) argue that differences in menu costs and product variety provide a partial explanation of variations in cost pass-through rates within product category with comparable levels of product

differentiation. Loy *et al.* (2015, 2016) find manufacturer, retailer and brand characteristics to be important determinants of dynamic (asymmetric) cost pass-through rates for dairy products in the German retail market. The authors specifically point to the role product differentiation plays in shaping the variation of cost pass-through dynamics in their data.

A higher degree of differentiation implies fewer close neighbours in the product attribute space. Such products may exhibit a higher degree of price staggering as pass-through of product-specific cost components dominates the retailer's pricing decisions, while price-setting decisions for products of lower differentiation may be more synchronised. These results motivate future research on retailers' incentives to set prices less frequently. For instance, in a framework of costly price adjustment (e.g. menu cost), where pricing decisions are endogenous, the nature of cost shocks has been shown to hold implications for the timing of adjustment decisions across competition firms (e.g. Levy *et al.* 1998). In turn, state-dependent models of pricing predict a synchronised price response to cost shocks (Caballero and Engel 1993, Dotsey *et al.* 1999), while staggered price adjustments prevail in strategic responses to relative shocks (Ball and Romer 1989, Lach and Tsiddon 1996, Fisher and Konieczny, 2000).

7. Conclusions

Many studies that estimate cost pass-through rates or vertical price transmissions for agricultural and food product markets have focused on short-run dynamics and in particular on asymmetric 'rocket and feathers' adjustments of prices to cost shocks. Over the past decade, researchers have started to investigate and to test different theories that explain the variation in cost pass-through dynamics. These theories are built on models of market power, menu costs or consumer search costs. More widely available data at higher frequency have greatly supported economists' ability to test specific and intricate hypotheses on cost pass-through dynamics.

While theoretical models on the relationship between product differentiation and cost pass-through have been available for some time, to our knowledge, Loy and Weiss (2019) are the first to explicitly introduce a measure of product differentiation into a reduced-form model of retail cost pass-through. The analysis in this paper builds on Loy and Weiss (2019) and uses a non-address oligopolistic model approach to derive and empirically estimate the impacts of product differentiation on cost pass-through for industry-wide and firm-specific cost shocks. We test the theory of cost pass-through in differentiated oligopolies for a unique data set of ready-to-eat soup products in the Canadian retail market that includes detailed information on product-specific purchase prices, which allow us to differentiate between and estimate industry-wide and firm-specific costs.

Detailed UPC-level product and brand information allows us to construct and estimate several distance measures of product differentiation and to

merge these variables with a large panel of purchase and retail prices. We employ panel and up-to-date dynamic time-series methods to simultaneously estimate the impacts of product differentiation on the pass-through of industry-wide and firm-specific cost shocks.

The theory points to the importance of separating industry-wide and firm-specific cost shocks and their interaction with product differentiation in the context of firms' (retailers') price-setting behaviour. Acknowledging these interactions has important implications when making inference about the welfare outcomes of cost pass-through processes. The results in this paper suggest that product differentiation is associated with higher margins, reduced cost pass-through rates for industry-wide cost shocks, and slower adjustments of retail prices to such cost shocks. Firm-specific cost shocks increase the cost pass-through for more differentiated products.

These results underline the well-known fact that product differentiation serves as a market segmentation tool to reduce price competition and is a source of retail market power, which directly impacts product-specific markups and cost pass-through. This study thereby closes a gap between the fields of microeconomic modelling and marketing.

Finally, the limitations of our study provide opportunities for future research. We already discussed the high pass-through rates of firm-specific costs and the inherent limitations in how we measure product differentiation. Although our distance measure reflects a broad set of product available attribute data, consumer choice and therefore retailer product placement and pricing decision may be influenced by other, unobserved factors. As Carlton and Perloff (2015, p. 227) put it: 'The consumer is always right'. This implies that consumers may evaluate physically identical products differently and vice versa. Thus, an open question remains, whether our measure of product differentiation correctly reflects consumers' product perceptions. Consumer choice experiments may answer this question. The many results of hedonic price analysis in the literature, however, may suggest a close relationship between these product attributes and consumers' perception.

Data availability statement

The data were obtained from SIEPR-Giannini Data Center via a research agreement between the retail chain and co-author Sven Anders at the University of Alberta. The data are not available due to third party restrictions.

References

Angrist, J.D. and Pischke, J.S. (2010). The credibility revolution in empirical economics: How better research design is taking the con out of econometrics, *Journal of Economic Perspectives* 24, 3–30.

