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What price-level data can tell us about pricing conduct in South Africa

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Abstract

This paper reports the findings of a study of two unique microdatasets, which are the unit level basis of South Africa's measured consumer price index (CPI) and producer price index (PPI), over the period December 2001 to December 2007. In particular, the findings, which are based on techniques that have been used in comparable international studies, report on the frequency and magnitude of price changes, the duration of prices and heterogeneity in pricing. The results for South Africa are compared to the stylised facts for pricing conduct that have been presented in recent international studies. The paper offers an illustration of how microdata-based findings on pricing conduct may impact on the modelling of monetary policy by introducing micro-founded results into an open economy dynamic stochastic general equilibrium (DSGE) model for the South African economy. The paper concludes by identifying areas for further research, where it has not as yet been determined how South African pricing conduct compares to certain stylised pricing facts identified in the international literature.

JEL classification: E30, E31, D40, E02, E52

Keywords: monetary policy, pricing conduct, pricing heterogeneity, pricing microdata,

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Contents

1. Introduction

The study of pricing microdata is emerging in the literature as an important method for understanding actual pricing conduct. Studies of the large price datasets used to compile consumer price index (CPI) and producer price index (PPI) measures have been undertaken in a number of countries, including Israel (Baharad et al. 2004), Spain (Alvarez et al. 2004), France (Baudry et al. 2007), the United States (US) (Bils and Klenow 2005), Portugal (Dias et al. 2007), Germany (Stahl 2005), Luxemburg (Lunnemann 2005), Austria (Baumgartner et al. 2005), Sierra Leone (Kovanen 2006), Italy (Sabbatini et al. 2006), Denmark (Hansen et al. 2006), Brazil (Gouvea 2007), France (Gautier 2008), Finland (Kurri 2007), the euro area (Alvarez et al. 2008), Colombia (Julio and Zarate 2008) and Slovakia (Coricelli and Horvath 2010).

An alternative methodology for studying pricing conduct is to undertake surveys, such as the seminal survey of price-setters in the US by Blinder et al. (1998). Surveys of price-setting behaviour have also been conducted in France (Loupias and Ricart 2004), Sweden (Apel et al. 2005), Austria (Kwapil et al. 2005), Portugal (Martins 2005), Luxemburg (Lunnemann and Matha 2006), Canada (Amirault et al. 2006), Holland (Hoeberichts and Stockman 2006), Spain (Alvarez et al. 2008) and Turkey (Sahinoz and Saracoglu 2008). Some studies of price-setting conduct, utilising price datasets based on supermarket scanner data, have also been undertaken, such as in the US (Chevalier et al., 2000) and in the United Kingdom (UK) (Bunn and Ellis 2009). An earlier generation of price studies tended to be based on relatively narrow sets of products, such as Carlton (1986), Cecchetti (1986) and Kashyap (1995), rather than the larger, wider datasets underlying CPI and PPI data and supermarket data.

Klenow and Malin, in a chapter in the *Handbook of Monetary Economics* (2011, 232), present a review of ten stylised facts on price-setting conduct. This contribution brings some order to the taxonomy of pricing conduct by suggesting "ten facts modellers may want to know about price setting". These ten facts are listed as follows:[1]

1. Prices change at least once a year.
2. There is substantial heterogeneity in the frequency of price changes across goods.
3. Price changes are big on average, but many small price changes occur.
4. Price changes are typically not synchronised over the business cycle.
5. Neither frequency nor size is increasing in the age of the price.
6. Sales and product turnover are often important for micro price flexibility.
7. "Reference prices" are often stickier and more persistent than regular prices
8. More cyclical goods prices change more frequently.
9. Relative price changes are transitory.
10. Price changes are linked to wage changes.

[1] The list of ten facts has been drawn from Klenow and Malin (2011, 323), but the order in which the ten facts are presented has been altered to facilitate their comparison to the findings on the South African pricing microdata.

A key objective of this paper is to show, in a number of instances, how findings on pricing conduct from South Africa's pricing microdata compare with Klenow and Malin's stylised facts. In summary, there is clear confirmation that pricing conduct in South Africa is consistent with the first three stylised facts (1–3) and there is some evidence from the South African data that appears to contradict stylised facts 4 and 5. The questions raised by stylised facts 6–10 have not as yet been canvassed in current South African research.

In comparing South African pricing conduct to the stylised facts suggested by Klenow and Malin (2011), this paper is an advance on Creamer and Rankin's (2008) earlier contribution. This paper utilises improved estimates of the duration of prices, as it corrects for sampling bias towards short price spells as highlighted by Dias et al. (2007), and it introduces the micro-founded results into an open economy dynamic stochastic general equilibrium (DSGE) model rather than into a closed economy investment–saving (IS)–Phillips curve–Taylor Rule model.

2. Datasets used in this study

For both the CPI and PPI components of this study, the underlying price data that are gathered in the measuring of the CPI and PPI are used in order to develop an analysis of pricing conduct at a micro-level. For the CPI component of the study, a total of around 5 million individual price records were available over the period 2001m12 to 2007m12. Based on the criteria that only data collected at a monthly frequency[2] and only data with an acceptable capture status were to be included in the study,[3] the CPI microdataset ultimately used in this study comprised 3 930 977 price records. For the PPI data, a total of about 430 000 price records were available over the same 73-month period and, of these, a total of 381 861 prices were included in the study.[4] Owing to changes in the statistical authority's price collection methodology from a facsimile-based approach to one of direct price collection by enumerators over the period and owing to related systems changes, there is a break in the data at 2006m3. The reason for this break is that there is no equivalence in the numeric outlet codes for the two price collection systems. So, it is not possible to compare price changes at store level during the first month of the new collection methodology.

Each individual price record corresponds to a precisely defined item sold in a particular outlet (CPI) or enterprise (PPI) at a given point in time. Therefore, the pricing of individual items can be followed over time within the same outlet. Along with each individual price record, the following additional information is provided: the year and month of the record; the item code (indicating the type of product), a unit code

[2] Prices collected quarterly, annually or at other non-monthly intervals have been excluded from the data to be analysed. Housing prices were excluded due the fact that during the period under review, certain housing subsector price information – including the prices of rental stock of housing, flats and townhouses – was based on a frequently updated price index rather than on actual pricing conduct.
[3] Excluded were such capture codes as those for out-of-stock goods and for incomparable goods due to changes in quality.
[4] Similarly, for the PPI, price records classified as items out of stock or no longer sold and records classified as outliers were not included in the study.

(indicating the specific variety of the product), a capture code (indicating the capture status of the item), and a numeric outlet code (which, in terms of relevant legal confidentiality requirements, does not enable the name of the outlet to be identified, but which enables the tracking of pricing activity at specific anonymous outlets or enterprises). For the CPI, there are 1 124 goods and services in the CPI basket used in the period and the CPI basket is further divided into 18 product categories.[5] The PPI dataset contains pricing information divided into 23 Standard Industrial Classification (SIC) subcategories, and is categorised into local imported and exported goods.

Dias et al. (2007) raise the important concern that price microdatasets, of the type used in this study, tend to over-sample short price spells and under-sample long price spells. This leads to a particular problem in the measuring of average price durations as products that change price more frequently would have a greater number of price spells overall. To overcome this, this study calculates an average duration for each cross-sectional unit at the most disaggregated level possible (i.e., for a specific product sold in a specific unit at a specific outlet). These are then used to calculate average durations. While this method assists in overcoming some of the potential bias highlighted by Dias et al. (2007), average duration calculations may still display some downward biased since censored price spells are excluded by this method. The exclusion of such records may mean that longer, censored price spells are excluded due to the arbitrary limits of the data period available to the study. Fortunately, this sample selection bias does not affect the results for either the frequency or magnitude of price changes which, in addition to price duration estimates, are an important aspect of this study.

