



School of Business  
Te Kura Pakihi

ISSN 1178-2293 (Online)

University of Otago  
Economics Discussion Papers  
No. 1107

August 2011

---

# Determinants of Relative Price Variability during a Recession: Evidence from Canada at the Time of the Great Depression

David Fielding,<sup>\*§</sup> Chris Hajzler<sup>§</sup> and Jim MacGee<sup>¶</sup>

---

\* Corresponding author. Address for correspondence: Department of Economics, University of Otago, PO Box 56, Dunedin 9054, New Zealand. E-mail [david.fielding@otago.ac.nz](mailto:david.fielding@otago.ac.nz); telephone +6434798653.

§ Department of Economics, University of Otago.

¶ Department of Economics, University of Western Ontario.

*Abstract*

Most studies find that relative price variability (RPV) is a U-shaped or V-shaped function of anticipated inflation, and a V-shaped function of unanticipated inflation. One exception is Reinsdorf (1994), who finds that RPV in the United States during the 1980s recession was monotonically decreasing in unanticipated inflation. We suggest a reason for this difference, and test our conjecture using data from inter-war Canada. Our results indicate that in recessionary conditions a positive inflation shock does reduce RPV. However, this reduction is unlikely to correspond to higher consumer utility; this has implications for the conduct of monetary policy during a recession.

*JEL classification:* E31, N12

*Key words:* relative price variability, inflation, Canada, Great Depression

## **1. Introduction**

Rates of consumer price inflation are not uniform across commodities or across regions within an economy. Moreover, the magnitude of price or inflation rate dispersion can vary over time. Both theoretical models and econometric evidence indicate that this dispersion, known as ‘relative price variability’ (RPV) is likely to be correlated with the aggregate inflation rate. However, different theories predict different functional forms for the relationship between inflation and RPV, and much of the econometric evidence suggests a non-monotonic relationship that depends on whether the inflation is anticipated or unanticipated.

Overall, there is some consistency in recent empirical results, with most studies finding a U-shaped or V-shaped relationship between RPV and anticipated aggregate inflation, at least during the Great Moderation, and a V-shaped relationship between RPV and unanticipated aggregate inflation. However, a key exception to the results regarding unanticipated inflation is Reinsdorf (1994), whose results imply not only that negative inflation shocks increase RPV, but also that positive inflation shocks reduce RPV: the relationship is not V-shaped but monotonic. These results come from data for the United States during the recession of the early 1980s.

This has potentially serious implications for monetary policy. Inflation-targeting central bank governors are most likely to deviate from their normal reaction function (most likely to generate a policy shock) in the depths of a recession, when their contract gives them permission to use discretion in closing the output gap. The effect of such a positive inflation shock for the whole economy will depend partly on whether the shock increases or reduces RPV. Moreover, whether a fall in RPV is associated with higher aggregate welfare depends on the mechanism generating the price dispersion. Evidence on the direction of the RPV change and on the likely mechanism underlying it can therefore inform forecasts of the welfare effects of an expansionary monetary shock in times of deep recession.

In this paper, we suggest a mechanism that explains why times of recession might be associated with a monotonic relationship between inflation shocks and RPV. This mechanism implies an asymmetry across different types of commodity in the process driving RPV. We then use data from Canada around the time of the Great Depression to test (i) whether the Reinsdorf result appears in a recessionary environment other than the United States in the early 1980s, and (ii) whether the Canadian data support our hypothesis about the asymmetry. Finally, we discuss some possible implications for the welfare effects of expansionary shocks in 21<sup>st</sup> century recessions. We begin with a review and interpretation of the existing literature.

## **2. Literature Review**

### *2.1. Theory*

Several different types of theory can be used to interpret a correlation between RPV and either anticipated or unanticipated aggregate inflation. Firstly, models which incorporate the signal extraction mechanism of Lucas (1973, 1994) predict a non-negative relationship between RPV and the absolute value of unanticipated inflation. This group includes Barro (1976), Hercowitz (1981) and Cukierman (1983). In these models, firms find it difficult to distinguish between aggregate demand shocks and their own idiosyncratic shocks. With a relatively large variance in aggregate demand shocks compared with idiosyncratic ones, firms will tend to interpret idiosyncratic shocks as aggregate ones, and so they will be more likely to adjust prices rather than output in response to a shock. If all firms face demand curves with the same elasticity, then an aggregate shock will not impact on RPV, because all firms will respond to the shock in the same way. However, asymmetries in price elasticity will cause asymmetries in firms' responses, in which case the shock will raise RPV. This is true of either a positive shock or a negative one, suggesting a V-shaped relationship between RPV and inflation with a turning point at zero.

Secondly, search models in the style of Bénabou and Gertner (1993) are consistent with a negative monotonic relationship between RPV and unanticipated inflation. This stems from the equilibrium search strategies of consumers, who can observe prices quoted by two firms in a duopoly only sequentially. If the first price observed exceeds a certain cut-off level, then consumers are willing to pay a fixed cost to find out the other firm's price. This motivates firms to adopt a strategy such that the monopoly price (which can vary between the firms) is charged when the firm's marginal cost falls below a certain level  $c^*$ , but the consumers' no-search threshold price is charged when the marginal cost exceeds  $c^*$ . The threshold level  $c^*$  is decreasing in search costs: when it is easy for consumers to search, firms are less willing to charge the monopoly price. If search costs are sufficiently low, then it will often be the case that a positive aggregate shock to firms' costs will push them above  $c^*$ , and both will charge the no-search threshold price, entailing low RPV. In the absence of such a shock, costs remain below  $c^*$ , and firms revert to monopoly pricing with high RPV.<sup>1</sup>

The range of theoretical predictions regarding the effect of anticipated inflation is even wider. Head and Kumar (2005) present a search model that predicts a U-shaped relationship

---

<sup>1</sup> Dana (1994) develops an oligopolistic pricing model with search which is similar in spirit. Here, nominal marginal costs are stochastic but perfectly correlated across firms. However, firms adopt a mixed strategy over a continuous price range (provided search costs are not too high), and consumers are unable to infer actual costs over this range. As in the Bénabou and Gertner model, consumers adopt a reservation price search strategy, and this price forms the upper bound of prices charged by firms (such that no search actually takes place in equilibrium.) This results in a negative average relationship between cost shocks and RPV. The reservation price implies that even when all firms' costs are higher than average, the upper end of the price range does not change. But because the minimum price they can charge does rise, the range of possible prices shrinks, resulting in less price dispersion on average when costs are relatively high than when they are relatively low.

between RPV and anticipated inflation.<sup>2</sup> In this model, search intensity is decided by households before prices are observed. Search yields a distribution of price offers, the number of prices depending on search intensity, and buyers make a purchase only if the lowest price observed is below a reservation price. They spend their entire budget on the lowest-priced good if this price is less than or equal to the ratio of the marginal utility of consumption to the marginal value of carrying money into the next period, and purchase none of the good otherwise. In the model, inflation raises firms' market power by lowering the return to holding money, resulting in a larger range of posted prices. At low inflation levels, the range of prices posted is small, and there is very little search. In this case, a small rise in aggregate inflation will raise prices at the top end of the distribution, leading to greater search intensity and a *lower* overall cross-sectional variance in prices. However, as aggregate inflation continues to rise further increases in search intensity are small, and the dominant effect is the increase in firms' market power, which increases price dispersion.

A non-monotonic relationship between RPV and anticipated inflation is also predicted by the menu cost model of Rotemberg (1982, 1983). In this model, each firm faces a demand curve that is decreasing in the firm's price relative to the market average and increasing in the real value of money (that is, the nominal money stock relative to the average price). With a quadratic cost curve and a log-linear demand curve, the elasticity of the optimal relative price with respect to the real value of money is positive and constant. If the money stock grows at a constant rate ( $\mu$ ), this will also be the growth rate of each firm's optimal price level. However, changes in this price entail a fixed cost, so each firm changes its price at discrete intervals. It can be shown that the optimal interval length ( $\lambda$ ) is equal to  $\alpha / \mu^{2/3}$ , where  $\alpha$  is a constant. Nevertheless, with a

---

<sup>2</sup> Head *et al.* (2005) extend this model to incorporate shocks to productivity and the monetary growth rate. The results with regard to the relationship between RPV and aggregate inflation are very similar.

uniform distribution of firms within this interval, the aggregate inflation rate will still be  $\mu$ . Danziger (1987) shows that in such a model, the observed variance of firms' prices will be approximately equal to  $\mu^2 \cdot t \cdot [\lambda - t]$ , where  $t$  is the interval between observations, as long as  $t \leq \lambda$ . Substituting the expression for  $\lambda$  produces the function  $\mu^2 \cdot t \cdot [\alpha / \mu^{2/3} - t]$ , which has an inverse-U shape: the variance of prices is equal to zero at  $\mu = 0$  and  $\mu = (\alpha / \lambda)^{3/2}$ , but positive in between. For  $t > \lambda$ , it can be shown that there is an inverse-U function in each interval  $[(n \cdot \alpha / \lambda)^{3/2}, ((1 + n) \cdot \alpha / \lambda)^{3/2}]$ , where  $n$  is an integer, with the variance falling to zero at the edges of the interval (see Figure 1). Depending on the parameter values and the observed inflation range, the observed relationship between RPV and inflation might be monotonic, inverse-U-shaped or V-shaped; it could also have more than one turning point.

*Figure 1 about here*

In a similar vein, Choi (2010) shows that a multi-sector, Calvo-type sticky price model can also account for a U-shaped relationship between anticipated inflation and RPV. In the model, sectors are distinguished by the degree of price stickiness, which is summarized by the exogenous probability that a firm will be able to change its price in any given period. Numerical simulations produce a U-shaped relationship with a turning point slightly greater than zero.

## *2.2. Evidence*

The empirical literature on cross-sectional RPV and inflation begins with Parks (1978), who analyzes the determinants of the variability in inflation across different components of the US consumer price index. Parks finds a significant positive correlation with the aggregate inflation rate, and this result is replicated in studies which use a similar methodology, for example Lach and Tsiddon (1992). Other papers, including Parsley (1996) and Debelle and Lamont (1997),

analyze the determinants of the variability of inflation across different US cities, again finding a significant positive correlation with aggregate inflation. Later papers explore a number of themes, including the relationship between the variability of price *levels* and aggregate inflation, the non-linearity of the relationship, its stability over time, and the difference between the effects of anticipated and unanticipated inflation. Data are taken not only from the US, but also from Europe and Latin America.

Evidence for the non-linear effect of aggregate inflation on RPV appears in studies such as Hartmann (1991), Dabus (2000), and Çağlayan and Filiztekin (2003). Other papers test for a non-linear effect of anticipated inflation by combining an RPV model with an inflation forecasting model. Reinsdorf (1994), Aarstol (1999), and Becker and Nautz (2009) find a positive relationship between RPV and the absolute value of anticipated inflation (or its square), suggesting a V-shaped relationship with a turning point at zero. However, studies fitting a model with a more flexible functional form, such as Choi and Kim (2010) and Becker and Nautz (2010), find a turning point at a positive inflation rate. Studies fitting an RPV model incorporating a non-parametric function of inflation, such as Fielding and Mizen (2008) and Choi (2010), find that the relationship is approximately U-shaped with a positive turning point. The positive turning point is consistent with several theoretical models, including a Calvo-type sticky price model, a menu cost model, and some search models. Overall, there is some degree of consistency in the empirical results regarding RPV and total or anticipated inflation, with most studies pointing to a U-shaped or V-shaped relationship, possibly with a turning point at a positive inflation rate. However, those studies which test for instability in the RPV-inflation relationship, such as Becker and Nautz (2009) or Choi (2010), typically find significant changes in the parameter values over time. For example, Choi provides evidence that the relationship was U-shaped during



the Great Moderation in the US, but positive at other times, which is consistent with a Calvo-type sticky price model.

Most econometric models that include terms in anticipated inflation also include a term in the absolute value of unanticipated inflation, sometimes with a coefficient that varies according to whether there is a positive or negative inflation shock. When the anticipated inflation series is constructed using an ARCH model of inflation, the RPV model also includes a term in the conditional variance of the inflation forecast. Almost all of these papers, including Aarstol (1999), Fielding and Mizen (2008), Çağlayan *et al.* (2008) and Becker and Nautz (2009, 2010), find positive and significant coefficients, although the coefficient for negative inflation shocks is sometimes insignificantly different from zero. In other words, there is general support for a V-shaped function. However, a substantial exception is Reinsdorf (1994), who finds that higher unanticipated inflation in the price of standard grocery items reduces RPV across US cities. Both V-shaped and downward-sloping functions are consistent with some of the relevant theories. For example, a V-shaped relationship is consistent with Lucas-type model and a negative monotonic one is consistent with a Bénabou and Gertner-type model. But what type of model is likely to be applicable to the Reinsdorf data and not to the data used in other papers?