Atil, A., Lahiani, A. and Nguyen, D.K. (2014). Asymmetric and nonlinear pass-through of crude oil prices to gasoline and natural gas prices, *Energy Policy* 65, 567–573.

Bai, J. and Ng, S. (2002). Determining the number of factors in approximate factor models, *Econometrica* 70, 191–221.

Baker, J. and Bresnahan, T.F. (2008). Economic evidence in antitrust: defining markets and measuring market power, in Buccirossi, P. ed, *Handbook of Antitrust Economics*. MIT Press, Cambridge, MA, pp. 1-42.

Bakucs, Z., Falkowski, J. and Ferto, I. (2014). Does Market Structure Influence Price Transmission in the Agro-food Sector? A Meta-analysis Perspective, *Journal of Agricultural Economic* 65(1), 1–25.

Ball, L. and Romer, D. (1989). The equilibrium and optimal timing of price changes, *The Review of Economic Studies* 56, 179–198.

Bittmann, T. and Anders, S. (2016). Seasonal asymmetries in wholesale–retail cost pass-through, *Applied Economics Letters* 23, 1065–1068.

Bonanno, A. (2012). Functional foods as differentiated products: the Italian yogurt market, *European Review of Agricultural Economics* 40, 45–71.

Bonnet, C. and Villas-Boas, S.B. (2016). An analysis of asymmetric consumer price responses and asymmetric cost pass-through in the French coffee market, *European Review of Agricultural Economics* 43, 781–804.

Borenstein, S. and Shepard, A. (2002). Sticky prices, inventories, and market power in wholesale gasoline markets, *RAND Journal of Economics* 33, 116–139.

Borenstein, S., Cameron, A.C. and Gilbert, R. (1997). Do gasoline prices respond asymmetrically to crude oil price changes?, *The Quarterly Journal of Economics* 112, 305–339.

Breitung, J. and Das, S. (2008). Testing for unit roots in panels with a factor structure, *Econometric Theory* 24, 88–108.

Bulow, J.I. and Pfleiderer, P. (1983). A note on the effect of cost changes on prices, *Journal of Political Economy* 91, 182–185.

Caballero, R.J. and Engel, E.M. (1993). Microeconomic rigidities and aggregate price dynamics, *European Economic Review* 37, 697–711.

Carlton, D.W. and Perloff, J.M. (2015). *Modern Industrial Organization*, 4th Edn. Pearson, New York.

Cha, S.H. (2007). Comprehensive survey on distance/similarity measures between probability density functions, *International Journal of Mathematical Models and Methods in Applied Sciences* 1, 300–307.

Dixit, A. (1979). A model of duopoly suggesting a theory of entry barriers, *The Bell Journal of Economics* 10, 20–32.

Dolado, J.J. and Lütkepohl, H. (1996). Making Wald tests work for cointegrated VAR systems, *Econometric Reviews* 15, 369–386.

Dotsey, M., King, R.G. and Wolman, A.L. (1999). State-dependent pricing and the general equilibrium dynamics of money and output, *The Quarterly Journal of Economics* 114, 655–690.

Durevall, D. (2018). Cost pass-through in the Swedish coffee market, *European Review of Agricultural Economics* 45, 505–529.

Fisher, T.C.G. and Konieczny, J.D. (2000). Synchronization of price changes by multiproduct firms: evidence from Canadian newspaper prices, *Economics Letters* 68, 271–277.

Gaudin, G. (2016). Pass-through, vertical contracts, and bargains, *Economics Letters* 139, 1–4.

Gaudin, G. (2018). Vertical bargaining and retail competition: What drives countervailing power?, *The Economic Journal* 128, 2380–2413.

Gower, J.C. (1971). A general coefficient of similarity and some of its properties, *Biometrics* 27, 857–871.

Gower, J.C. and Warrens, M.J. (2017). Similarity, dissimilarity, and distance, measures of. Statistics Reference Online, Wiley StatsRef.

Hanck, C. (2008). The error-in-rejection probability of meta-analytic panel tests, *Economics Letters* 101, 27–30.

Haucap, J., Heimeshoff, U., Thorwarth, S. and Wey, C. (2015). Die Sektoruntersuchung des Bundeskartellamts zur Nachfragemacht im Lebensmitteleinzelhandel. Ein Kommentar aus ökonomischer Perspektive. DICE Consult GmbH, Düsseldorf. www.dice-consult.de.

Hausman, J.A. (1978). Specification tests in econometrics, *Econometrica* 46, 1251–271.

Holtz-Eakin, D., Newey, W. and Rosen, H.S. (1988). Estimating vector autoregressions with panel data, *Econometrica: Journal of the Econometric Society* 56, 1371–1395.