3. Pricing stylised facts and South Africa

3.1 Prices change at least once a year

On average, individual prices change more frequently than once per year. For the CPI microdata the average price duration is 5,0 months (median 3,7 months) and for the PPI microdata the average price duration is 6,1 months (median 4,5 months). Such price durations are based on the direct measurement of the duration of uncensored spells during which prices are unchanged.[6]

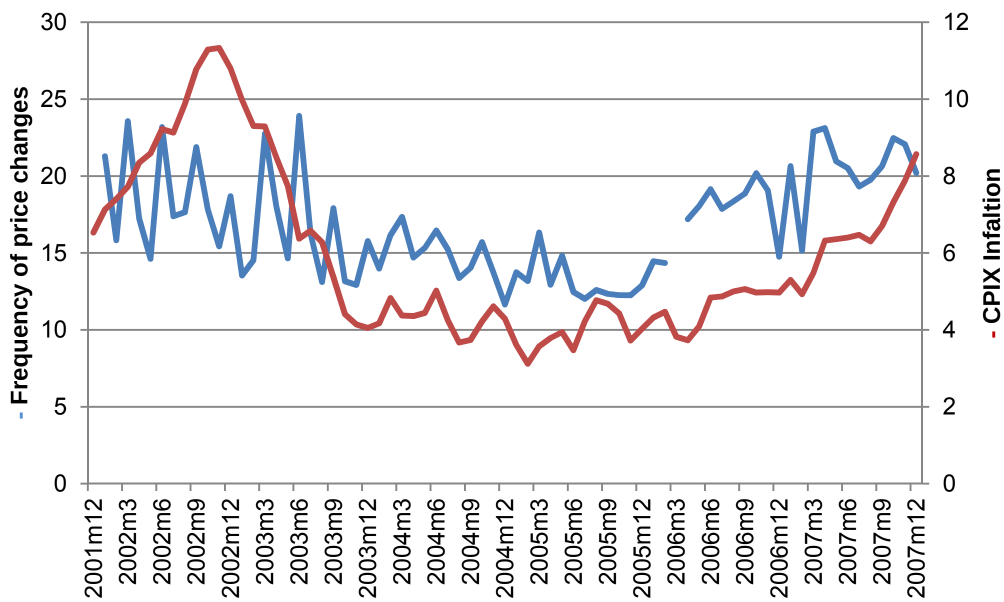
[5] From 2009 the basket of goods and services in the CPI has been reduced to fewer than 400 goods. Also from 2009 the Classification of Individual Consumption by Purpose (COICOP) system was introduced, which resulted in some changes to the categorisation of the CPI basket. At the same time, new weights for the goods and services in the CPI were introduced, based mainly on the consumption patterns found in the Income and Expenditure Survey of 2005/06.

[6] Censored price spells include those spells that are censored at the beginning or end of the spell by a break in the data, such as for monthly price records as follows: commence. R1,00. R2,00. R2,00. R2,00. R2,00. break (amounting to a censored price spell of four months, underlined). Whereas, uncensored price spells would not include the above as they would only include price spells that are not censored by a break in the data. Uncensored spells must include both the beginning and the end of the price spell, such as for monthly price records as follows: commence. R1. R2,00. R2,00. R2,00. R2,00. R3. break (amounting to an uncensored price spell of four months, underlined).

The CPI dataset offers evidence of a varying frequency of price changes, and related price durations, over time. This variation in the frequency of price changes over time is plotted in Figure 1, together with the series of the prevailing consumer price index excluding mortgage interest cost for metropolitan and other urban areas (CPIX) inflation rate. Over the period, the average monthly price change frequency is 16,8 per cent, meaning that each month an average 16,8 per cent of prices changes.[7] Re-weighting of the data according to the various product category weights underlying the CPI and CPIX results in an increased aggregate frequency of price adjustments to 17,1 per cent for both the CPI and CPIX weightings. Even though food items, with a relatively high price change frequency, are over-represented in the sample, the reweighting increases the weighting of transport, and other goods and services, which also have relatively high price change frequencies (outlined in Table 4).

Figure 1 shows that the frequency of price changes initially trended downwards until the end of 2005, and then began to rise.[8] As expected, such shifts in the frequency of price changes broadly follow the increases and decreases in the rate of inflation, with higher frequencies of price changes being associated with higher inflation rates and reduced frequencies of price changes being associated with lower inflation rates. Price stickiness would appear to increase as the rate of inflation falls.

Figure 1: Frequency of price changes per month and CPIX inflation
Per cent



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[7] This is based on the calculation, $f = (\sum_{i=1}^{N} (pc_{it}) / N$ where pc_{it} = 1 if $p_t \neq p_{t-1}$, or pc_{it} = 0 otherwise, f is the average monthly frequency of price change (pc_t) and N is the total number of observations.

[8] The highest frequency of price changes occurred in 2003m6 at 23,9 per cent and the lowest frequency of price changes occurred in 2004m12 at 11,6 per cent.

The CPI data also offer evidence of asymmetry in pricing as price increases (11,1 per cent on both the CPI and CPIX weightings) occur more frequently than price decreases (6 per cent).

The frequency of price changes at the aggregate level for South Africa (17,1 per cent) using CPI and CPIX weighted data, would appear to be broadly similar to findings for Spain (15 per cent), the euro area (15,1 per cent) and France (18,9 per cent). The US economy would appear to have a significantly greater frequency of price changes (24,8 per cent), including higher frequencies both of price increases and price decreases. Similarly, Brazil has experienced a significantly higher frequency of price changes (37 per cent) than South Africa, and a higher frequency of both price increases and price decreases. Price-setting conduct in South Africa bears little resemblance to high-inflation Sierra Leone where the frequency of price change averaged 51 per cent in the period between 1999 and 2003. In the case of Sierra Leone the frequency of price changes declined from 90 per cent in 1999 to about 40 per cent in 2003. A brief summary of selected comparative findings for CPI data is presented in Table 1.[9]

Table 1: Comparison of frequency of price changes from CPI microdata studies	**Frequency of price changes (%)**	**Frequency of price increases (%)**	**Frequency of price decreases (%)**
South Africa (2001–2007)	17,1	11,1	6,0
Euro area (1996–2001)	15,1	8,3	5,9
United states (1998–2003)	24,8	16,1	13,2
Spain (1993–2001)	15	9	6
France (1994–2003)	18,9	9,7	6,5
Brazil (1996–2006)	37	22,2	19
Sierra Leone (1998–2003)	51	20,1	21,4

An important technical reason for higher price change frequencies in the US data, as compared to the euro area and South Africa, is that the US data include a relatively large proportion of temporary sales prices, compared to other jurisdictions. Klenow and Malin (2011) report that in the US, 1 in 5 price changes relate to temporary sales prices, whereas in France 1 in 8 price changes relate to temporary sales prices. In other euro area countries, sales prices are not recorded in the pricing microdata.