### *2.3. An interpretation of the conflicting evidence*

One distinctive characteristic of the Reinsdorf data is that they represent a strongly recessionary period in the US (1980-1982). Annual real US GDP growth in 1982 was  $-2\%$ , the lowest value between the Great Depression and the Global Financial Crisis. The annual real rate of growth of credit to the private sector in 1980 was  $-11\%$ , the lowest value in the post-war record. In this year, the bank prime loan rate reached  $20\%$ , and the gap between this rate and the treasury bill

rate reached 8%, the highest values ever recorded.<sup>3</sup> Many US households are likely to have been credit-constrained in the early 1980s. It is not clear why the existing search models that predict a negative monotonic relationship between RPV and unanticipated inflation should be especially relevant to this recessionary period. However, there is another theoretical explanation for a link between the negative correlation and the recession.

Shocks to prices require variation in nominal consumption, if households are to maintain a smooth pattern of real consumption. In a recession characterized by credit constraints, this requires variation in a buffer stock of liquid assets. However, there is some disagreement about whether households do hold such stocks, and whether variations in money demand can be represented by a buffer stock model (Carr and Darby, 1981; Judd and Scadding, 1982; Cuthbertson, 1997; Mizen, 1997). The model might not be applicable, if the opportunity cost of holding money is high even in low-inflation environments. For example, this might be the case if the chosen level of saving is the result of an intra-household bargaining process between household members with different inter-temporal discount rates. In such an environment, more liquid assets are at a greater risk of expropriation by the member with the highest discount rate. There is evidence for such an effect in some low-income countries where credit markets are permanently lacking; see for example Andersen and Baland (2002).<sup>4</sup> In this case, there might be no buffer stock of money, and households might have no way of immediately adjusting their total nominal consumption after observing a price shock.

---

<sup>3</sup> These figures are based on chained real GDP, consumer credit, CPI and interest rate data in the Federal Reserve Bank of St Louis FRED database.

<sup>4</sup> In these countries, inter-temporal discount rates appear to be correlated with gender: the wife wants to save money and the husband wants to spend it all.

This creates a feature which is not present in any of the existing theoretical models of RPV. If nominal money balances and nominal consumption are set before price shocks are observed, then a positive (negative) shock will reduce (increase) real consumption and increase (reduce) the marginal utility of consumption relative to the opportunity cost of search. If the change in marginal utility is large enough, there will be an increase (reduction) in the amount of search in equilibrium, leading to lower (higher) RPV. In Appendix 1, we present a simple theoretical model which formalizes this intuition. Such an explanation for a negative correlation between RPV and unanticipated inflation entails a further prediction: the correlation will be stronger for non-storable goods, and weaker or non-existent for storable goods. Even if households choose not to hold a buffer stock of liquid assets, they can still hold a buffer stock of storable goods (which are illiquid assets). If the storage costs are negligible, then for storables the negative correlation should disappear.

In the following sections, we present an econometric model of monthly RPV in inter-war Canada. Our sample period incorporates years before, during and after the Great Depression. We will test for parameter stability over the sample period, but there is some reason to suppose that Canadian households faced credit constraints for much of the period, not just during the Great Depression. Access to bank loans during this period was limited, and Canadian households relied much more heavily on the informal sector than did households in the US (Harris and Ragonetti, 1998). A lack of bank finance led to the rise of credit unions in 1930s Canada, but these organisations often channeled all of their resources into loans for small businesses, providing no consumer credit (Levasseur and Rousseau, 2001). If our hypothesis regarding the negative correlation between RPV and unanticipated inflation in the 1980s is correct, then we should also see such a correlation in the Canadian inter-war data. Moreover, the correlation should be stronger for non-storable goods than for storable ones.

### 3. The Canadian Data

Following Hajzler and MacGee (2011), our data are taken from monthly issues of the *Canada Labour Gazette*, which are available for the period November 1922 – November 1940. This publication lists the monthly prices of a variety of grocery items in a number of Canadian cities. The data are therefore similar in character to those used by Reinsdorf (1994), except for differences in the range of items covered, as one might expect in data collected fifty years previously. Not all prices are available for all cities, but the prices of 42 items are reported for 69 cities over the whole period with just a few missing observations; it is these prices that form our data set.<sup>5</sup> The cities are listed in Table 1 and the grocery items in Table 2. Table 2 separates the items into storable and non-storable goods, a distinction that will be important in our data analysis. Our rule of thumb in classifying items is that goods which are likely to be unusable a month after purchase (bearing in mind that most Canadian households during this period lacked a refrigerator) are non-storable.<sup>6</sup> A few items, such as lard, might be allocated to either group, but the results presented below are not sensitive to our classification of these items.

For each of the 42 items, we construct a Canada-wide inflation series and a geographical RPV series as follows. Let  $p_{ijt}$  be the price of item  $i$  in city  $j$  in month  $t$ , and  $p_{it}$  the mean price across all cities.<sup>7</sup> Then the item-specific aggregate inflation series is:

---

<sup>5</sup> Newfoundland did not become part of Canada until 1949, so there are no Newfoundland cities in the data set.

<sup>6</sup> It is possible that storability increases during the winter months, and we did explore whether the parameters of our model were subject to seasonal variation. However, no clear pattern emerged from this exercise. Generally, it seems that non-storable good prices do not take on the characteristics of storable good prices during winter.

<sup>7</sup> Note that these are unweighted averages. We do not have data on the value of consumption in each city.

$$\pi_{it} = \ln(p_{it}) - \ln(p_{it-1}) \quad (1)$$

and the corresponding RPV series is:

$$v_{it} = \ln \left( \frac{1}{69} \sqrt{\sum_j \left( \frac{p_{ijt}}{p_{it}} - 1 \right)^2} \right) \quad (2)$$

That is, RPV is measured as the logarithm of the standard deviation of relative prices. The distributions of the two series are illustrated in Figures 2-3. The distribution of aggregate inflation is centred on zero; it also has very fat tails, so Figure 3 shows the distribution trimmed at  $\pm 10\%$ . The maximum and minimum values are around  $\pm 40\%$ , and section 5 includes a discussion of how we explore the sensitivity of our results to these outliers.

Some typical inflation and RPV series are illustrated in Figures 4-5. These figures show some of the heterogeneity in the data, and can be compared with the annual CPI inflation and real GNP growth series illustrated in Figure 6. Figure 4 includes annual average aggregate inflation rate series for four items: veal, lard, oats and prunes. For all four items there is substantial price deflation during the Great Depression era (1929-1933), corresponding to the aggregate CPI deflation and negative real GNP growth. However, the length of the deflationary period varies across items, as does the magnitude of the deflation relative to the variation in other parts of the sample period. For example, the deflation in veal prices lasts for the whole of the Great Depression and is the dominant feature in the time series, but the deflation in oat prices is more short-lived and not such a dominant feature. Inflation rates are not highly correlated across items, a characteristic that is also apparent in the RPV series in Figure 5. Further, comparison of Figures 4 and 5 does not reveal any obvious common pattern in the relationship between RPV and inflation.

*Tables 1-2 and Figures about 2-6 here*

The marked heterogeneity in the RPV and inflation in Figures 4-5 suggests that the parameters of the RPV-inflation relationship are likely to vary across items. If there is any persistence in RPV, this may also vary across items. The econometric model discussed in the next section is designed to take account of this heterogeneity.

#### 4. The RPV Model

Our model is designed to identify the effects of anticipated and unanticipated inflation on inter-war Canadian RPV, as measured by the variable  $v_{it}$  in equation (2). First of all, we deseasonalize the RPV and inflation time series; the deseasonalized variables  $v_{it}^D$  and  $\pi_{it}^D$  are residuals from regressions of  $v_{it}$  and  $\pi_{it}$  on monthly dummy variables. Next,  $\pi_{it}^D$  is decomposed into an anticipated component ( $\pi_{it}^A$ ) and an unanticipated component ( $\pi_{it}^U$ ), as in papers such as Fielding and Mizen (2008) and Becker and Nautz (2009). The decomposition is based on an ARCH model of aggregate inflation:

$$\pi_{it}^D = \pi_{it}^A + \pi_{it}^U \quad (3)$$

$$\pi_{it}^A = \gamma_{0i} + \gamma_{1i} \cdot \pi_{it-1}^D + \gamma_{2i} \cdot \pi_{it-2}^D + \gamma_{3i} \cdot t \quad (4)$$

$$\pi_{it}^U \sim N(0, h_{it}^2) \quad (5)$$

$$h_{it}^2 = \delta_{0i} + \delta_{1i} \cdot (\pi_{it-1}^U)^2 \quad (6)$$

Here,  $h_{it}^2$  is the conditional variance of the inflation forecast, capturing inflation uncertainty. Note that the  $\gamma$  and  $\delta$  parameters are specific to each item  $i$ ; in other words, the dynamics of inflation are allowed to vary from one item to another. The fitted coefficient values are not reported here, but are available on request; they vary significantly across the items, as one would expect given the heterogeneity apparent in Figure 4.

Then we fit a number of alternative RPV regression equations, each having the following general form:

$$v_{it}^D = \alpha_{0i} + \alpha_{1i} \cdot v_{it-1}^D + \alpha_{2i} \cdot \pi_{it}^{U+} + \alpha_{3i} \cdot \pi_{it}^{U-} + \alpha_{4i} \cdot t + \beta(\pi_{it}^A) + \theta(h_{it}) + u_{it} \quad (7)$$

There is one regression equation for each item. The residuals  $u_{it}$  may be correlated across the items, so the equations are fitted using a SUR estimator. In equation (7), unanticipated inflation is decomposed into its positive values:

$$\pi_{it}^{U+} = 0.5 \left( \pi_{it}^U + \left| \pi_{it}^U \right| \right) \quad (8)$$

and its negative ones:

$$\pi_{it}^{U-} = 0.5 \left( \pi_{it}^U - \left| \pi_{it}^U \right| \right) \quad (9)$$

This allows for the possibility *either* that the relationship between RPV and unanticipated inflation takes its more usual V-shaped form, *or* that there is a monotonic relationship, as in Reinsdorf (1994).<sup>8</sup>

Different theories suggest a wide range of functional forms for the relationship between RPV and anticipated inflation, and this range is reflected in the variety of functional forms in existing empirical studies. In other papers, the relationship is U-shaped or V-shaped, and the

---

<sup>8</sup> Equation (7) imposes the restriction that if RPV is a non-monotonic function of unanticipated inflation, then the turning point is at  $\pi^U = 0$ . In none of the relevant theoretical models is there a non-zero turning point. In this respect, the range of theoretical predictions regarding the effect of unanticipated inflation is narrower than the range of predictions regarding the effect of anticipated inflation. Theory does not require the effect of unanticipated inflation to be linear, but if terms in  $(\pi^{U+})^2$  and  $(\pi^{U-})^2$  are added to equation (7), the resulting parameter estimates are insignificantly different from zero.

turning point is not necessarily at  $\pi^A = 0$ . For this reason, we fit alternative versions of equation (7) with different parameterizations of the  $\beta$  function. One of these is a quadratic function:

$$\beta = \alpha_{5i} \cdot \pi_{it}^A + \alpha_{6i} \cdot (\pi_{it}^A)^2 \quad (10)$$

In Appendix 2, we show that a non-parametric estimate of the  $\beta$  function produces a curve that is approximately quadratic, lending some support to equation (10), and the results presented below will focus mainly on estimates using this quadratic function. However, we will also compare the quadratic parameterization with a piecewise-linear parameterization that has been used in some other papers:

$$\beta = 0.5 \alpha_{5i} \cdot (\pi_{it}^A + |\pi_{it}^A|) + 0.5 \alpha_{6i} \cdot (\pi_{it}^A - |\pi_{it}^A|) = \alpha_{5i} \cdot \pi_{it}^{A+} + \alpha_{6i} \cdot \pi_{it}^{A-} \quad (11)$$

This function allows for a V-shaped curve. Finally, we allow for the possibility that RPV depends either on the standard deviation of the inflation forecast, so  $\theta = \alpha_{7i} \cdot h_{it}$ , or on the variance, so  $\theta = \alpha_{7i} \cdot h_{it}^2$ .

Equation (7) allows the effect of inflation on RPV to vary across the different grocery items; it also allows for heterogeneity in RPV dynamics, as captured by the parameter  $\alpha_{1i}$ . Having allowed for such heterogeneity in our model, we can then calculate consistent estimates of the mean parameter values  $\frac{1}{42} \sum_i \alpha_{ki}$ ,  $k = 1, \dots, 7$ , and the corresponding t-ratios can be calculated using the Delta Method. Means and corresponding t-ratios can also be calculated for the storable and non-storable sub-groups separately.



