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Taking prices seriously in the measurement of inequality

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Abstract

In the measurement of inequality, adjustments for differences across households in their demographic composition and in the price regimes they face are usually very simple. Often, nominal expenditure (or income) is adjusted with an expenditure-independent price deflator and a price-independent equivalence scale. I show that using more flexible expenditure-dependent price deflators and price-dependent equivalence scales affects the level of, and trend in, measured family expenditure inequality in Canada over 1969–1997. For example, standard methods show a significant decrease in inequality between 1969 and 1978, but more flexible methods show a significant increase in inequality over this period.

JEL classification: D63; D31; D11

1. Introduction

Research which seeks to measure economic inequality usually ignores relative prices. However, relative price changes may affect poor households differently from rich households and small households differently from large households. While these may affect the distribution of well-being, most analyses of economic inequality incorporate them in simple and potentially misleading ways. Adjustment for differences in the prices faced by households is typically made by dividing nominal expenditure (or income) by an expenditure-independent price deflator. This price deflator is typically computed at the level of the country even though

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there is substantial variation in prices across regions in the country. Adjustment for demographic differences across households is typically made by dividing expenditure (or income) by a price-independent equivalence scale. This paper shows that using more general procedures changes measured inequality in Canada over 1969–1997.

In particular, I explore two refinements to the measurement of family expenditure inequality: (1) I let the price deflator depend on total expenditure (and of course relative prices), allowing rich and poor families to respond to price changes differently; and (2) I let the equivalence scale depend on relative prices, allowing large and small families to respond to price changes differently. Further, I allow both price deflators and equivalence scales to respond to variation in prices across regions and over time.

These refinements to the assessment of expenditure inequality in Canada over 1969–1997 make a difference, especially at the beginning and end of this period. Standard methods indicate that the Gini coefficient for family expenditure inequality decreased significantly by 0.4 percentage points between 1969 and 1978 and increased significantly by 0.6 percentage points between 1978 and 1982. More general methods indicate that the Gini coefficient increased significantly by 0.5 percentage points between 1969 and 1978, and decreased insignificantly by 0.2 percentage points between 1978 and 1982, reversing the time-pattern shown by standard methods. In the 1990s, standard methods show a significant decline in the the Gini coefficient of 0.4 percentage points between 1992 and 1997, but more general methods show no significant difference between the Gini coefficients for these two years.

2. Theory

Define the expenditure function $E(p,u,z)$ as the minimum expenditure necessary to give each member[1] of a household with demographic characteristics z facing prices $p = (p_1,..,p_m)$[2] a utility level of u. Define the indirect utility function $V(p,x,z)$ as the utility attained by each member of a household with expenditure x and demographic characteristics z facing prices p. Define $\bar{p} = 1_m \cdot 100$ as a base vector of prices, to be used as a basis for price adjustments. Define z^R as reference vector of household characteristics, to be used as the reference for demographic

[1] I assume for this paper that utility is equally distributed within households. This would result if households maximised a maximin household welfare function over individual member utilities, or if u gave the equally-distributed equivalent utility for the household.

[2] In the empirical work in this paper, I do not account for one very important price, that of leisure. Apps and Savage (1989) take account of leisure in their assessment of inequality among married dual earner couples. However, because many people do not work, and therefore do not have an observed price for their leisure, I do not take leisure into account in this paper.

adjustments, and let z^R denote a single childless adult. Define $V^R(p, x) = V(p, x, z^R)$ as the indirect utility function and $E^R(p, u) = E(p, u, z^R)$ as the expenditure function of the reference household type.

The measurement of inequality requires an expenditure measure that has been adjusted for differences in the demographic characteristics of, and price regimes faced by, various households. If it is possible to 'adjust' for differences in demographics and prices across households, then the adjusted distribution with no differences in demographic characteristics or price regimes must be equivalent in some way to the unadjusted distribution with differences in characteristics and/or price regimes. With consequentialist distributive ethics, for example utilitarianism, two distributions are equivalent if the set of individual utilities in each distribution is identical. Thus, define *adjusted expenditure* denoted $y = Y(p,x,z)$ as the expenditure that would give the reference household facing base prices the same utility as each member of a household with expenditure x and characteristics z facing prices p. Donaldson (1992) shows that adjusted expenditure[3] is given by

$$y = Y(p, x, z) = E^R(\bar{p}, u) = E^R(\bar{p}, V(p, x, z)). \tag{1}$$

To measure inequality, assign adjusted expenditure y to each member of each household and calculate inequality measures, such as Gini coefficients, on the basis of these adjusted expenditure data. Donaldson (1992) cautions that welfare and inequality measures computed from adjusted expenditure will in general depend on the choice of base price vector $\bar{p}$, and suggests that utility $u = V(p, x, z)$ should be used in welfare measurement, rather than proxies such as y. However, in the absence of a cardinalisation of V, adjusted expenditure y is certainly an improvement over nominal expenditure x.