In South Africa in the period under study, the statistical authorities initially gave no indication in the data as to whether or not prices were related to temporary sales, but from 2006m3 it is recorded for the CPI data whether or not a particular price record constitutes a sales price. If the US' experience is anything to go by, such a change in South Africa's price-recording methodology is a likely contributor to the higher price change frequencies evident in the second part of the period under review, as the current

[9] Data for South Africa are based on the results of the current study. Data for the euro area are from Dhyne et al., (2005) and for the US from Bils and Klenow (2004), and from Klenow and Kryvstov (2008). Data for Spain is from Alvarez et al. (2004), for France from Baudry et al. (2007), for Brazil from Gouvea (2007) and for Sierra Leone from Kovanen (2006). Owing to the adoption of differing methodologies in the various studies, not all results are strictly equivalent, yet the results allow for general comparisons of pricing conduct in a number of economies.

study does not differentiate between sales prices and non-sales prices. Results for the two sample periods are as follows: for the first period, the overall average frequency of price changes is 16 per cent and for the second period it is 19,6 per cent, with the price increase frequency being measured at 10,5 per cent and 12,1 per cent respectively and the price decrease frequencies being 5,5 and 7,5 per cent respectively.

For the PPI, there is evidence of a varying frequency of price changes and related price durations, over time, with an unweighted average monthly price change frequency of 20,2 per cent over the period. When price changes are weighted as per the weight given to each industrial sector in the PPI, then the aggregate frequency of price changes is given as 27,8 per cent.

As with the CPI data, the PPI data also reveal evidence of asymmetry in pricing, as price increases occur more frequently than price decreases. The frequency of price increases based on the PPI weightings is 16,4 per cent (or 12,2 per cent unweighted) and the frequency of price decreases is 11,4 per cent for the re-weighted data (or 8,1 per cent unweighted). A brief summary of selected comparative findings on the frequency of price changes for PPI data is presented in Table 2.[10]

Table 2: Comparison of findings on monthly price change frequencies using PPI microdata

	Frequency of price changes (%)	Frequency of price increases (%)	Frequency of price decreases (%)
South Africa (2001–2007)	27,8	16,4	11,4
Euro Area (various studies 1991–2005)	21,0	12,0	10,0
United States (1998–2005)	24,8	not reported	not reported
Spain (1991–1999)	21,0	12,0	9,0
France (1994–2005)	25,0	not reported	not reported
Colombia(1999–2006)	20,2	not reported	not reported

3.2 There is heterogeneity in the frequency of price changes

There is evidence of substantial heterogeneity in pricing conduct across goods and industries, and typically services prices change less frequently than goods prices, mainly due to the fact that services prices decrease less frequently than goods prices.

For the CPI data, the frequency of price changes for goods (17,0 per cent) is higher than that for services (14,9 per cent). In the CPI microdataset over 90 per cent of price records are classified as goods rather than services and, as a result, there is a close similarity between the findings for pricing conduct for goods and findings for the aggregate data. Table 3 indicates that the prices of services generally change less

[10] Sources for the table are: for South Africa from the current study, for the euro area from Vermeulen et al. (2007), for the US Nakumura and Steinsson (2008), for Spain Alvarez et al. (2008), for France from Gautier (2008), and for Colombia from Julio and Zarate (2008).

frequently than the prices of goods. However, services prices increase more frequently than goods prices, but decrease less frequently than goods prices.

Table 3: Average frequency of price changes for goods and services

	Aggregate	Goods	Services
Frequency of price changes	16,8	17,0	14,9
Frequency of price increases	10,8	10,8	11,4
Frequency of price decreases	6,0	6,1	3,5

For the CPI the findings on the frequency of price changes by product category show that there is significant heterogeneity in pricing conduct for different product categories. Table 4 shows, by product category, the average frequency of price changes, the average frequency of price increase and the average frequency of price decreases. For all products categories, except for footwear, the frequency of price increases is greater than the frequency of price decreases.

Table 4: Average frequency of price changes by product categories

Product category	Frequency of price change	Frequency of price increase	Frequency of price decrease
Food	20,5	13,1	7,5
Non-alcoholic beverages	13,1	9,2	3,8
Alcoholic beverages	12,5	9,7	2,9
Cigarettes tobacco and cigars	17,8	15,9	1,9
Clothing	8,8	5,0	3,8
Footwear	7,1	3,3	3,8
Fuel and power	14,7	10,5	4,2
Furniture and equipment	12,0	7,4	4,6
Household operation	14,7	9,9	4,8
Medical care and health expenses	17,2	10,6	6,6
Transport	19,3	12,2	7,1
Communications	6,7	5,5	1,2
Recreation and entertainment	13,8	7,5	6,3
Reading matter	13,6	12,1	1,5
Personal care	12,3	8,1	4,2
Other goods and services	25,2	17,6	7,6

For the PPI microdata, there is also heterogeneity in pricing across various product categories, with the frequency of price changes for imported products (23,2 per cent) being higher than for local products (18,8 per cent) and for exported products (18,7 per cent) as outlined in Table 5.

Table 5: Comparing price change frequencies for local, imported and exported products

Frequency	Aggregate (%)	Local products (%)	Imported products (%)	Exported products (%)
Frequency of price changes	20,23	18,75	23,15	18,71
Frequency of price increases	12,18	12,34	12,25	11,13
Frequency of price decreases	8,05	6,41	10,90	7,58

This is mainly due to the relatively high frequency of price decreases for imported goods (10,9 per cent) as compared to exported goods (7,58 per cent) and local products (6,41 per cent). The relatively high frequency of price decreases for imported goods is consistent with the currency appreciation during the period under review. The frequency of price increases for local (12,34 per cent) and imported goods (12,25 per cent) is very similar, with the frequency of price increases for exported products being somewhat lower (11,13 per cent), possibly hinting at competitive pressures facing exporters.

There is also evidence in the PPI microdata of pricing across industry sectors as outlined in Table 6, which shows the overall frequency of price changes, the average frequency of price increases and the average frequency of price decreases.

Table 6: PPI price changes by industrial sector

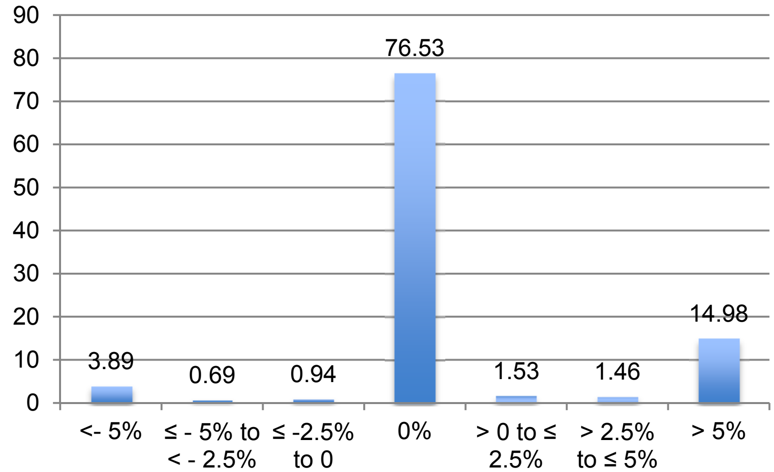
Sector	Frequency of price changes	Frequency of price increases (%)	Frequency of price decreases (%)
Agriculture	50,76	27,52	23,24
Forestry and fishing	11,91	10,99	0,92
Mining and quarrying	50,55	27,01	23,55
Food at manufacturing	26,11	15,65	10,46
Beverages	10,36	8,00	2,36
Tobacco products	16,89	12,33	4,57
Textiles and made-up goods	13,15	7,86	5,29
Wood and wood products	13,60	10,15	3,45
Paper, paper products and printing	21,33	11,88	9,45
Products of petroleum and coal	67,82	40,09	27,73
Chemicals and chemical products	16,55	10,26	6,30
Rubber and plastic products	15,25	10,44	4,81
Non-metallic mineral products	16,91	11,05	5,86
Basic metals	33,44	19,86	13,58
Metal products	15,94	9,96	5,99
Non-electrical machinery and equipment	15,46	9,06	6,40
Electrical machinery and apparatus	15,32	10,22	5,10
Radio, television, communications equipment and apparatus	13,41	7,16	6,25
Transport equipment	18,23	10,63	7,60
Furniture	8,32	6,80	1,52
Other manufactures	11,14	6.92	4.22
Electricity	46.90	30,19	16,71
Construction	26,49	14,73	11,76