## 5. Results

Table 3 reports estimates of the mean values of the parameters in the four different parametric models as follows:

<i>Panel</i>	<i><math>\beta</math> function</i>	<i><math>\theta</math> function</i>
A:	$\beta = \alpha_{5i} \cdot \pi_{it}^A + \alpha_{6i} \cdot (\pi_{it}^A)^2$	$\theta = \alpha_{7i} \cdot h_{it}^2$
B:	$\beta = \alpha_{5i} \cdot \pi_{it}^{A+} + \alpha_{6i} \cdot \pi_{it}^{A-}$	$\theta = \alpha_{7i} \cdot h_{it}^2$
C:	$\beta = \alpha_{5i} \cdot \pi_{it}^A + \alpha_{6i} \cdot (\pi_{it}^A)^2$	$\theta = \alpha_{7i} \cdot h_{it}$
D:	$\beta = \alpha_{5i} \cdot \pi_{it}^{A+} + \alpha_{6i} \cdot \pi_{it}^{A-}$	$\theta = \alpha_{7i} \cdot h_{it}$

After taking lags in the inflation forecasting model, the sample period is February 1923 – November 1940. For each version of the model and for each parameter estimate, three values are reported: the mean across all items, the mean across non-storables and the mean across storables; the table also includes the corresponding t-ratios. These estimates are based on untrimmed inflation figures, and there is a concern that these estimates might be driven by extreme values of inflation. For this reason, Table 4 includes a parallel set of parameter estimates in which  $\pi_{it}$  in equation (1) is replaced by an inflation series trimmed at  $\pm 10\%$ . The figures appearing in Table 4 are very similar to those appearing in Table 3, so our results appear not to be sensitive to the treatment of outliers. This conclusion is confirmed in Appendix 2, which reports similar results using a semi-parametric estimator that is robust to the existence of outliers in the distribution of the explanatory variable.

### 5.1. The mean parameter estimates for all items

In Table 3 panels A and C, the first columns (for all items) include estimates of the mean  $(\pi^A)^2$  parameter that are significantly greater than zero. However, the  $\pi^A$  parameter is statistically insignificant. In other words, RPV is increasing in anticipated inflation, and there is significant non-linearity in the relationship, but in this model the value of anticipated inflation at which RPV is minimized is insignificantly different from zero. Figure 7 illustrates the relationship by plotting the size of  $dv^D / d\pi^A$  implicit in the mean parameter values in panel A for different levels of  $\pi^A$ . The value of this derivative is increasing in  $\pi^A$ . For a value as high as 2.5% per month,  $dv^D / d\pi^A \approx 10$ ; that is, a further increase in  $\pi^A$  by 0.1 percentage points is estimated to lead to a 1% increase in RPV. (The plot corresponding to the panel C parameter estimates is very similar, and is not shown in the figure.) The curve in Figure 7 is similar in shape to equivalent curves in several recent papers using modern North American data, such as Fielding and Mizen (2008) and Choi (2010).

Panels B and D of Table 3 include estimates of the alternative parameters in the piecewise-linear model. These are broadly consistent with those in the quadratic model, insofar as  $dv^D / d\pi^{A+} > 0 > dv^D / d\pi^{A-}$ . However,  $dv^D / d\pi^{A+}$  is insignificantly different from zero. If in fact  $dv^D / d\pi^{A+}$  is increasing in  $\pi^{A+}$ , as suggested by Figure 7, then the imprecision of the estimate of the mean  $\pi^{A+}$  coefficient is not too surprising.

Estimates of the mean effects of unanticipated inflation are very different from the  $\pi^A$  effects. In all four panels in Table 3, both the mean  $\pi^{U+}$  parameter for all items and the mean  $\pi^{U-}$  parameter for all items are significantly less than zero. The absolute value of the  $\pi^{U+}$  parameter is slightly less than one; that of the  $\pi^{U-}$  parameter is slightly greater than 0.5. A positive one percentage point inflation shock reduces RPV by about 1%; a negative one

percentage point inflation shock increases RPV by about 0.5%. These two effects are significantly different in magnitude: the extent to which positive inflation shocks reduce RPV, on average, is significantly greater than the extent to which negative inflation shocks increase RPV. Nevertheless, the effect of unanticipated inflation on RPV is uniformly negative: there is no V-shaped relationship. This is consistent with the findings of Reinsdorf (1994), as anticipated in section 2, and with some theoretical models, such as the search model with liquidity constraints in Appendix 1.

Finally, greater uncertainty in the inflation forecast, measured either by the conditional variance of the forecast (panels A and B of Table 3) or by the conditional standard deviation (panels C and D), is associated with significantly *less* RPV. This effect is at odds with some theoretical models, for example, the signal extraction model. However, it is not necessarily inconsistent with a search model. Greater *ex ante* inflation uncertainty might make the return to search effort more uncertain, but it can also make the return to substitutes for search effort (for example, holding larger nominal money balances) more uncertain, so search effort does not necessarily fall in times of greater uncertainty.

*Tables 3-4 and Figure 7 about here*

### *5.2. The mean parameter estimates for storable and non-storable*

Table 3 indicates that there are significant differences between storable and non-storable with respect to the effects of both anticipated inflation and unanticipated inflation. For anticipated inflation, the mean estimates of  $(\pi^A)^2$  parameter in panels A and C are significantly larger for storable, indicating significantly more convexity in the function, on average. Moreover, mean estimates of the  $\pi^A$  parameter for non-storable are significantly less than zero, indicating a

turning point of the function at a significantly positive inflation rate; this is not the case for storables. These differences are illustrated in the plots of  $dv^D / d\pi^A$  in Figures 8-9. There is no straightforward theoretical reason why this difference should arise.

More relevant to our hypothesis are the differences in the effect of unanticipated inflation. In all four models, the mean non-storables  $\pi^{U+}$  parameter is about  $-0.75$ , significantly less than zero and significantly less than the mean storables  $\pi^{U+}$  parameter; the storables parameter is insignificantly different from zero.<sup>9</sup> By contrast, estimates of the mean  $\pi^{U-}$  parameter vary very little between storables and non-storables: in all cases the estimated parameter value is about  $-0.9$ . Qualitatively, the results for non-storables here resemble the results for all items, and the Reinsdorf results. Moreover, the insignificant storables  $\pi^{U+}$  parameter is consistent with the conjecture that the explanation for the significantly negative non-storables  $\pi^{U+}$  parameter lies in a search model with liquidity constraints: households do not hold large buffer stocks of liquid assets, but they can nevertheless hold some buffer stocks of storable goods. This does not necessarily mean that the relationship between storables RPV and unanticipated inflation is entirely unaffected by liquidity constraints. One possibility is that for storables the effect is relatively weak and offset by some other mechanism that creates a positive relationship in the absolute value of inflation shocks (as for example in a signal extraction model). This would explain why the storables  $\pi^{U-}$  parameter is significantly negative. Nevertheless, there is a large

---

<sup>9</sup> There are negative estimates of the  $\pi^{U+}$  coefficient for all 25 non-storable items except the following: butter solids, eggs (cooking), onions, potatoes, and round steak. None of the positive coefficient estimates for these items is significantly different from zero; the largest, for round steak, is  $0.13$  ( $t = 0.40$ ). Among the 17 storable items, there are three individual  $\pi^{U+}$  coefficient estimates that are significantly greater than zero at the 5% level (canned peaches, oats, and raisins) and four that are significantly less than zero (canned salmon, both types of sugar, and coffee).

and significant difference between the effect of inflation shocks on storables RPV and their effect on non-storables RPV.

*Figures 8-9 about here*

### 5.3 Parameter stability

As Figure 6 shows, our sample period covers not only the Great Depression in Canada, but also the relatively prosperous years before and afterwards. It is possible that the variation in macroeconomic conditions affects the relationship between RPV and inflation, so that some of the parameters in equation (7) are not stable over time. For this reason, we also fit the model to eight-year sub-samples, the first ending in December 1930, the second in December 1931, and so on to the last sub-sample, ending in November 1940. Each of sub-sample has 96 observations except the first one (missing January 1923) and the last one (missing December 1940), which have 95 observations. These first and last subsamples both exclude the trough of the depression (1931-1932); other subsamples include the trough. If the Great Depression affects the relationship between RPV and inflation, this should be apparent in differences in parameter estimates across subsamples.

The charts in Figures 10-11 illustrate the  $\pi^A$  and  $\pi^U$  parameter estimates in model A. The stylized facts discussed here also apply to the parameter estimates in the other models, which are not shown. Figure 10 includes the mean  $\pi^A$  and  $(\pi^A)^2$  parameter estimates for all items, storables and non-storables; Figure 11 includes the mean  $\pi^{U+}$  and  $\pi^{U-}$  parameter estimates for these groups. The estimates are indicated by the black lines, with the 95% confidence interval in gray. In each chart, the horizontal axis indicates the last year in the sub-sample corresponding to the parameter estimate measured on the vertical axis. Overall, there does seem to be some change in the relationship between RPV and anticipated inflation, as shown in Figure 10. The all-items

$(\pi^A)^2$  parameter is significantly greater than zero in subsamples ending in 1935 or earlier, but its value falls over time, and is insignificantly different from zero in later subsamples. By the final subsample, neither the mean  $\pi^A$  parameter nor the mean  $(\pi^A)^2$  parameter is significantly different from zero. The same is true for storables and non-storables. Figure 11 shows that the relationship between RPV and anticipated inflation is somewhat more stable. All of the  $\pi^{U-}$  parameter estimates are significantly below zero in all subsamples, as are the all-items and non-storables  $\pi^{U+}$  parameter estimates. The storables  $\pi^{U+}$  parameter estimate is insignificantly different from zero in all subsamples.

It appears that there is some change in the relationship between RPV and anticipated inflation over time. The relationship is much stronger before the Great Depression than afterwards. There are a number of possible explanations for this change, given the variety of theoretical models that could account for the relationship. Whatever mechanism explains the U-shaped curve of the 1920s – be it a menu cost model or a search model – seems to have been part of a market structure that underwent a permanent transformation during the depression, so that there is no significant relationship in the late 1930s. By contrast, the conditions giving rise to the negative effect of inflation shocks on RPV apply to the whole of our sample period. This result contrasts with existing evidence from modern data, in which the monotonic relationship between inflation shocks and RPV appears only in the depths of a recession.

*Figures 10-11 about here*

## **6. Summary and Conclusion**

Fitting a model of RPV to data from inter-war Canada, we find that the effect of anticipated changes in aggregate inflation on price dispersion is similar to effects found in modern data: there is an approximately U-shaped relationship, with dispersion minimized at an aggregate inflation

rate close to (but possibly slightly greater than) zero. However, this relationship appears to weaken in the years immediately before World War II. Overall, we also find that the effect of unanticipated changes in aggregate inflation in inter-war Canada is similar to the effect found by Reinsdorf (1994), who uses data from the recession of the early 1980s. This effect is different from effects found in other modern data. In inter-war Canada and in the recession, there is a negative monotonic relationship between RPV and inflation shocks; in other modern data, there is a V-shaped relationship. However, in inter-war Canada, the negative effect on RPV of positive inflation shocks is much stronger for non-storable items than for storable ones.

One possible interpretation of these findings is that the response of RPV to inflation shocks during a recession – or in economies at an earlier stage of financial development – is explained by a model in which households face severe liquidity constraints. A positive shock to prices threatens to reduce their real consumption, and they respond by devoting more time and effort to searching for lower prices. The increased search effort reduces RPV. For storable items the liquidity constraint has less effect, because households can run down their stocks of these items when they become unexpectedly expensive.

If this is the correct explanation for the patterns observed in the data, then there are implications for the conduct of monetary policy during a recession.<sup>10</sup> An expansionary monetary policy shock will reduce RPV: it will generate more inflation in those locations where prices and costs are relatively low – in other words, where the recession is most severe. There will be more of a price stimulus for firms in the most economically depressed locations. This suggests that a monetary policy intervention will be automatically well targeted, which adds weight to arguments favoring monetary policy over other interventions in such circumstances. However,

---

<sup>10</sup> Since Canada had free banking for most of our sample period, it is not possible for us to test hypotheses about monetary policy directly using our data.

the monetary policy shock also imposes costs on households, reducing their real consumption below the level planned. This effect is likely to be larger for non-storable items (perishable food; rent and services) than for storable ones (preserved food; consumer durables). *If* the share of non-storable items in the consumption of the poorest households is relatively large, then these households will bear a large part of the costs of monetary policy shock.

A single data set such as ours can provide only circumstantial evidence for a given theory. However, the onset of the Global Financial Crisis means that there will soon be data available from a wide range of liquidity-constrained economies that can be used to test the hypotheses explored in this paper. The replication of our results using contemporary data would indicate that more attention should be paid to the effect of liquidity constraints on the cross-sectional distribution of prices.