Ten Kate, A. and Niels, G. (2005). To what extent are cost savings passed on to consumers? An oligopoly approach, *European Journal of Law and Economics* 20, 323–337.

Kaufman, L. and Rousseeuw, P.J. (2009). *Finding Groups in Data: An Introduction to Cluster Analysis*. Wiley Series in Probability and Statistics, 344. John Wiley & Sons, Chichester, UK.

Kim, D. and Cotterill, R.W. (2008). Cost pass-through in differentiated product markets: The Case of U.S. processed cheese, *The Journal of Industrial Economics* 56, 32–48.

Krzanowski, W.J. (1993). The location model for mixtures of categorical and continuous variables, *Journal of Classification* 10, 25–49.

Lach, S. and Tsiddon, D. (1996). Staggering and synchronization in price setting: an empirical analysis of disaggregated data, *American Economic Review* 86, 1175–1196.

Lancaster, K.J. (1966). A new approach to consumer theory, *Journal of Political Economy* 74, 132–157.

Leamer, E.E. (1983). Let's take the con out of econometrics, *The American Economic Review* 73, 31–43.

Levy, D., Dutta, S., Bergen, M. and Venable, R. (1998). Price adjustment at multiproduct retailers, *Managerial and Decision Economics* 19, 81–120.

Lloyd, T. (2017). Forty years of price transmission research in the food industry: Insights, challenges and prospects, *Journal of Agricultural Economics* 68, 3–21.

Love, I. and Zicchino, L. (2006). Financial development and dynamic investment behavior: Evidence from panel VAR, *Quarterly Review of Economics and Finance* 46, 190–210.

Loy, J.-P. and Weiss, C. (2019). Product Differentiation and Cost Pass-Through, *Journal of Agricultural Economics*. https://doi.org/10.1111/1477-9552.12320.

Loy, J.-P., Holm, T., Steinhagen, C. and Glauben, T. (2015). Cost pass-through in differentiated product markets: a disaggregated study for milk and butter, *European Review of Agricultural Economics* 42, 441–471.

Loy, J.-P., Weiss, C.R. and Glauben, T. (2016). Asymmetric cost pass-through? Empirical evidence on the role of market power, search and menu costs, *Journal of Economic Behavior & Organization* 123, 184–192.

McFadden, D. (1974). The measurement of urban travel demand, *Journal of Public Economics* 3, 303–328.

Milligan, G.W. and Cooper, M.C. (1988). A study of standardization of variables in cluster analysis, *Journal of Classification* 5, 181–204.

Mintel (2015). Global New Product Database. Proprietary market database available athttp://gnpd.mintel.com.

Newey, W.K. and West, K.D. (1994). Automatic lag selection in covariance matrix estimation, *Review of Economic Studies* 61, 631–653.

Ng, S. and Perron, P. (1995). Unit root tests in ARMA models with data-dependent methods for the selection of the truncation lag, *Journal of the American Statistical Association* 90, 268–281.

Pesaran, M.H. (2004). General diagnostic tests for cross section dependence in panels. CESifo Working Paper No. 1229.

Pesaran, M.H. (2006). Estimation and inference in large heterogeneous panels with a multifactor error structure, *Econometrica* 74, 967–1012.

Pinske, J. and Slade, M.E. (2004). Mergers, brand competition, and the price of a pint, *European Economic Review* 48, 617–643.
Pinske, J., Slade, M. and Brett, C. (2002). Spatial price competition: A semiparametric approach, *Econometrica* 3, 1111–1153.
Richards, T.J., Gómez, M.I. and Lee, J. (2014). Pass-through and consumer search: An empirical analysis, *American Journal of Agricultural Economics* 96, 1049–1069.
Rojas, C. (2008). Price Competition in US Brewing, *The Journal of Industrial Economics* 56, 1–31.
Rojas, Ch and Peterson, E.B. (2008). Demand for differentiated products: Price and advertising evidence from the U.S. beer market, *International Journal of Industrial Organization* 26, 288–307.
Rosen, S. (1974). Hedonic prices and implicit markets: product differentiation in pure competition, *Journal of Political Economy* 82, 34–55.
Shubik, M. and Levitan, R. (1980). *Market Structure and Behavior*. Harvard University Press, Cambridge, MA.
Stanford Institute for Economic Policy Research (SIEPR) (2012). *SIEPR-Giannini Data Center*. Available from URL: http://are.berkeley.edu/sgdc.
Tappata, M. (2009). Rockets and feathers: Understanding asymmetric pricing, *The RAND Journal of Economics* 40, 673–687.
Toda, H.Y. and Yamamoto, T. (1995). Statistical inferences in vector autoregressions with possibly integrated processes, *Journal of Econometrics* 66, 225–250.
Walters, C., Williams, I., Durand, B., Majumdar, A., Alvim, N., Armstrong, R., Manganelli, A., Reuter, T., Sodano, V., Stoddart, P. and Whelan, C. (2014). Cost pass-through: theory, measurement, and potential policy implications. Report prepared for the Office of Fair Trading, London. RBB Economics. London 2014.
Wang, X.H. and Zhao, J. (2007). Welfare reductions from small cost reductions in differentiated oligopoly, *International Journal of Industrial Organization* 25, 173–185.
Westerlund, J. (2007). Testing for error correction in panel data, *Oxford Bulletin of Economics and Statistics* 69, 709–748.
Weyl, E.G. and Fabinger, M. (2013). Pass-through as an economic tool: Principles of incidence under imperfect competition, *Journal of Political Economy* 121, 528–583.
Zimmerman, P.R. and Carlson, J.A. (2010). Competition and cost pass-through in differentiated oligopolies. MPRA Paper No. 25931. Available form URL: http://mpra.ub.uni-muenchen.de/25931/.