3.3 Price changes are big, on average, but many small price changes occur

The magnitude of price changes is defined as the absolute value of the one-period difference, month on month, of the natural log of prices. The magnitude of price increases is the differenced natural log of price changes that are greater than zero and the magnitude of price decreases is the differenced natural log of price decreases.[11]

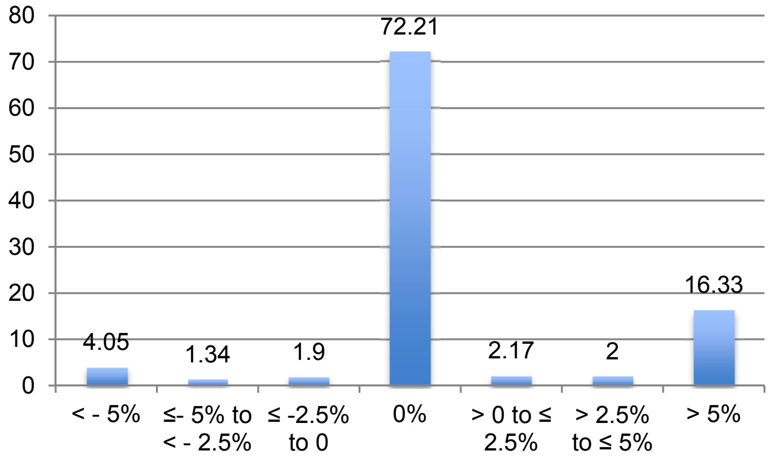
From the histograms presented in Figures 2 and 3, it can be seen that for most observations prices do not change, in the CPI data sample 76,5 per cent of prices indicate no change on the previous period's price and for the PPI data sample 72,2 per cent of prices do not change. The histograms have relatively fat tails, providing evidence of the proposition that where price changes do occur, they are big, on average, with about 18,9 per cent of the CPI data showing price change magnitudes of more than 5 per cent in absolute terms. Similarly, for the PPI, price change magnitudes of more than 5 per cent in absolute terms occur for around 20,4 per cent of the price records.

Figure 2: Histogram of magnitude of price changes (CPI)



[11] The average monthly magnitude of price increases (M^+) is calculated as follows:
$M^+ = \sum_{i=1}^{N} (\ln p_{i,t} - \ln p_{i,t-1}) / N$, for $\ln p_{i,t} - \ln p_{i,t-1} > 0$ where *p* is the magnitude of the price of a specific item at a specific store, and *N* is the number of observations where price magnitudes increase (as price changes of 0 are not included in the calculation of average price change magnitudes).

Figure 3: Histogram of magnitude of price changes (PPI)



South Africa's pricing microdata reveal relatively large magnitude price changes. For the CPI, the weighted average magnitude of price increases is 10,7 per cent, compared to an average monthly CPI inflation of 5,4 per cent over the period. For the PPI, the average weighted magnitude of price increase is 14,7 per cent compared to an average monthly PPI inflation of 6,8 per cent over the period.

For the CPI, comparative analysis is facilitated through Table 7, which shows that while the average magnitude of price increases in South Africa is 10,7 per cent, this is larger than for the euro area (8,2 per cent) and Spain (8,2 per cent). The average size of South Africa's price increases (10,7 per cent) is smaller than those of Brazil (16 per cent), the US (12,7 per cent) and France (12,5 per cent). The average size of price decreases in South Africa (12,3 per cent) is similar to the average size of price decreases in Brazil (12,6 per cent), is larger than for the euro area (10 per cent), France (10 per cent) and Spain (10,3 per cent), but smaller than for the US (14,1 per cent).

Table 7: Comparison of magnitude of price changes from CPI microdata studies

Country	Average size of price increases (%)	Average size of price decreases (%)
South Africa (2001–2007)	10,7	12,3
Euro area (1996–2001)	8,2	10
United states (1998–2003)	12,7	14,1
Spain (1993–2001)	8,2	10,3
France (1994–2003)	12,5	10
Brazil (1996–2006)	16	12,6

There is evidence in the South African data of many small price changes, where price changes are decomposed into those price changes that are partially indexed to the prevailing rate of inflation and those prices that are reoptimised. A possible method for decomposing price changes into price reoptimisations, on the one hand, and inflation-indexed price changes, on the other, is to identify as price reoptimisations those price changes that have a price change magnitude *above* a certain 'indexation threshold' and to identify as inflation-indexed price changes those prices that have a price change magnitude *below or equal to* a certain 'indexation threshold'.[12]

Where the 'indexation threshold' (δ) is based on $\delta = 0{,}25$ (as is the case in Steinbach et al.'s (2009) open economy DSGE model), 10,56 per cent of all price changes in the CPI microdata can be described as inflation-indexed price changes which, by definition, have a magnitude of less than or equal to 25 per cent of the prevailing CPI inflation rate. For the imported component of the PPI microdata, where the 'indexation threshold' is based on $\delta = 0{,}25$, 12,53 per cent of all price changes have a magnitude of less than, or equal to, 25 per cent of the prevailing PPI inflation rate. If the 'indexation threshold' is raised to $\delta = 0{,}5$, then for the CPI data, the number of price changes below that threshold rises to 15,4 per cent of all price changes and for the imported component of PPI to 18,14 per cent. If the 'indexation threshold' is raised to $\delta = 1$ (i.e., the magnitude of price changes is less than, or equal to, the prevailing rate of inflation) then 25,4 per cent of CPI prices and 25,5 per cent of PPI prices fall within the threshold.

Klenow and Malin (2011) suggest that such a finding – of a large number of small price increases – may be at odds with menu cost models and may favour time-dependent or information-constrained pricing models.

3.4 Price changes are typically not synchronised over the business cycle

Klenow and Malin's (2011) fourth stylised fact may be contradicted by the South African price microdata, which reveal some evidence that price changes are synchronised, in that the frequency of price increases rises with the prevailing rate of inflation. Using a basic regression model,[13] it was found that for the South African CPI price data, over the period 2001m12 to 2007m12, the frequency of price changes and price increases was positively and significantly associated with current CPI inflation and with the CPI inflation rate after a three-month lag. These regression results are reported in Table 8.