## **References**

- Aarstol, M. (1999) 'Inflation, inflation uncertainty and relative price variability,' *Southern Economic Journal* 66: 414-23.
- Andersen, S. and Baland, J-M. (2002) The economics of ROSCAs and intrahousehold resource allocation,' *Quarterly Journal of Economics* 117(3): 963-995.
- Barro, R.J. (1976) 'Rational expectations and the role of monetary policy,' *Journal of Monetary Economics* 2: 1-32.
- Becker, S.S. and Nautz, D. (2009) 'Inflation and relative price variability: new evidence for the United States,' *Southern Economic Journal* 76:146-64.
- Becker, S.S. and Nautz, D. (2010) 'Inflation, price dispersion and market integration through the lens of a monetary search model,' SFB 649 Discussion Paper 2010-058, Freie Universität Berlin, Germany.



- Bénabou, R. and Gertner, R. (1993) 'Search with learning from prices: does increased inflationary uncertainty lead to higher markups?' *Review of Economic Studies* 60: 69-94.
- Çağlayan, M. and Filiztekin, A. (2003) 'Nonlinear impact of inflation on relative price variability,' *Economics Letters* 79: 213-218.
- Çağlayan, M., Filiztekin, A. and Rauh, M.T. (2008) 'Inflation, price dispersion, and market structure,' *European Economic Review* 52: 1187-1208.
- Carr, J. and Darby, M.R. (1981) 'The role of money supply shocks in the short-run demand for money,' *Journal of Monetary Economics* 8(2): 183-199.
- Choi, C-Y. (2010) 'Reconsidering the relationship between inflation and relative price variability,' *Journal of Money, Credit and Banking* 42(5): 769-798.
- Choi, C-Y. and Kim, Y.S. (2010) 'Is there any asymmetry in the effect of inflation on relative price variability?' *Economics Letters* 108: 233-236.
- Cukierman, A. (1983) 'Relative price variability and inflation: a survey and further results,' in K. Brunner and A.H. Meltzer (eds.) *Variability in Employment, Prices and Money*, Amsterdam: Elsevier.
- Cuthbertson, K. (1997) 'Microfoundations of the demand for money,' *Economic Journal* 107: 1186-1201.
- Dabus, C. (2000) 'Inflationary regimes and relative price variability: evidence from Argentina,' *Journal of Development Economics* 62: 535-47.
- Dana, J.D. (1994) 'Learning in an equilibrium search model,' *International Economic Review* 35: 745-771.
- Danziger, L. (1987) 'Inflation, fixed cost of price adjustment and the measurement of relative price variability,' *American Economic Review* 77: 704-13.

- Debelle, G. and Lamont, O. (1997) 'Relative price variability and inflation: evidence from U.S. cities,' *Journal of Political Economy* 105: 132-52.
- Fielding, D. and Mizen, P. (2008) 'Evidence on the functional relationship between relative price variability and inflation with implications for monetary policy,' *Economica* 75: 683-99.
- Hajzler, C. and MacGee, J. (2011) 'Retail price differences across U.S. and Canadian cities during the interwar period,' paper presented to the Australian Conference of Economists, Australian National University, Canberra, 11-13 July.
- Harris, R. and Ragonetti, D. (1998) 'Where credit is due: residential mortgage finance in Canada, 1901 to 1954,' *Journal of Real Estate Finance and Economics* 16(2): 223-238.
- Hartman, R. (1991) 'Relative price variability and inflation,' *Journal of Money, Credit and Banking* 23: 185-205.
- Head, A. and Kumar, A. (2005) 'Price dispersion, inflation, and welfare,' *International Economic Review* 46: 533-572.
- Head, A., Kumar, A. and Lapham, B. (2005) 'Innovation, search, and incomplete price adjustment,' mimeo, Department of Economics, Queen's University, Kingston, Ontario, Canada.
- Hercowitz, Z. (1981) 'Money and Dispersion of Relative Prices,' *Journal of Political Economy* 89: 328-56.
- Judd, J.P. and Scadding, J.L. (1982) 'Dynamic adjustment in the demand for money: tests of alternative hypotheses,' *Economic Review* (Fall): 19-30.
- Lach, S. and Tsiddon, D. (1992) 'The behavior of prices and inflation: an empirical analysis of disaggregated price data,' *Journal of Political Economy* 100: 349-89.

- Levasseur, R. and Rousseau, Y. (2001) 'Social movements and development in Quebec: the experience of the Desjardins Movement,' *Annals of Public and Cooperative Economics* 72(4): 549-579.
- Lucas, R.E. (1973) 'Some international evidence on output-inflation trade-offs,' *American Economic Review* 63: 326-34.
- Lucas, R. E. (1994) 'On the welfare cost of inflation,' CEPR Working Paper 394.
- Mizen, P. (1997) 'Microfoundations for a stable demand for money function,' *Economic Journal* 107: 1202-1212.
- Parks, R.W. (1978) 'Inflation and relative price variability,' *Journal of Political Economy* 86: 79-95.
- Parsley, D. (1996) 'Inflation and relative price variability in the short and long run: new evidence from the United States,' *Journal of Money, Credit, and Banking* 28: 323-41.
- Reinsdorf, M. (1994) 'New evidence on the relationship between inflation and price dispersion,' *American Economic Review* 84: 720-31.
- Rotemberg, J.J. (1982) 'Monopolistic price adjustment and aggregate output,' *Review of Economic Studies* 49: 517-31.
- Rotemberg, J.J. (1983) 'Aggregate consequences of fixed costs of price adjustment,' *American Economic Review* 50: 433-6.

## Appendix 1: A Theoretical Model of RPV with Liquidity Constraints.

This model captures the intuition that if people choose nominal money balances before they observe prices, and if goods are non-storable, then a positive inflation shock will reduce consumption and increase the marginal utility from consumption. If this effect is strong enough, and if there is any difference in firms' costs (and therefore prices) across locations, then the marginal opportunity cost of searching to find better prices will be lower after an inflationary shock. The increased search will result in higher demand for the low-cost firms. If firms are perfectly competitive but face increasing marginal costs of production, this will cause the marginal cost of the low-cost firms to rise relative to the marginal cost of the high-cost ones, so there will be lower price dispersion.<sup>11</sup>

The model is as follows, with the parameterization of cost and utility functions chosen so as to facilitate a tractable solution. There is large number of locations, each inhabited by households with mass equal to 1 and a firm which behaves perfectly competitively. (Even though there is a single firm, it does not exploit any local monopoly power, because its local market is contestable.) In any given month, the firm in each location can be one of two types, defined by the marginal cost functions  $mc_1(x)$  and  $mc_2(x)$ , where  $x$  is demand. We assume that  $mc_1(x) > mc_2(x)$  and that firms in each location set price equal to marginal cost, which is increasing in  $x$ . Denote the corresponding equilibrium prices set by each firm type as  $\{p_1, p_2\}$ . Each household must set its nominal money balances ( $m$ ) at the beginning of the month, before the local firm's type is revealed. Then, having observed the type of the local firm (but not of any other),

---

<sup>11</sup> In an alternative version of the model, available on request, firms behave monopolistically. There is no closed-form solution to this model, but simulations show that higher unanticipated inflation and consequently higher search erode the market power (and mark-ups) of the relatively high cost firms to a greater degree, again reducing price dispersion.

households can engage in search. Households living in the location of a high-cost firm may choose to search for a low-price location.

We assume that the *ex ante* probability of a firm facing high costs is 0.5, and that the marginal cost curve for such a firm is:

$$mc_1 = \frac{1 + \phi + \theta}{1 - \gamma \cdot x_1}, \quad \phi > 0, \quad \theta > 0, \quad \gamma > 0 \quad (\text{A1})$$

Here,  $x_1$  is the demand faced by all type-1 firms. The marginal cost curve for a low-cost firm is:

$$mc_2 = \frac{1 + \phi}{1 - \gamma \cdot x_2} \quad (\text{A2})$$

Here,  $x_2$  is the demand faced by all type-2 firms. Since money balances are chosen before the firm type is known,  $m$  is the same in all locations. We will interpret changes in  $\theta$  as city-specific cost shocks and changes in  $\phi$  as common cost shocks.

Next, consider household search effort in a location where the firm faces higher costs this month. (Households in low-cost location have no incentive to search.) Having decided how hard to search, there is a certain probability of finding a low-cost location; this probability is denoted  $s$ . With probability  $(1 - s)$  the search for a lower price is unsuccessful, and the household purchases its goods from the local high-cost firm. Hence  $s$  also represents the fraction of households in a high-cost location that spend their money on goods produced in a different location. This means that demand for a firm in a high-cost location is:

$$x_1 = \frac{(1 - s) \cdot m}{p_1} \quad (\text{A3})$$

and demand for a firm in a low-cost location is:

$$x_2 = \frac{(1+s) \cdot m}{p_2} \quad (\text{A4})$$

Now, let the disutility from search effort associated with the probability  $s$  be equal to  $\beta \cdot s^2/2$ , and the utility from consumption ( $c$ ) be equal to  $\frac{c^{1-\lambda}}{1-\lambda}$ . Households in high-cost locations will choose search effort to maximize expected utility:

$$U(c, s) = E \left[ \frac{c^{1-\lambda}}{1-\lambda} - \frac{\beta \cdot s^2}{2} \right] \quad (\text{A5})$$

Having set  $m$  and subsequently observed the local firm price, households in high-cost locations will solve:

$$\max_s (1-s) \cdot \left( \frac{(m/p_1)^{1-\lambda}}{1-\lambda} \right) + s \cdot \left( \frac{(m/p_2)^{1-\lambda}}{1-\lambda} \right) - \frac{\beta \cdot s^2}{2} \quad (\text{A6})$$

There are two ways in which the aggregate price level affects the optimal amount of search effort. Firstly, a proportionate increase in both  $p_1$  and  $p_2$  will reduce the extra consumption that can be expected from a given increase in search effort. For a given marginal utility of consumption, this will lower the optimal amount of search effort. However, this effect is offset by the fact that a higher aggregate price level entails lower total consumption and therefore a higher marginal utility from consumption. If the increase in marginal utility is large enough, then there will be an increase in search effort. In order to see this, consider firstly the case in which  $\lambda = 0$ : marginal utility from consumption is constant. In this case, the optimal value of  $s$  is given by:

$$s = \left( \frac{1}{p_2} - \frac{1}{p_1} \right) \cdot \frac{m}{\beta} \quad (\text{A7})$$

A proportionate increase in both  $p_1$  and  $p_2$  will reduce  $s$ . Next, consider the case in which  $\lambda = 1$ : marginal utility declines moderately as consumption increases. In this case, the optimal value of  $s$  is given by:

$$s = \ln\left(\frac{p_1}{p_2}\right) \cdot \frac{1}{\beta} \quad (\text{A8})$$

A proportionate increase in both  $p_1$  and  $p_2$  will have no effect on  $s$ . Finally, consider the case in which  $\lambda = 2$ : marginal utility declines more rapidly as consumption increases. In this case, the optimal value of  $s$  is given by:

$$s = \frac{p_1 - p_2}{\beta \cdot m} \quad (\text{A9})$$

A proportionate increase in both  $p_1$  and  $p_2$  will increase  $s$ . In this case, shocks which increase aggregate inflation will increase the amount of search effort and reduce price dispersion in equilibrium. In order to see this, we can substitute equations (A3-A4) and equation (A9) into equations (A1-A2) with  $mc_1 = p_1$  and  $mc_2 = p_2$  to produce the following two price equations:

$$p_1 = \frac{(1 + \varphi + \theta) \cdot \beta/\gamma + \beta \cdot m + p_2}{1 + \beta/\gamma} \quad (\text{A10})$$

$$p_2 = \frac{(1 + \varphi) \cdot \beta/\gamma + \beta \cdot m + p_1}{1 + \beta/\gamma} \quad (\text{A11})$$

Let the unweighted consumer price index be  $p = [p_1 + p_2] / 2$ . Combining equations (A11) and (A12), we have:

$$p = \frac{1 + \varphi + \theta/2 + \gamma \cdot m}{(1 + \gamma/\beta) \cdot (1 - (1 + \beta/\gamma)^{-1})} \quad (\text{A12})$$

As we would expect, the aggregate price is increasing in  $\varphi$  and  $m$ : higher common costs and larger money balances are inflationary. With some manipulation, equations (A10-A11) also produce an equation for the relative price differential:

$$\frac{p_1 - p_2}{p_2} = \frac{\theta}{\theta \cdot \gamma / \beta + (1 + 2\gamma / \beta) \cdot (1 + \varphi + \gamma \cdot m)} \quad (\text{A13})$$

This shows that the differential is decreasing in  $\varphi$  and  $m$ . So, for example, shocks to common costs that raise aggregate prices also reduce relative price dispersion, if people have already chosen their money balances.<sup>12</sup>

---

<sup>12</sup> However, the negative correlation of  $(p_1 - p_2) / p_2$  with  $p$  does depend on the source of the inflationary shocks. For example, it requires that the variance of aggregate cost shocks ( $\varphi$ ) is sufficiently large relative to that of city-specific cost shocks ( $\theta$ ).