Appendix

The impact of product differentiation in the Cournot (*sCN**) and Bertrand (sB*) oligopoly model by Shubik and Levitan (1980)

Shubik and Levitan (1980) start with a utility function of the following form. As above, we use the slope coefficient of 1 in the utility and demand function.

$$U = \alpha q - \frac{1}{2} q^2 \frac{2\sigma^2}{1+\phi} - \sum_i p_i q_i \tag{A1}$$

From this preference function, they derive the demand functions and calculate the price equilibria for the cost symmetric case under Cournot (*sCN**) in A2 and under Bertrand (*sB**) competition in A3).

$$p_i^{sCN*} = \frac{(n-\phi)\alpha + n(1-\phi)c}{2n-\phi(n+1)} \tag{A2}$$

$$p_i^{sB*} = \frac{n\alpha + c(n-(n-1)\phi)}{2n-\phi(n1)} \tag{A3}$$

Differentiation with respect to the marginal costs results in the cost pass-through in A4 and A5.

$$\frac{\partial p_i^{sCN*}}{\partial c} = \frac{n(1-\phi)}{2n-(1+n)\phi} > 0 \tag{A4}$$

$$\frac{\partial p_i^{sB*}}{\partial c} = \frac{n-(n-1)\phi}{2n-(n-1)\phi} > 0 \tag{A5}$$

Differentiation of A4 and A5 with respect to product differentiation (ϕ) results in the effect of product differentiation on cost pass-through rates in A6 and A7.

$$\frac{\partial^2 p_i^{sCN*}}{\partial c \partial \phi} = -\frac{n(n-1)}{(2n-(n+1)\phi)^2} < 0 \tag{A6}$$

$$\frac{\partial^2 p_i^{sB*}}{\partial c \partial \phi} = -\frac{n(n-1)}{(2n-(n-1)\phi)^2} < 0 \tag{A7}$$

Appendix

Table A1 Alternative random-effects panel cost pass-through model (equation (11))

Dependent variable: Retail prices Independent variables, Coeff.	Random Effects Euclidian,ϕ^E	Random Effects Sq. Diff.,ϕ^{SD}	Random Effects Block D.,ϕ^{BD}
Product differentiation (PD),β^{ϕ}	0.01 (0.01)	−0.01 (0.01)	0.03*** (0.01)
Industry-wide costs,$\beta^{\overline{w}}$	0.91*** (0.02)	0.89*** (0.02)	0.89*** (0.02)
Industry-wide c. * PD,$\beta^{\overline{w}\phi}$	0.05** (0.03)	0.09*** (0.03)	0.05*** (0.02)
Firm-specific costs,β^{w*}	0.96*** (0.05)	0.94*** (0.05)	0.98*** (0.05)
Firm-specific c. * PD,$\beta^{w*\phi}$	−0.12*** (0.03)	−0.10*** (0.03)	−0.12*** (0.03)
Constant,β^0	0.34*** (0.01)	0.36*** (0.01)	0.33*** (0.01)
R^2, within	0.83	0.84	0.83
R^2, between	0.90	0.90	0.90
R^2, overall	0.84	0.84	0.85
Observations	101,104	101,104	101,104

Note: Robust standard errors are reported in parenthesis. ***, ** and * denote 1%, 5% and 10% significance.
Source: Own calculation based on SIEPR (2012) and Mintel (2015).