[12] This is not a method without its flaws, but it serves as a tractable and useful proxy measure of the degree of indexation that is prevalent in overall pricing conduct. One weakness with the method is that in certain circumstances the non-change of prices, or small price changes below the 'indexation threshold', may, in fact, constitute an optimal (or profit-maximising) pricing strategy. However, the proposed decomposition method would (by definition) classify such price changes as indexations rather than re-optimisations. By contrast, the proposed decomposition method has the advantage that it is relatively easy to implement, it correctly treats large magnitude price changes as non-indexed price changes, suggesting that such price changes represent price reoptimisations, and it offers a useful insight into how information on pricing conduct, as revealed by pricing microdata, can be used to guide the parameterisation of economic models.

[13] $F_t / M_t = a + \sum_{i=1}^{12} B + \kappa CPI_t + \varepsilon_t$ where F_t is variously the frequency of price changes, increases or decrease and M_t is the magnitude of price changes, increases or decreases.

Table 8:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Frequency of price changes				Frequency of price increases			
January	0.010	0.007	0.008	0.006	0.022	0.018	0.020	0.018
	(0.60)	(0.36)	(0.45)	(0.33)	(1.76)*	(1.17)	(1.35)	(1.11)
February	-0.004	-0.007	-0.006	-0.008	0.004	0.001	0.002	0.001
	(0.23)	(0.35)	(0.34)	(0.39)	(0.36)	(0.09)	(0.14)	(0.05)
March	0.051	0.048	0.050	0.050	0.055	0.052	0.054	0.054
	(2.96)***	(2.40)**	(2.67)***	(2.44)**	(4.32)***	(3.14)***	(3.56)***	(3.17)***
April	0.019	0.015	0.018	0.016	0.024	0.019	0.022	0.021
	(1.14)	(0.77)	(0.99)	(0.82)	(1.95)*	(1.24)	(1.55)	(1.30)
May	0.011	0.008	0.010	0.008	0.014	0.011	0.013	0.012
	(0.65)	(0.40)	(0.55)	(0.42)	(1.18)	(0.71)	(0.93)	(0.74)
June	0.039	0.036	0.038	0.036	0.031	0.028	0.030	0.029
	(2.37)**	(1.89)*	(2.16)**	(1.87)*	(2.56)**	(1.78)*	(2.11)**	(1.76)*
July	0.009	0.006	0.009	0.007	0.012	0.008	0.011	0.009
	(0.55)	(0.30)	(0.49)	(0.35)	(0.95)	(0.51)	(0.77)	(0.56)
August	0.003	0.001	0.003	0.002	0.005	0.003	0.005	0.004
	(0.18)	(0.03)	(0.18)	(0.09)	(0.45)	(0.18)	(0.38)	(0.26)
September	0.021	0.019	0.021	0.020	0.022	0.020	0.022	0.021
	(1.25)	(1.01)	(1.18)	(1.01)	(1.78)*	(1.28)	(1.52)	(1.27)
October	0.013	0.011	0.014	0.012	0.018	0.016	0.019	0.017
	(0.78)	(0.59)	(0.76)	(0.64)	(1.49)	(1.04)	(1.30)	(1.07)
November	0.003	0.003	0.003	0.002	0.009	0.009	0.009	0.008
	(0.16)	(0.15)	(0.16)	(0.12)	(0.70)	(0.56)	(0.60)	(0.50)
CPI_t	0.005				0.005			
	(4.83)***				(6.97)***			
CPI_{t-3}			0.004				0.004	
			(3.52)***				(4.34)***	
ΔCPI_t		0.011				0.012		
		(1.91)*				(2.62)**		
ΔCPI_{t-3}				0.005				0.007
				(0.80)				(1.51)
Constant	0.127	0.156	0.134	0.156	0.060	0.091	0.069	0.091
	(9.60)***	(11.63)***	(9.47)***	(11.35)***	(6.13)***	(8.23)***	(6.02)***	(7.92)***
Observations	71	71	71	71	71	71	71	71
R-squared	0.42	0.24	0.33	0.20	0.57	0.29	0.40	0.24

Absolute value of *t* statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

Klenow and Malin (2011) note that in the US there is little evidence of synchronisation as movements in inflation are associated with variations in the magnitude of price changes rather than variations in the frequency of price adjustments. In Mexico, however, in the context of volatile inflation, the frequency of price changes has shown meaningful synchronisation across firms (drawing on Gagnon, 2009).[14]

3.5 Neither frequency nor size is increasing in the age of the price

An estimated hazard function shows the probability of a change in a price, conditional on the price having been unchanged for a certain number of periods.[15] An upward-sloping hazard function would indicate that within a certain timeframe the longer the period that has passed since a price has changed, the greater is the likelihood that the price will change. For a variety of specific consumer products at specific stores, hazard functions are broadly upward-sloping, indicating the expected result that within a certain timeframe the likelihood of price changes increases with the passing of time.

At the aggregate level, the South African evidence for the CPI and PPI microdata (as outlined in Figures 4 and 5) is in line with Klenow and Malin's (2011, 276) finding for the US and euro area that "the hazard rate of price changes is falling over the first few months . . . . and [is] largely flat thereafter". As explained by Alvarez et al. (2005b, 9), such a downward-sloping hazard function at the aggregate level is a likely result of heterogeneity in pricing conduct. Intuitively, this is because "[t]he probability of observing price changes is lower for firms with sticky price schemes than for firms following flexible pricing rules, while the aggregate hazard considers price changes for all firms. Therefore, when the aggregate hazard function is obtained, the share of price changes corresponding to firms with more flexible pricing rules decreases as the horizon increases and, consequently, the hazard rate also decreases."

For example, the hazard functions in Figures 4 and 5 show the probability of a price change for each price duration, for the CPI dataset and PPI datasets respectively.[16] The

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[14] Klenow and Kryvstov (2008) decompose monthly inflation into the fraction (fr_t) of items with price changes and the average size (sz_t) of those changes, that is, $\pi_t = fr_t \times sz_t$. They find that for the US between 1988 and 2004, movements in inflation are mainly due to changes in the magnitude of price changes, rather than the changes in the frequency of price changes. Gagnon (2009) further decomposes inflation into terms due to price increases and price decreases such that $\pi_t = fr^+ sz^+ + fr^- sz^-$, where $fr = fr^+ + fr^-$, and fr^+ and fr^- (sz^+ and sz^-) denote the frequency (absolute size) of price increases and price decreases, respectively. Using this technique, the finding for Mexico from 1994 to 2002 is that when the annual inflation rate was below 10–15 per cent, the average frequency (size) of prices changes co-moves weakly (strongly) with inflation due to offsetting movements in the frequency of price increases and decreases. By contrast, when inflation rose above 15 per cent, a few price decreases were observed and both the frequency and the average size of price changes were found to be important determinants of inflation.

[15] More formally, where the hazard rate *(h)k* is expressed as the probability that a price (p_t) will change after *k* periods conditional on it having remained constant during the previous *k-1* periods, that is, $h(k) = Pr\{p_{t+k} \neq p_{t+k-1} | p_{t+k-1} = p_{t+k-2} = \ldots = p_t\}$.

[16] Both censored price spells and uncensored price spells are used in deriving the hazard functions in Figures 4 and 5. Censored price spells include those spells that are censored at the beginning or end of the spell by a break in the data, such as, for monthly price records, as follows: commence. R1,00. R2,00,

hazard functions are downward-sloping with an uptick at 12 months. The slope of the hazard functions indicates heterogeneity in pricing conduct and the uptick in the probability of price changes at 12 months indicates a degree of time-dependence in pricing conduct, in the form of annual pricing for consumer prices. The small peaks at between 40 and 50 months is based on a very small sample, and is unlikely to indicate any significant generalised finding on pricing conduct in South Africa.