## Appendix 2: A Semi-parametric Model of RPV

As noted in the literature review, there is some diversity in the way that existing papers parameterize the relationship between RPV and anticipated inflation, and the quadratic and piecewise-linear functions in equations (10-11) do not encompass all of them. (For example, these equations do not allow for a V-shaped function with a turning point at a positive inflation rate.) For this reason, we also fit a semi-parametric model similar to the ones used by Fielding and Mizen (2008) and Choi (2010). In this model, the parameterizations of the  $\beta$  function in equations (10-11) are replaced by a non-parametric estimate of the function, using the method described by Robinson (1988) and Härdle (1990, chapter 9.1). Robust estimation of a semi-parametric model requires a large number of observations, so now the data are pooled across all grocery items, and the following regression equation is fitted to the panel:

$$v_{it}^D = \alpha_0 + \alpha_1 \cdot v_{it-1}^D + \alpha_2 \cdot \pi_{it}^{U+} + \alpha_3 \cdot \pi_{it}^{U-} + \alpha_4 \cdot t + \beta(\pi_{it}^A) + \alpha_5 \cdot h_{it} \quad (\text{A14})$$

$$+ \eta_1 \cdot v_{i0}^D + \eta_2 \cdot \pi_i^{U+} + \eta_3 \cdot \pi_i^{U-} + \eta_4 \cdot \pi_i^A + \eta_5 \cdot h_i + u_{it}$$

The second row of the equation contains a term in the initial value of RPV, and terms in the mean values of the different regressors:  $x_i = \frac{1}{T} \sum_t x_{it}$ . These terms are included to control for unobserved heterogeneity across the different grocery items.

The first step in fitting equation (A1) to the data is to create transformed regressors that are orthogonal to  $\pi_{it}^A$ . This is achieved by fitting a non-parametric regression equation for each of the regressors other than  $\pi_{it}^A$ :

$$x_{it} = \beta_x(\pi_{it}^A) + \tilde{x}_{it} \quad (\text{A15})$$

Here,  $\tilde{x}_{it}$  is a regression residual. The non-parametric  $\beta_x(\cdot)$  function is fitted in the same way as the  $\beta(\cdot)$  function described below. The  $\alpha$  and  $\eta$  parameters in equation (A1) are then estimated using the following regression equation:

$$v_{it}^D = \alpha_0 + \alpha_1 \cdot \tilde{v}_{it-1}^D + \alpha_2 \cdot \tilde{\pi}_{it}^{U+} + \alpha_3 \cdot \tilde{\pi}_{it}^{U-} + \alpha_4 \cdot \tilde{t} + \alpha_5 \cdot \tilde{h}_{it} + \eta_1 \cdot \tilde{v}_0^D + \eta_2 \cdot \tilde{\pi}_i^{U+} + \eta_3 \cdot \tilde{\pi}_i^{U-} + \eta_4 \cdot \tilde{\pi}_i^A + \eta_5 \cdot \tilde{h}_i + \varepsilon_{it} \quad (\text{A16})$$

Here,  $\varepsilon_{it}$  is a regression residual. Finally, the shape of  $\beta(\cdot)$  is estimated using the following non-parametric regression equation:

$$\varepsilon_{it} = \beta(\pi_{it}^A) + u_{it} \quad (\text{A17})$$

There are several different kernel density estimators that could be used to estimate the shape of  $\beta(\cdot)$ . The results reported below are based on one particular kernel density function, but results using other kernel density functions that are robust to outliers (such as the Epanechnikov Kernel) produce similar results.<sup>13</sup> First, we choose specific values of anticipated inflation at which the derivative of  $\beta(\cdot)$  is to be estimated. These values are equidistant points within the observed range of  $\pi_{it}^A$ . (The estimate at each point is independent of the others; enough points are chosen for the shape of  $\beta(\cdot)$  to be clear.) At any particular point  $\pi_0$ , the derivative  $\beta_0$  is estimated by fitting a linear regression equation using Weighted Least Squares. The regression equation is:

$$\varepsilon_{it} = \alpha_0 + \beta_0 \cdot \pi_{it}^A + u_{it} \quad (\text{A18})$$

and the weights  $w_{it}$  are as follows:

---

<sup>13</sup> The kernel density function here is used for example in Deaton and Paxson (1998). See Fan (1992, 1993) for a discussion of the properties of alternative kernel density functions.

$$w_{it} = \frac{15}{16} \left( 1 - \left( \frac{\pi_0 - \pi_{it}}{4z} \right)^2 \right)^2 \quad \text{if } |\pi_0 - \pi_{it}| < 4z, \text{ else } w_{it} = 0 \quad (\text{A19})$$

Here,  $z$  is a smoothing parameter, and the truncation of the weighting function at  $\pi_0 \pm 4z$  ensures that extreme outliers do not influence the estimates. The standard error of  $\beta_0$  is estimated using a bootstrap with 100,000 replications.

Figures A1-A3 illustrate the derivative of the  $\beta(\cdot)$  function at different anticipated inflation rates for alternative values of  $z$  between 1% and 2%,<sup>14</sup> along with the corresponding standard error bars. (The ‘butterfly shape’ of the error bars arises from the fact that there are fewer observations at more extreme values of anticipated inflation.) Note that the figures are drawn to different scales, so that each function occupies the whole chart and its shape is clear. As the value of  $z$  increases, the function becomes flatter but the error bars become smaller. In principle, it is possible to select an ‘optimal’ value of  $z$  based on an in-sample forecast error criterion. However, in our case the mean squared forecast error changes very little within the range of  $z$  shown.<sup>15</sup> Nevertheless, all of the figures show a line that is approximately straight, that is, a  $\beta(\cdot)$  function that is approximately quadratic. The turning point of the  $\beta(\cdot)$  function is indicated by the point at which the line in the figure crosses the  $y$ -axis; this is always at a positive anticipated inflation rate. Comparing Figures A1-A3 with Figure 7 in the main text, the imposition of a quadratic functional form produces a curve similar to the one in the semi-parametric model with  $z = 1\%$ . The error bars associated with the semi-parametric model are smaller, so a turning point at zero can be rejected with more confidence.

---

<sup>14</sup> Recall that the inflation data are monthly, so typical absolute anticipated inflation rates are less than 1%.

<sup>15</sup> For smaller values of  $z$ , the forecast errors are larger.

Table A1 reports estimated values of  $\alpha_2$  and  $\alpha_3$  (the unanticipated inflation parameters) for the different values of  $z$ , along with the corresponding t-ratios. Other parameter estimates are available on request. The parameter estimates are not very sensitive to the choice of  $z$ ; they have the same sign as the estimates reported in Table 3 of the main text, and are significantly different from zero. Their absolute value is somewhat smaller than in Table 3, and in the case of  $\alpha_3$  this difference is statistically significant. However, the overall conclusions regarding the effect of unanticipated inflation on RPV are unchanged.

It is also possible to fit a semi-parametric model to two different sub-samples comprising the storable and non-storable items. This produces results that are again broadly consistent with the quadratic model reported in the main text.

## References

- Deaton, A. and Paxson, C. (1998) 'Economies of scale, household size, and the demand for food,' *Journal of Political Economy* 106: 897-930.
- Fan, J. (1992) 'Design-adaptive nonparametric regression,' *Journal of the American Statistical Association* 87: 998-1004.
- Fan, J. (1993) 'Local linear regression smoothers and their minimax efficiencies,' *Annals of Statistics* 21: 196-216.
- Härdle, W. (1990) *Applied Nonparametric Regression*, Cambridge University Press: Cambridge, UK.
- Robinson, P. (1988) 'Root-N-consistent semiparametric regression,' *Econometrica* 56: 931-54.

**Table 1.** Cities Included in the Data Set

<i>City</i>	<i>Province</i>	<i>City</i>	<i>Province</i>
Amherst	Nova Scotia	Stratford	Ontario
Halifax	Nova Scotia	Sudbury	Ontario
New Glasgow	Nova Scotia	Timmins	Ontario
Sydney	Nova Scotia	Toronto	Ontario
Truro	Nova Scotia	Windsor	Ontario
Windsor	Nova Scotia	Woodstock	Ontario
Charlottetown	Prince Edward Island	Hull	Quebec
Bathurst	New Brunswick	Montreal	Quebec
Fredericton	New Brunswick	Quebec	Quebec
Moncton	New Brunswick	Saint Hyacinthe	Quebec
Saint John	New Brunswick	Saint John's	Quebec
Belleville	Ontario	Sherbrooke	Quebec
Brantford	Ontario	Sorel	Quebec
Brockville	Ontario	Thetford Mines	Quebec
Chatham	Ontario	Trois-Rivières	Quebec
Cobalt	Ontario	Brandon	Manitoba
Fort William	Ontario	Winnipeg	Manitoba
Galt	Ontario	Moose Jaw	Saskatchewan
Guelph	Ontario	Prince Albert	Saskatchewan
Hamilton	Ontario	Regina	Saskatchewan
Kingston	Ontario	Saskatoon	Saskatchewan
Kitchener	Ontario	Calgary	Alberta
London	Ontario	Drumheller	Alberta
Niagara Falls	Ontario	Edmonton	Alberta
North Bay	Ontario	Lethbridge	Alberta
Orillia	Ontario	Medicine Hat	Alberta
Oshawa	Ontario	Fernie	British Columbia
Ottawa	Ontario	Nanaimo	British Columbia
Owen Sound	Ontario	Nelson	British Columbia
Peterborough	Ontario	New Westminster	British Columbia
Port Arthur	Ontario	Prince Rupert	British Columbia
Saint Catharine's	Ontario	Trail	British Columbia
Saint Thomas	Ontario	Vancouver	British Columbia
Sarnia	Ontario	Victoria	British Columbia
Sault Sainte Marie	Ontario		

**Table 2.** Grocery Items Included in the Data Set

<i>non-storables</i>	<i>storables</i>
Bacon (unsliced)	Coffee
Bacon (sliced)	Corn (canned)
Butter (creamery)	Corn syrup
Butter solids	Currants
Cheese	Flour
Eggs (cooking)	Peaches (canned)
Eggs (fresh)	Peas (canned)
Finnan haddie	Prunes
Ham (sliced)	Raisins
Lard	Rice
Leg of lamb	Rolled oats
Milk	Salmon (canned)
Mutton leg roast	Sugar (granulated)
Onions	Sugar (yellow)
Potatoes (15lb bag)	Tapioca
Potatoes (100lb bag)	Tea
Rib roast	Tomatoes (canned)
Round steak	
Salt cod	
Salt mess pork	
Shoulder roast	
Sirloin steak	
Soda biscuits	
Stewing beef	
Veal shoulder	