Figure 4: Hazard function (CPI data)

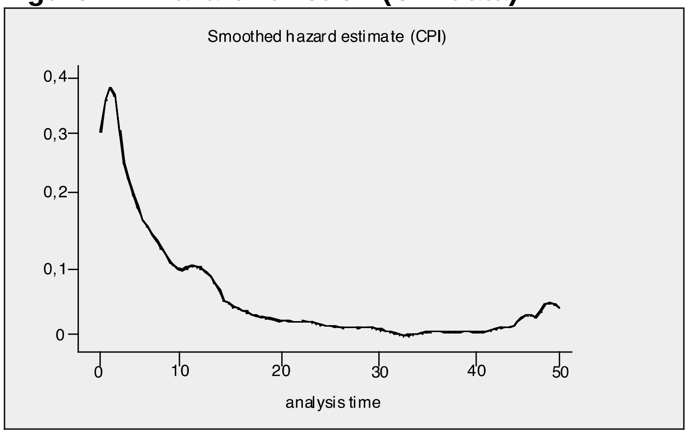
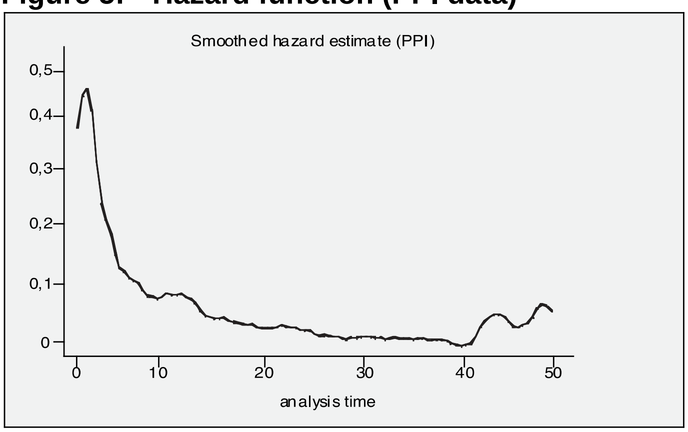


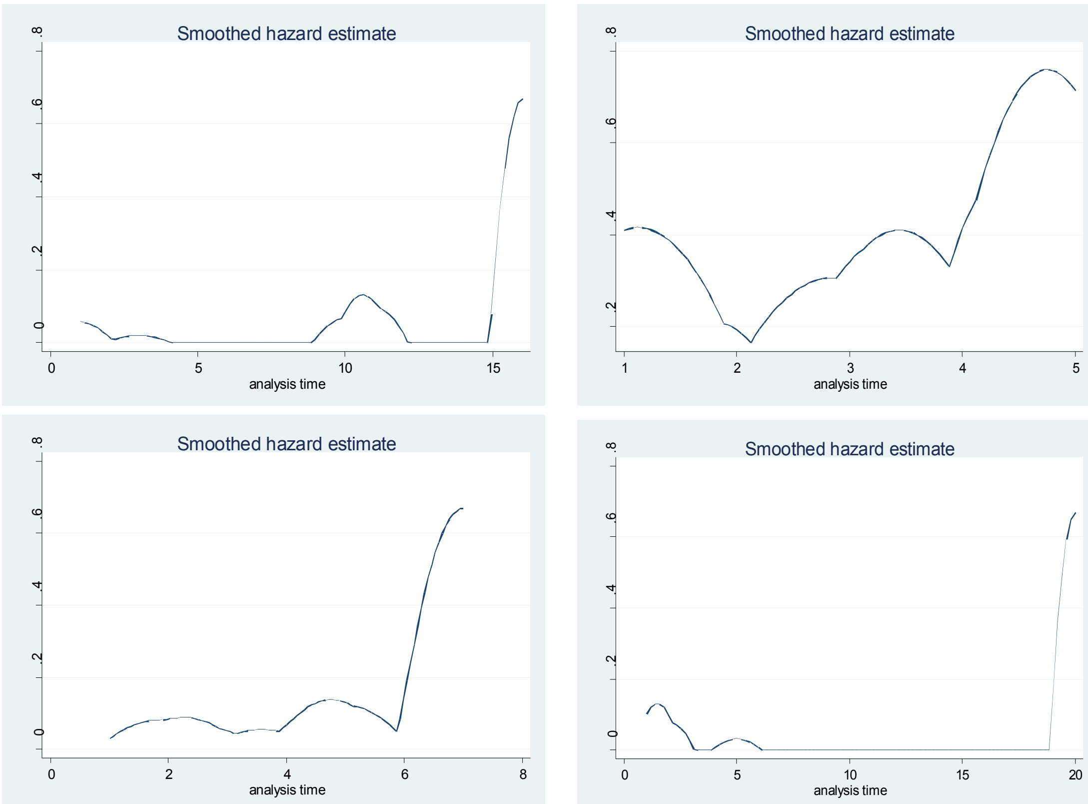
Figure 5: Hazard function (PPI data)



At the level of specific products at specific stores, hazard functions tend to be upward-sloping, indicating an increasing likelihood of price changes as time passes. Klenow and Malin (2011, 276) suggest that such rising hazard functions (see examples in Figure 6) may be indicative of state-dependent pricing as over time "shocks accumulate and the desired price level drifts further away from the current price".

R2,00, R2,00, R2,00. break (amounting to a censored price spell of four months, underlined). Whereas, uncensored price spells would not include the above as they would only include price spells that are not censored by a break in the data. Uncensored spells must include both the beginning and the end of the price spell, such as, for monthly price records, as follows: commence, R1. R2,00, R2,00, R2,00, R2,00, R3. break (amounting to an uncensored price spell of four months, underlined).

Figure 6: Disaggregated hazard functions for specific products (clockwise from top left: brown bread, oranges, imported whiskey and instant coffee)



With regard to the magnitude of price changes, Klenow and Malin's (2011, 276) fifth stylised fact that there is "little connection between the size of price changes and the duration of price spells", seems to be clearly contradicted by the evidence from the South African microdata which shows that the magnitude of CPI and PPI price changes is increasing in the age of the price (see Figures 7 and 8).[17] This finding offers evidence of time dependence in pricing conduct in the South African context.

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[17] These figures are descriptive and are the median size of a price change conditional on a price change and a price duration of a specific length (i.e., the median value of a price change for all those prices with a duration of one month that change).