**Table 3.** Average Coefficient Values in the Parametric Models of  $v_{it}^D$  (Using Untrimmed Inflation)

*T-ratios are reported in italics.*

	A			B			C			D		
	all items	non-storables	storables	all items	non-storables	storables	all items	non-storables	storables	all items	non-storables	storables
$v_{it-1}^D$	0.757 <i>110.535</i>	0.724 <i>83.930</i>	0.832 <i>101.329</i>	0.756 <i>110.352</i>	0.723 <i>83.941</i>	0.832 <i>101.293</i>	0.758 <i>110.797</i>	0.725 <i>84.061</i>	0.833 <i>101.257</i>	0.756 <i>110.339</i>	0.723 <i>83.841</i>	0.833 <i>101.217</i>
$\pi_{it}^A$	-0.332 <i>-0.793</i>	-0.508 <i>-2.946</i>	-0.766 <i>-0.847</i>				-0.324 <i>-0.776</i>	-0.537 <i>-3.127</i>	-0.842 <i>-0.932</i>			
$100 \cdot (\pi_{it}^A)^2$	2.119 <i>2.964</i>	0.520 <i>2.964</i>	2.845 <i>1.870</i>				2.035 <i>2.848</i>	0.405 <i>2.449</i>	2.300 <i>1.525</i>			
$\pi_{it}^{A+}$				0.127 <i>0.307</i>	0.678 <i>2.188</i>	-1.251 <i>-1.362</i>				0.016 <i>0.037</i>	0.605 <i>1.890</i>	-1.267 <i>-1.357</i>
$\pi_{it}^{A-}$				-2.090 <i>-6.852</i>	-1.617 <i>-5.413</i>	-2.356 <i>-4.256</i>				-2.018 <i>-6.516</i>	-1.590 <i>-5.151</i>	-2.243 <i>-4.033</i>
$\pi_{it}^{U+}$	-0.505 <i>-5.581</i>	-0.728 <i>-9.311</i>	-0.112 <i>-0.633</i>	-0.513 <i>-5.688</i>	-0.733 <i>-9.420</i>	-0.107 <i>-0.601</i>	-0.513 <i>-5.659</i>	-0.731 <i>-9.352</i>	-0.122 <i>-0.687</i>	-0.514 <i>-5.694</i>	-0.736 <i>-9.483</i>	-0.108 <i>-0.604</i>
$\pi_{it}^{U-}$	-0.922 <i>-9.922</i>	-0.911 <i>-10.334</i>	-0.936 <i>-5.052</i>	-0.924 <i>-9.982</i>	-0.899 <i>-10.275</i>	-0.944 <i>-5.088</i>	-0.911 <i>-9.788</i>	-0.897 <i>-10.162</i>	-0.935 <i>-5.037</i>	-0.915 <i>-9.890</i>	-0.889 <i>-10.173</i>	-0.950 <i>-5.111</i>
$100 \cdot (h_{it})^2$	-3.896 <i>-4.686</i>	-3.208 <i>-2.988</i>	-4.686 <i>-4.030</i>	-2.679 <i>-4.686</i>	-1.807 <i>-2.988</i>	-3.857 <i>-4.030</i>						
$100 \cdot h_{it}$							-0.125 <i>-3.484</i>	-0.148 <i>-2.503</i>	-0.101 <i>-3.951</i>	-0.083 <i>-4.087</i>	-0.088 <i>-2.963</i>	-0.082 <i>-3.541</i>
Average R <sup>2</sup>	0.781	0.732	0.852	0.781	0.733	0.852	0.781	0.732	0.851	0.781	0.733	0.851

**Table 4.** Average Coefficient Values in the Parametric Models  $v_u^D$  (Using Trimmed Inflation)

*T-ratios are reported in italics.*

	A			B			C			D		
	all items	non-storables	storables	all items	non-storables	storables	all items	non-storables	storables	all items	non-storables	storables
$v_{it-1}^D$	0.758 <i>111.064</i>	0.726 <i>84.430</i>	0.831 <i>101.126</i>	0.757 <i>110.711</i>	0.724 <i>84.137</i>	0.832 <i>101.050</i>	0.758 <i>110.928</i>	0.725 <i>84.415</i>	0.831 <i>100.945</i>	0.757 <i>110.653</i>	0.724 <i>84.136</i>	0.832 <i>100.906</i>
$\pi_{it}^A$	-0.272 <i>-0.649</i>	-0.518 <i>-3.023</i>	-0.689 <i>-0.762</i>				-0.293 <i>-0.701</i>	-0.515 <i>-3.008</i>	-0.822 <i>-0.909</i>			
$100 \cdot (\pi_{it}^A)^2$	2.150 <i>3.005</i>	0.391 <i>2.289</i>	2.983 <i>1.961</i>				2.033 <i>2.845</i>	0.343 <i>2.121</i>	2.384 <i>1.580</i>			
$\pi_{it}^{A+}$				0.080 <i>0.193</i>	0.428 <i>1.404</i>	-1.180 <i>-1.284</i>				0.076 <i>0.182</i>	0.397 <i>1.269</i>	-1.107 <i>-1.187</i>
$\pi_{it}^{A-}$				-2.068 <i>-6.754</i>	-1.481 <i>-4.942</i>	-2.412 <i>-4.345</i>				-2.052 <i>-6.629</i>	-1.451 <i>-4.711</i>	-2.373 <i>-4.263</i>
$\pi_{it}^{U+}$	-0.564 <i>-6.003</i>	-0.749 <i>-8.977</i>	-0.202 <i>-1.114</i>	-0.568 <i>-6.069</i>	-0.754 <i>-9.066</i>	-0.185 <i>-1.017</i>	-0.568 <i>-6.035</i>	-0.753 <i>-9.020</i>	-0.208 <i>-1.143</i>	-0.564 <i>-6.026</i>	-0.758 <i>-9.138</i>	-0.181 <i>-0.996</i>
$\pi_{it}^{U-}$	-0.933 <i>-9.901</i>	-0.929 <i>-10.250</i>	-0.924 <i>-4.952</i>	-0.925 <i>-9.851</i>	-0.917 <i>-10.186</i>	-0.936 <i>-5.001</i>	-0.931 <i>-9.874</i>	-0.928 <i>-10.233</i>	-0.925 <i>-4.947</i>	-0.924 <i>-9.842</i>	-0.915 <i>-10.179</i>	-0.939 <i>-5.002</i>
$100 \cdot (h_{it})^2$	-3.681 <i>-4.892</i>	-2.572 <i>-2.618</i>	-4.580 <i>-4.375</i>	-2.788 <i>-4.400</i>	-1.759 <i>-2.307</i>	-3.704 <i>-3.850</i>						
$100 \cdot h_{it}$							-0.113 <i>-3.200</i>	-0.084 <i>-1.463</i>	-0.099 <i>-3.848</i>	-0.091 <i>-2.418</i>	-0.058 <i>-0.918</i>	-0.080 <i>-3.424</i>
Average R <sup>2</sup>	0.770	0.721	0.842	0.770	0.721	0.841	0.770	0.721	0.842	0.770	0.721	0.841



**Table A1.** Unanticipated Inflation Coefficients in the Semi-parametric Model of  $v_{it}^D$

*T-ratios are in italics.*

	$z = 1.0$	$z = 1.5$	$z = 2.0$
$\pi_{it}^{U+}$	-0.361 <i>-6.311</i>	-0.314 <i>-6.035</i>	-0.300 <i>-6.044</i>
$\pi_{it}^{U-}$	-0.503 <i>-7.684</i>	-0.601 <i>-10.086</i>	-0.616 <i>-10.841</i>

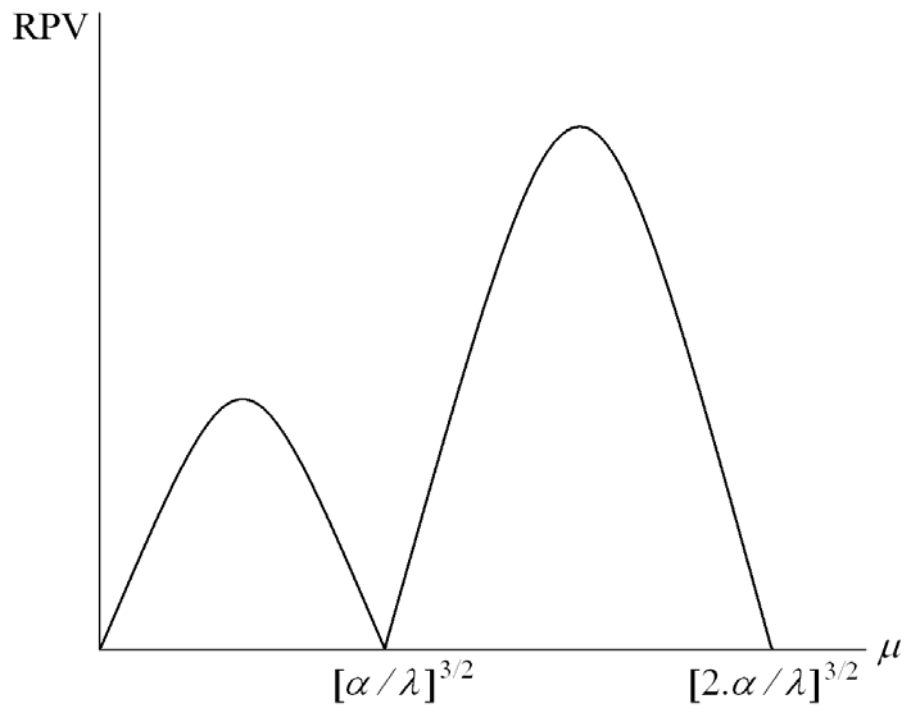


Figure 1: Inflation and RPV in the Rotemberg-Danziger Model

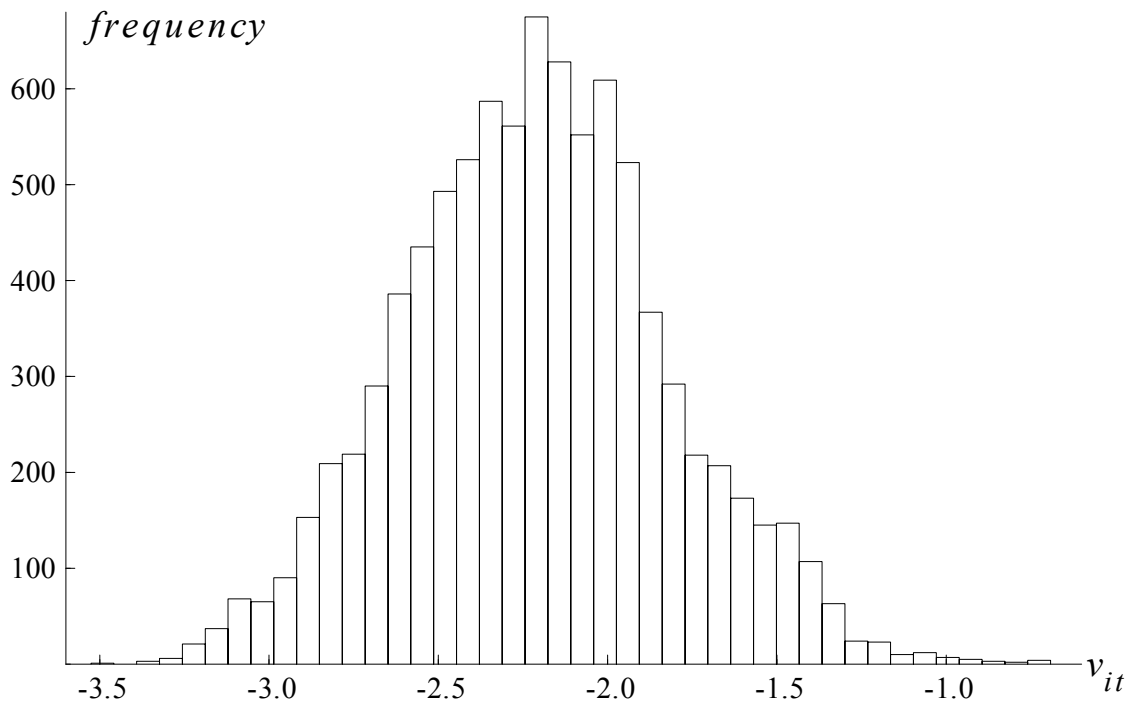


Figure 2: distribution of relative price variability ( $v_{it}$ )

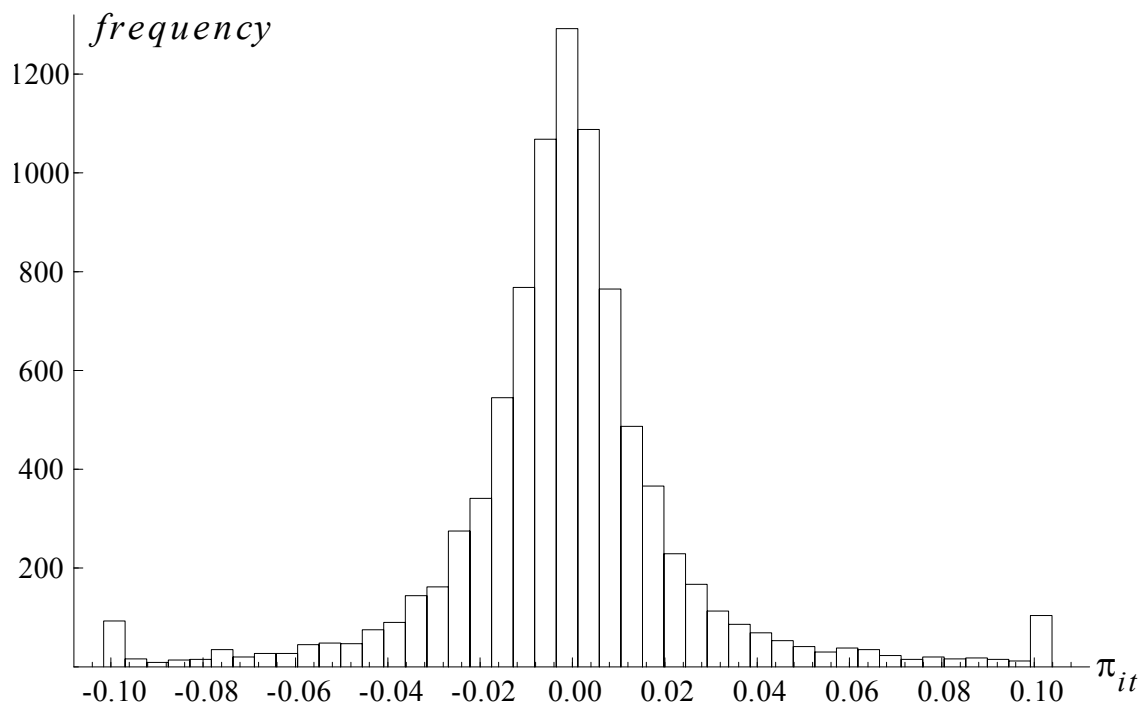


Figure 3: distribution of inflation ( $\pi_{it}$ ) trimmed at  $\pm 10\%$

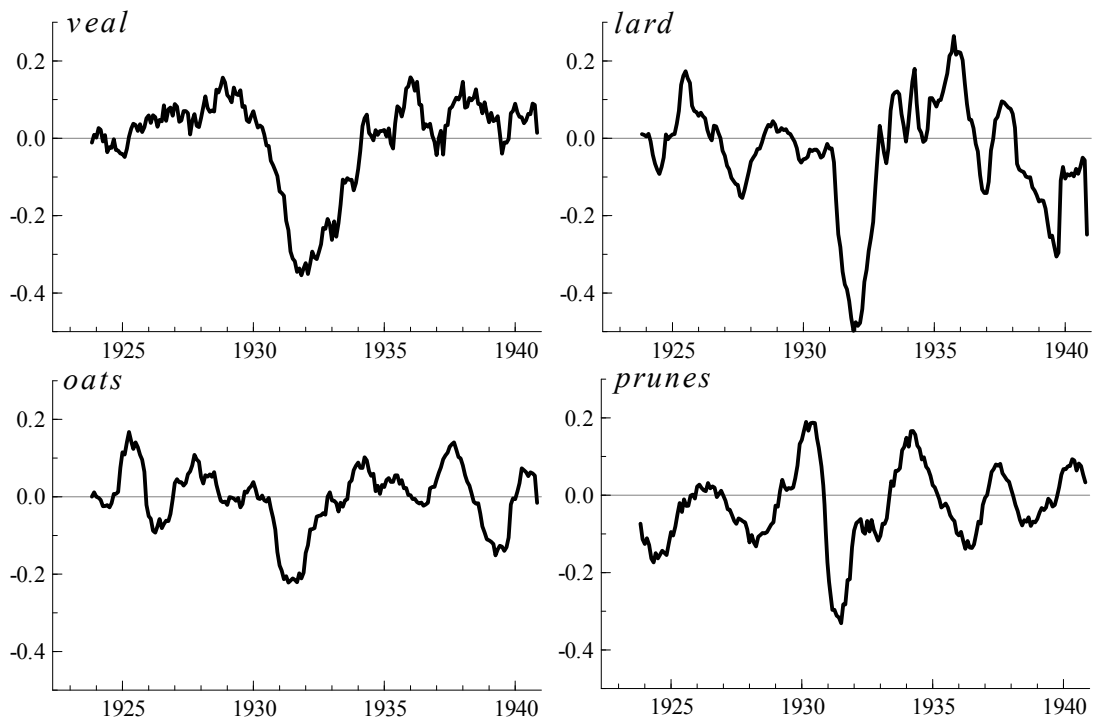


Figure 4: annual inflation  $\sum_{k=0}^{k=11} \pi_{it-k}$  for selected items

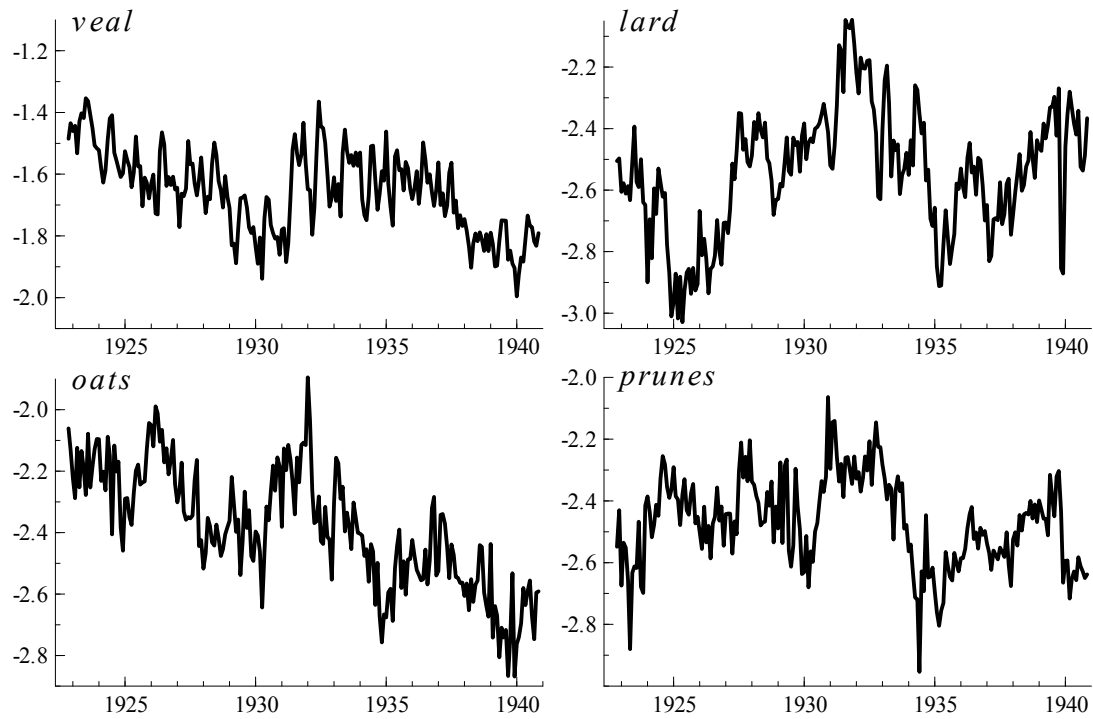


Figure 5: relative price variability  $v_{it}$  for selected items



Figure 6: annual inter-war Canadian consumer price inflation and real GNP growth rates

Source: *Historical Statistics of Canada* ([www.statcan.gc.ca](http://www.statcan.gc.ca))

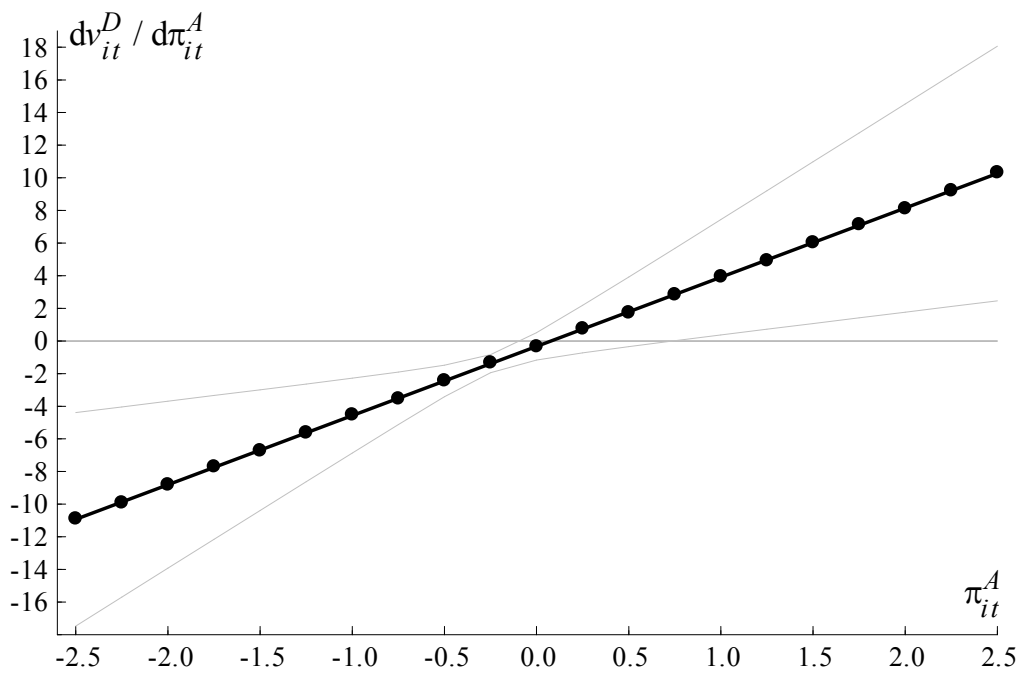


Figure 7: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (quadratic model, all items)

*Inflation ( $\pi^A$ ) is measured in percentage points per month.*

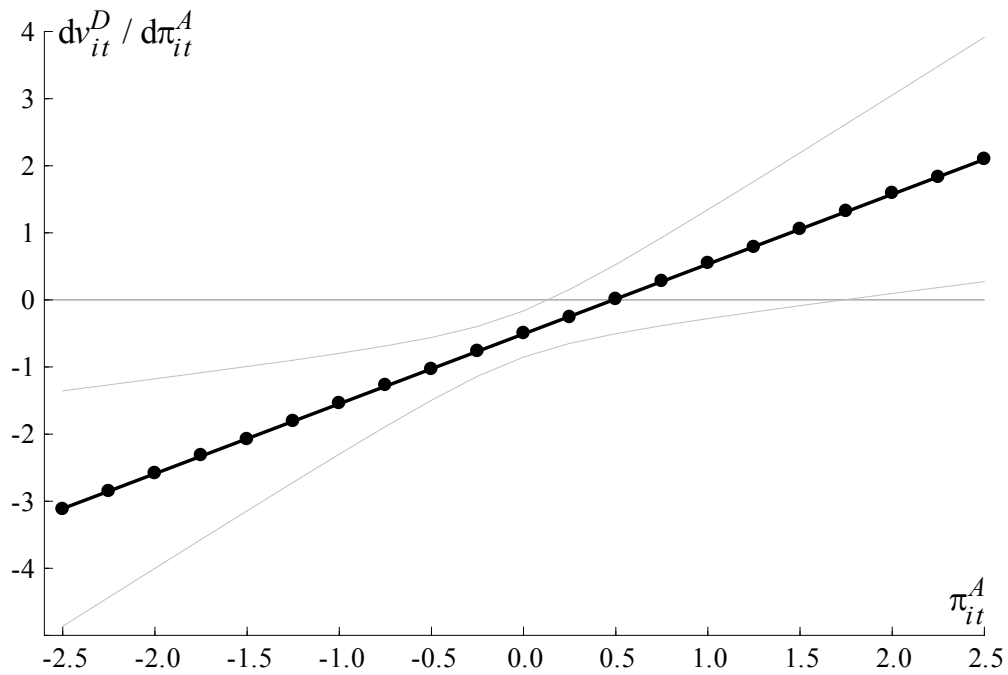


Figure 8: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (quadratic model, non-storables)

*Inflation ( $\pi^A$ ) is measured in percentage points per month.*

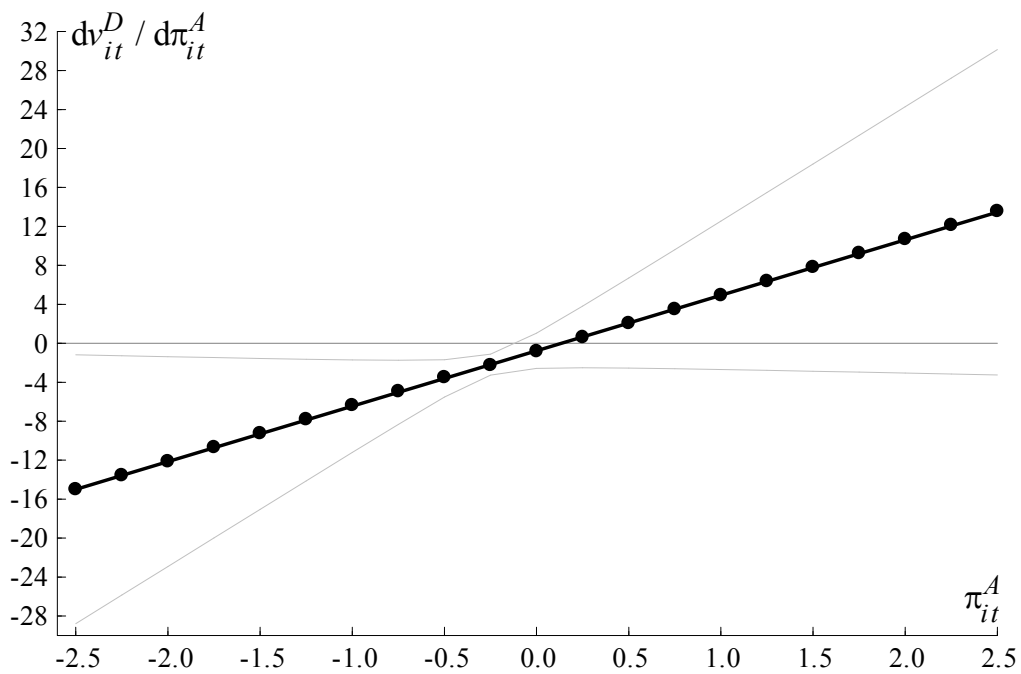


Figure 9: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (quadratic model, storables)