Figure 7: Median size of price change (absolute %) against duration (in months) since previous price change for CPI dataset

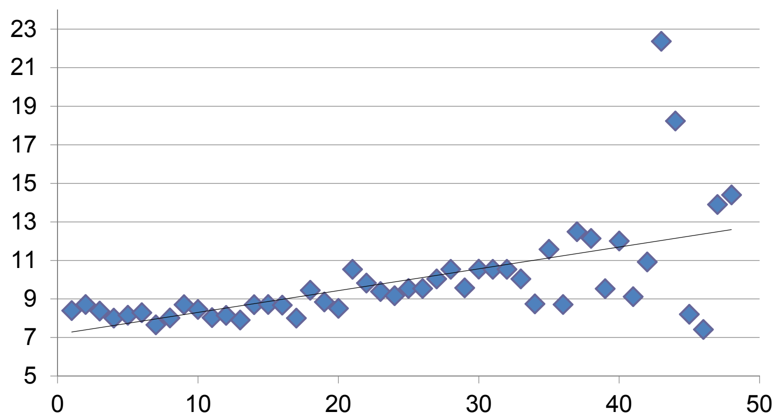
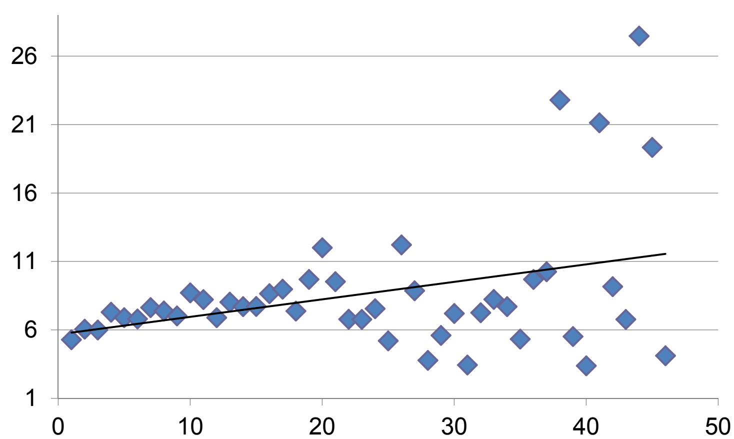


Figure 8: Median size of price change (absolute %) against duration (in months) since previous price change for PPI dataset



4. Potential modelling implications of findings on pricing conduct

An illustration of how a micro-founded understanding of pricing conduct for South Africa can usefully be integrated into macroeconomic modelling is provided by adjusting the pricing assumptions of the open economy DSGE model of the South African economy developed by Steinbach et al. (2009). This is a two-country New Keynesian DSGE model, with South Africa the domestic economy and the foreign economy being the rest of the world. South Africa is modelled as a small open economy along the lines proposed by Monacelli (2003), and Justiniano and Preston (2004), with incomplete pass-through of exchange rate changes and real rigidity in the form of external habit formation in consumption. Furthermore, and particularly important here, the model features Calvo (1983) price and wage setting, partial indexation of domestic prices to past inflation, and partial indexation of wages to past consumer price inflation. The foreign economy is assumed to be large in that it is not affected by developments in South Africa. Bayesian estimation techniques are used to estimate pricing conduct.

In this section we show how it is possible to incorporate direct measures of pricing conduct from our microdata studies into the model, with the ultimate finding that the microdata evidence contradicts the assumptions of the DSGE model in the South African case.

In Steinbach et al.'s model, the New Keynesian Phillips curve for domestic inflation (π_h) is derived through a combination of an optimal price-setting rule and the price level:

$$\pi_{h,t} = \frac{\delta}{1+\delta\beta}\pi_{h,t-1} + \frac{\beta}{1+\delta\beta}\pi_{h,t+1} + \frac{(1-\theta_h)(1-\theta_h\beta)}{\theta_h(1+\delta\beta)}mc_t$$

Nominal rigidity enters into the system, as it is assumed that intermediate goods firms set prices in a staggered manner as per Calvo (1983), whereby in each period *t*, each firm is allowed to reset its price with a probability of (1 – θ_h). Therefore, in a given period *t*, not all firms are able to react to shocks immediately. The higher is θ_h, the stickier are prices (the lower the frequency of price changes), as a smaller number of firms, given by (1 – θ_h), are able to reoptimise their prices in each period. The result of $\frac{(1-\theta_h)(1-\theta_h\beta)}{\theta_h(1+\delta\beta)}$ is that for higher values of θ_h, shocks in marginal cost (mc_t) will result in smaller changes to domestic inflation (π_h) as prices are relatively sticky. For lower levels of θ_h, prices are less sticky and increases in marginal cost will result in larger inflationary effects. In the special case where prices are flexible, that is, θ_h = 0, then all firms would be able to change their prices. By using the findings of the price microdata study, it is possible to set θ_h at different levels in order to use the model to assess the impact of varying degrees of price stickiness.[18]

In order to understand the comparative dynamic effects of various pricing scenarios, the open economy DSGE model can then run using alternative parameter levels for θ_h and θ_f. For the CPI microdata, the average price duration is measured at 5,0 months, compared to the DSGE model's estimate of an average duration of 6,5 months for domestic consumer goods.[19] For the imported component of PPI microdata, the

[18] Overall CPI inflation is based on a combination of domestic inflation (π_h) and imported inflation (π_f), which is log-linearised as follows: $\pi_t = (1-\gamma)\pi_{h,t} + \gamma\pi_{f,t}$. Imported inflation ($\pi_f$) can be expressed as a Phillips curve-type relationship, including the import price index and optimal price setting by importing retailers, where higher price stickiness among importing retailers is indicated by higher values for θ_f:

$$\pi_{f,t} = \beta\pi_{f,t+1} + \frac{(1-\theta_f)(1-\theta_f\beta)}{\theta_f}\psi_{f,t}.$$

[19] In this instance the result for the entire sample of the CPI microdata is used to estimate θ_h as the disaggregation of the CPI microdata into a domestic and an imported component is not provided for in the

average price duration is measured at 6,1 months, compared to the DSGE model's estimate of 9,14 months average price duration for imported price. It is then possible to compare the open economy DSGE model's "baseline" scenario – where θ_h = 0,539 and θ_f = 0,672 with the "alternative" scenario, informed by the pricing microdata, where: θ_h = 0,4 and θ_f = 0,5.[20]

The implication of the shorter measured durations is that due to the higher degree of price flexibility implied by the data, the conduct of interest rate policy should generally be *more aggressive, but less persistent,* in response to a range of positive and negative shocks, than the impulse responses implied by the price conduct parameters estimated by Steinbach et al. (2009); for example, Figure 9 offers a graphic indication of the impulse response implications for a range of macroeconomic variables, in the context of a positive cost shock, depending on which pricing assumption is used, that is, either the pricing conduct estimated using Bayesian techniques (the 'baseline' scenario), which is the same as that utilised in Steinbach et al. (2009), or the pricing conduct measured from the CPI and PPI microdata (the 'alternative' scenario).