*Inflation ( $\pi^A$ ) is measured in percentage points per month.*

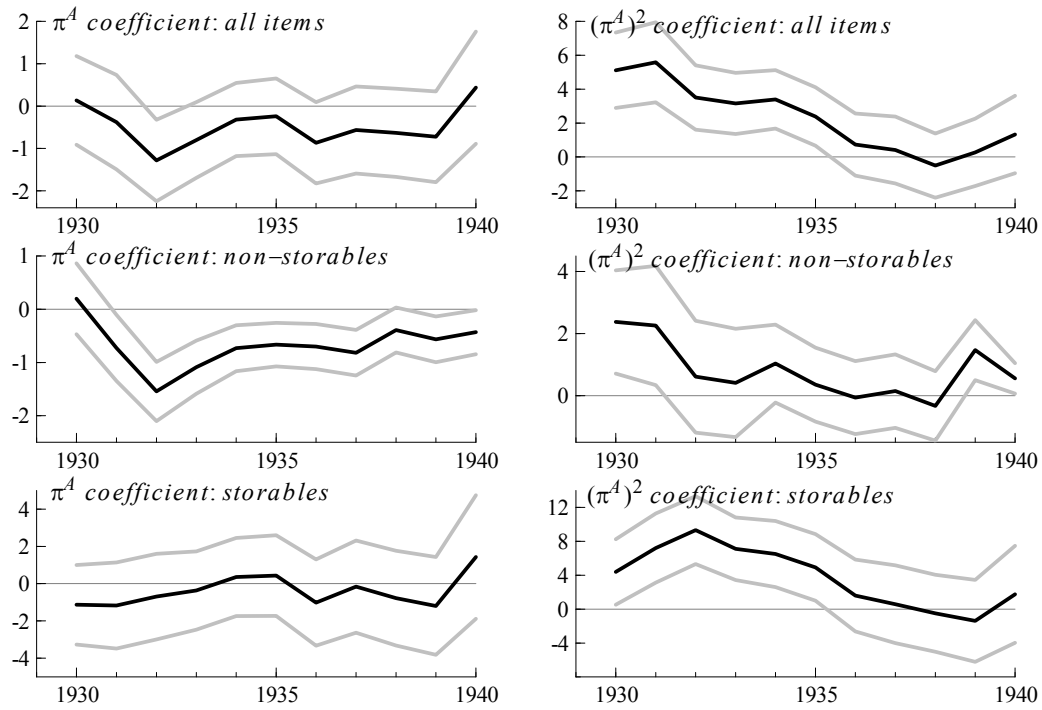


Figure 10: recursive parameter estimates for the  $\pi^A$  coefficients with a 95% confidence interval

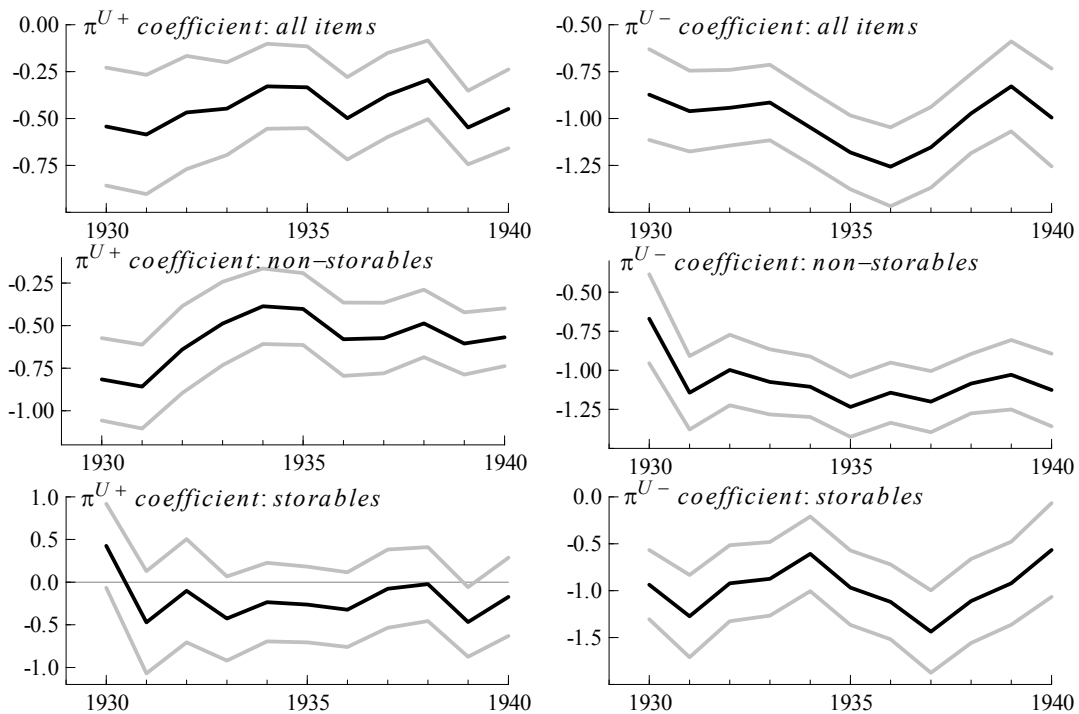


Figure 11: recursive parameter estimates for the  $\pi^U$  coefficients with a 95% confidence interval

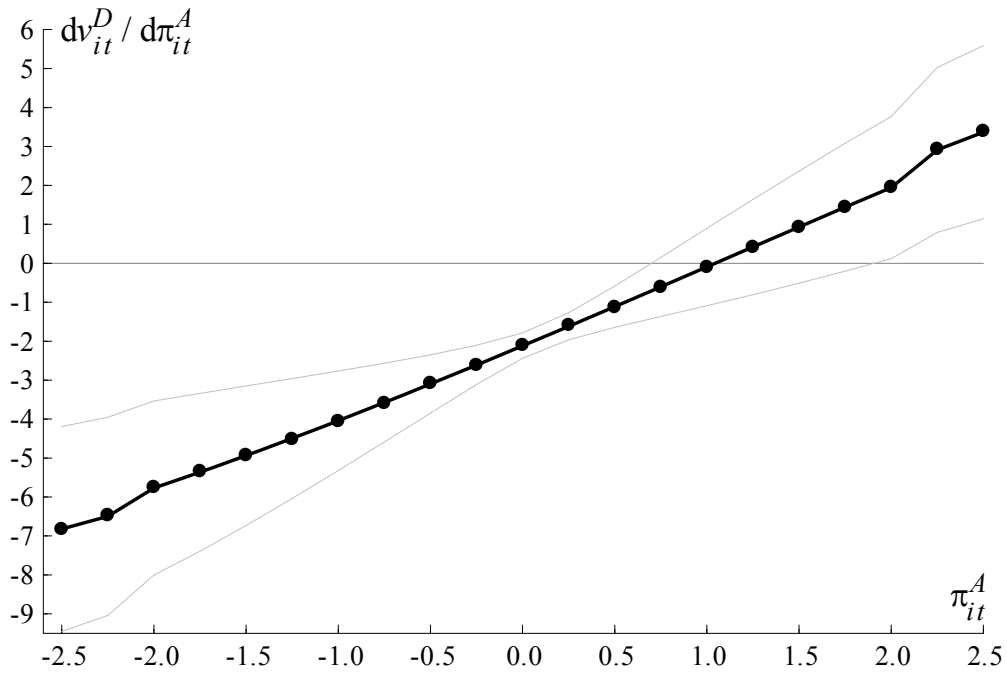


Figure A1: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (semi-parametric model,  $h = 1\%$ )

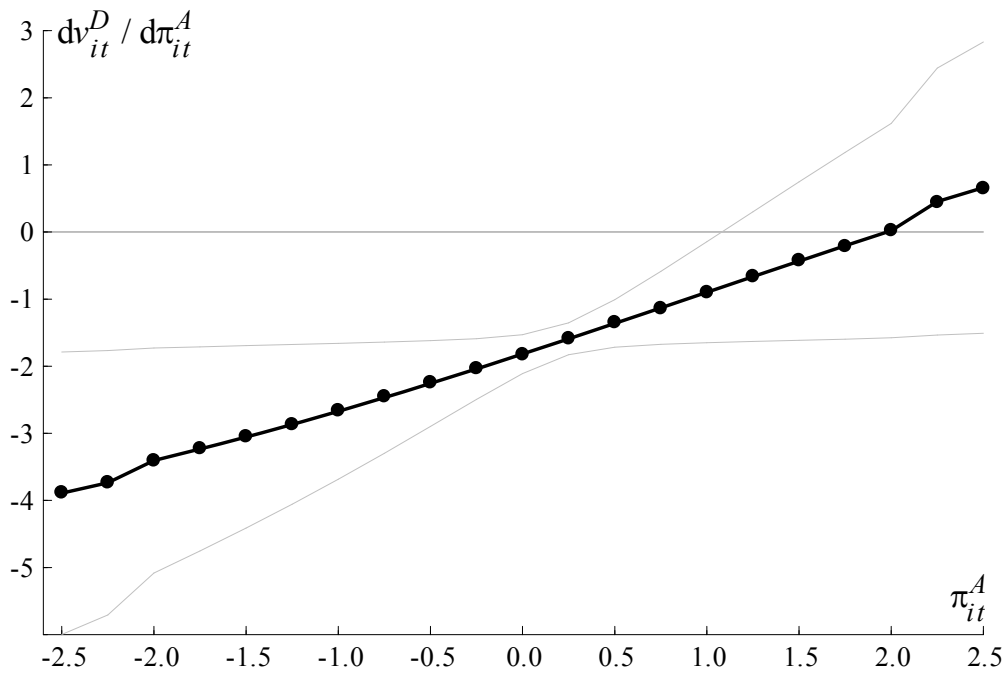


Figure A2: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (semi-parametric model,  $h = 1.5\%$ )



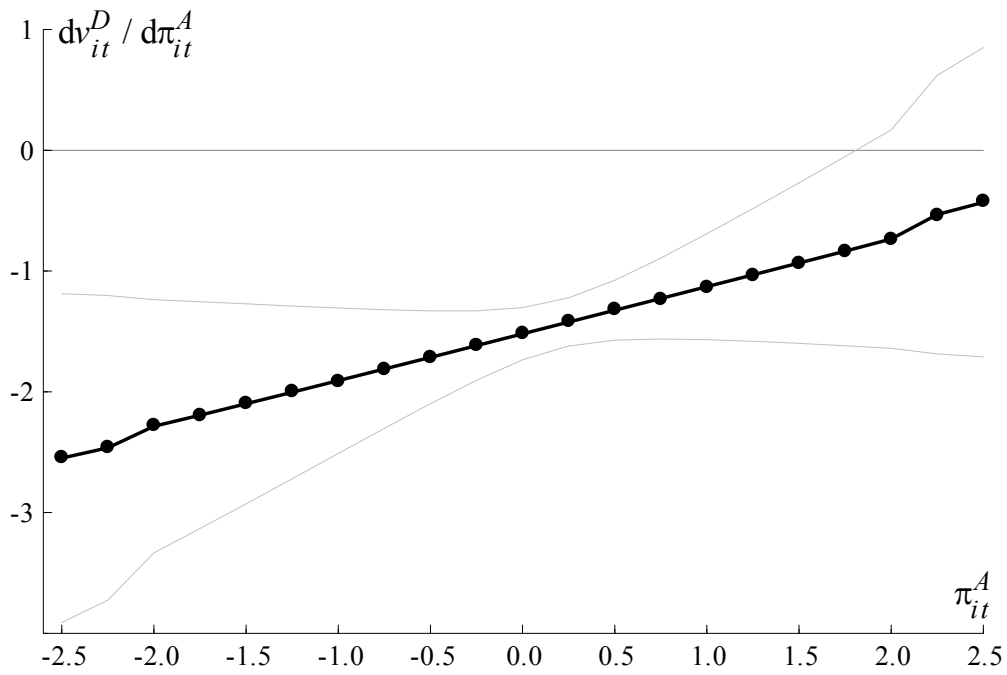


Figure A3: values of  $dv_{it}^D / d\pi_{it}^A$  with a 95% confidence interval (semi-parametric model,  $h = 2\%$ )