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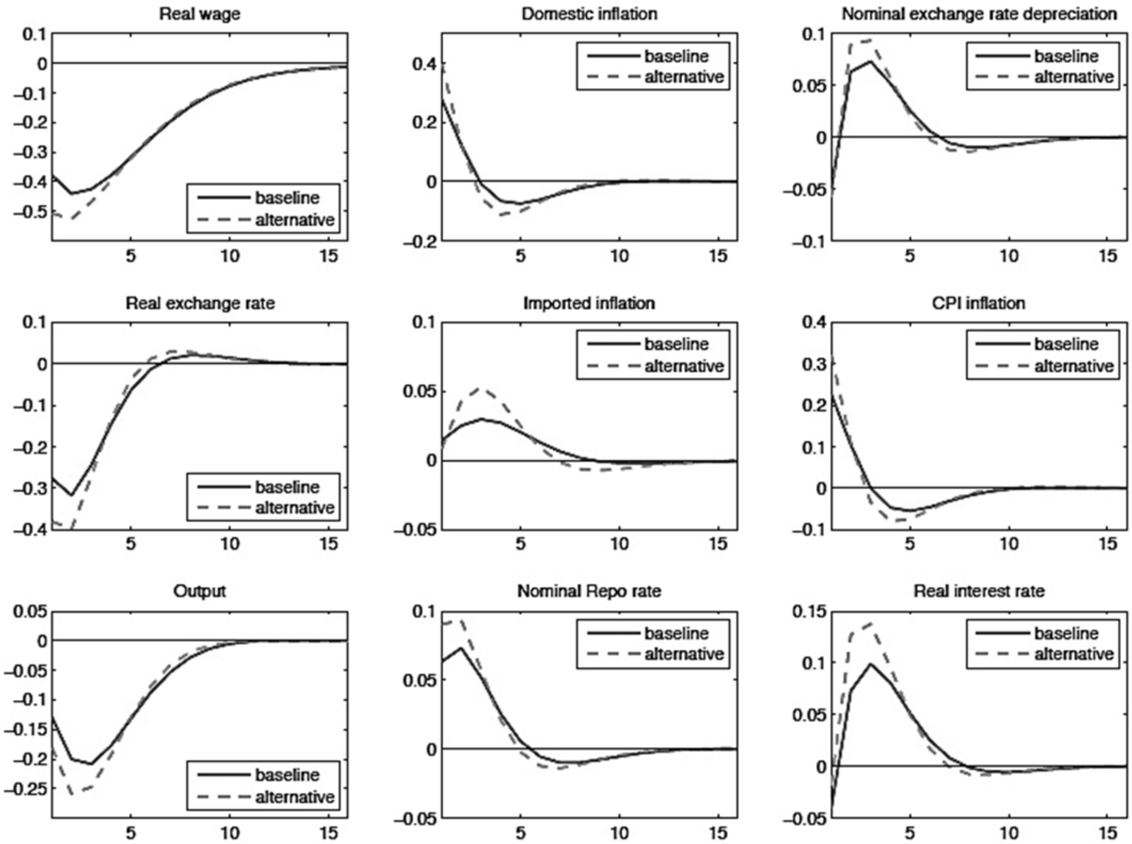
underlying data, although a study by Blignaut et al. (2006) estimates that imports make up around 15 per cent of the CPIX and around 14 per cent of the CPI, implying that about 85 per cent (86 per cent) of the movement in the CPIX (CPI) is accounted for by domestic factors.

[20] The open economy DSGE model estimates Calvo parameters for domestic firms (θ_h) at 0,539 and for importing firms (θ_f) at 0,672. The inverses of (1-θ_h) and (1-θ_f) are used to calculate the monthly duration of prices from the Calvo parameter. This translates to an average price duration of $(1-0{,}539)^{-1}$, equalling 2,2 quarters (6,5 months) for domestic prices and average price duration $(1-0{,}672)^{-1}$, equalling 3,1 quarters (9,14 months) for imported prices. Therefore, the price microdata offer evidence that, on average, measured price durations are shorter than the mean price durations estimated in Steinbach et al.'s (2009) open economy DSGE model. In summary, measured prices are less sticky than those used in the DSGE model. For the DSGE model, pricing by domestic firms (θ_h) estimates a 2,2 quarter average price duration (where θ_h=0,539), whereas from the CPI microdata the average price duration for domestically producing firms is about 1,6 quarters (or θ_h= 0,4). For the DSGE model, pricing by importing firms (θ_f) estimates a 3,1 quarter average price duration (where θ_f= 0,672), whereas from the PPI microdata the average price duration for importing firms is about 2,0 quarters (or θ_f= 0,5).

Figure 9: Comparing the 'baseline' scenario with 'alternative' micro-founded scenario in the context of a positive price shock or cost shock



In response to a positive price shock or cost shock, the impulse responses of the 'alternative' scenario (as compared to the 'baseline' scenario) reveal the following:

- For all measures of inflation (i.e., domestic, imported and overall CPI) there is a sharper, but less persistent, increase in inflation in the 'alternative' scenario as compared to the 'baseline' scenario. In all cases there is also some overshooting before the inflation rate returns to its steady state.
- The decline in output is sharper and marginally less persistent in the 'alternative' scenario as compared to the 'baseline' scenario.
- There is an initial appreciation, followed by an overshooting depreciation of the nominal and real exchange rates, with the overshooting depreciation of the nominal exchange rate being significantly sharper in the 'alternative' scenario as compared to the 'baseline' scenario.
- A decrease in real wages is a feature of both the 'alternative' scenario and the 'baseline' scenario, although there is no significant difference in the degree of persistence of the decrease in real wages between the two scenarios, the decrease is sharper in the 'alternative' scenario as a result of the sharper inflation response.
- The increase in the repo rate (the policy rate) and real interest rate is of a significantly larger magnitude in the 'alternative scenario' and is marginally less

persistent, mainly due to the fact that the ‘alternative’ scenario’s greater price flexibility results in less persistent output and inflation responses.

This example – and others such as the differential responses of the ‘baseline’ and ‘alternative’ scenarios to demand shocks, interest rate shocks and wage shocks – indicates the potential role that price microdata studies could play in deepening the understanding of pricing conduct and the modelling of such conduct.

Furthermore, the microeconomic evidence regarding the heterogeneity of pricing conduct raises important issues for macroeconomic modelling and the evaluation of monetary policy. Very few existing pricing models in the monetary policy literature allow for pricing heterogeneity. Alvarez (2008) reviews 25 pricing models, which include sticky information, menu cost, time-dependent, cost of adjustment and customer anger models, and finds that only four allow for price heterogeneity. It has been argued that (microdata-implied) heterogeneous pricing conduct by firms is unlikely to be adequately captured by standard models that assume a ‘representative firm’, and attempts to calibrate the latter to models with many heterogenous firms “may be wishful thinking” (Maćkowiak and Smets 2008, 7). Furthermore, models that attempt to incorporate price heterogeneity into the analysis of monetary policy suggest that neglecting heterogeneity has a significant quantitative effect.

The Calvo (1983) pricing assumption, often modified to incorporate “dynamic indexation” to generate inflation persistence in quantitative DSGE models, faces a particularly stern challenge from the microdata evidence on the frequency of price adjustment. The Calvo parameters estimated from New Keynesian Phillips curves imply levels of price rigidity that are generally inconsistent with the microdata,[21] leading some economists to argue that the profession’s thinking about pricing should not be focused on the Calvo model (King, 2009).

5. Conclusion: Identifying areas for further research

Further research is required to test whether there is evidence in the South African pricing microdata to support the remaining four of Klenow and Malin’s (2011) stylised facts (6–10). Such a research agenda would assist in answering, among other things, the following questions:

– How important are temporary sales prices in micro-level price flexibility, do they play an important role in South Africa, as in the US, or a lesser role, as in the countries in the euro area?
– If a broad set of short-lived prices (including temporary sales prices) is excluded from the data, does a stickier ‘reference’ price emerge that changes about once per year, as in the US data?

[21] For the US, for example, compare Eichenbaum and Fisher (2007) to the microdata evidence from Bils and Klenow (2004) and Nakamura and Steinsson (2008).

– Is there evidence in the South African data that cyclical goods (such as cars and clothing) exhibit greater micro price flexibility than goods that display less-cyclical behaviour (such as medical care)?
– Is there evidence in South Africa that relative price changes are transitory, as this will shed light on the relative importance of idiosyncratic versus aggregate shocks and will have implications for price-setting models?
– What linkages exist between changes in prices and change in wages?

Answers to these questions would undoubtedly contribute further to our knowledge of pricing conduct in the South African economy and, armed with such knowledge, further refinements could be made in the arenas of macroeconomic modelling and policy implementation.

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