Define the price deflator D to give the ratio of reference expenditure to reference expenditure at base prices, so that $D(p, u) = E^R(p, u)/E^R(\bar{p}, u)$. Note that $D(p, u)$ can be expressed as a function of x rather than u by substituting indirect utility V for utility u. Define the equivalence scale S to give the ratio of expenditures to reference expenditures, so that $S(p, u, z) = E(p, u, z)/E^R(p, u)$. Since $x \equiv E(p, u, z)$ and $u \equiv V(p, x, z)$, we can write adjusted expenditure without loss of generality as

$$Y(p, x, z) = \frac{x}{D(p, u)S(p, u, z)}. \tag{2}$$

Almost all empirical work that measures economic inequality[4] uses a 'naive adjusted expenditure function', Y_N, of the form

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[3]Donaldson (1992) and King (1983) refer to adjusted expenditure as an 'extended money metric'.

[4]For example, several papers on consumption inequality use the naive adjusted expenditure function and assume that all observations in each year face the same prices: Barrett et al. (1999a); Pendakur (1998); Gouveia and Tavares (1995); Cutler and Katz (1992); Slesnick (1998); and Blundell and Preston (1998).

$$Y_N(p,x,z)=\frac{x}{D_N(p)S_N(z)} \tag{3}$$

where D_N is a 'naive price deflator' which depends only on prices and S_N is a 'naive equivalence scale' which depends only on demographic characteristics.

If the adjusted expenditure function is of the form (3), then the expenditure function must be decomposable as follows (see Diewert, 1993; Blackorby and Donaldson, 1993):

$$E(p,u,z)=D_N(p)f(u)S_N(z). \tag{4}$$

Here, preferences are identically homothetic across demographic types, so that expenditure share equations are independent of expenditure and demographic characteristics.

Diewert (1993) shows that if the naive price deflator D_N corresponds to the true price deflator, then expenditure share equations are independent of expenditure. Pendakur (1999) shows if the naive equivalence scale S_N corresponds to the true equivalence scale, then expenditure share equations must be identical across demographic types except for translation in $\ln x$ (scaling in expenditure).

A large body of empirical work suggests that expenditure shares for many commodities are highly dependent on expenditure and on demographic characteristics (see, for example, Banks et al., 1997). Thus, it is undesirable to use naive price deflators and equivalence scales in the measurement of inequality.

In this paper, I consider a 'flexible adjusted expenditure function', Y_F, of the form:

$$Y_F(p,x,z)=\frac{x}{D_F(p,u)S_F(p,z)}=\frac{x}{D_F(p,V(p,x,z))S_F(p,z)} \tag{5}$$

where D_F is a 'flexible price deflator' and S_F is a 'flexible equivalence scale'. Here, the price deflator depends on prices and utility (which depends on prices, expenditure and demographics), and the equivalence scale depends on prices and demographics[5]. I estimate a demand system to recover the functions D_F and S_F, and use these flexible deflators and scales to assess patterns in family expenditure inequality in Canada over 1969 to 1997.

The functional forms for D_F and S_F are determined by choice of a functional form for the expenditure function and the definitions of D and S. For this paper, I use a version of the Quadratic Almost Ideal (QAI) model (see Banks et al., 1997) in which the expenditure function is given by

[5]In empirical work on economic inequality, equivalence scales for demographic adjustment are almost always assumed to be expenditure-independent as well as price-independent. In this paper, I consider only the relaxation of price-independence. See Donaldson and Pendakur (1999) for a discussion of expenditure- and price-dependent equivalence scales.

$$\ln E(p,u,z) = \ln a(p,z) + \frac{b(p)u}{1+q(p)u}, \tag{6}$$

and the dual indirect utility function is given by

$$V(p,x,z) = \left(\left(\frac{\ln x - \ln a(p,z)}{b(p)}\right)^{-1} - q(p)\right)^{-1}, \tag{7}$$

and base prices satisfy the restrictions $b(\bar{p}) = 1$ and $q(\bar{p}) = 0$.

The flexible equivalence scale S_F associated with (6) is given by

$$\ln S(p,u,z) = \ln S_F(p,z) = \ln \frac{E(p,u,z)}{E^R(p,u)} = \ln a(p,z) - \ln a(p,z^R), \tag{8}$$

and the flexible price deflator D_F associated with (6) is given by

$$\ln D(p,u) = \ln D_F(p,u) = \ln \frac{E^R(p,u)}{E^R(\bar{p},u)}$$
$$= \ln a(p,z^R) - \ln a(\bar{p},z^R) + \frac{b(p)u}{1+q(p)u} - u. \tag{9}$$

The QAI demand system features expenditure share equations that are quadratic in ln x so that the homotheticity restrictions associated with naive price deflators are relaxed. Further, share equations exhibit 'shape invariance' (see Blundell et al., 1998 and Pendakur, 1999) which allows for expenditure share equations that vary across demographic groups by translations in ln x and the addition of equation-specific constants, so that the translation restrictions associated with naive equivalence scales are relaxed.

3. Data and demand system estimation

This analysis uses the 1969, 1978, 1982, 1986, 1992 and 1996 Canadian Family Expenditure Surveys and the 1997 Canadian Survey of Household Spending[6]. I focus the analysis on urban residents of Canada[7] for two reasons: (1) it does not

[6]The Survey of Household Spending includes observations of households which exist for only part of the year. Since the Family Expenditure Surveys contain only full-year observations, part-year observations are dropped from the 1997 data.

[7]Urban residents are defined as households living in cities with 30,000 or more residents. Urban residents account for most households in each sample year:

	1969	1978	1982	1986	1992	1996	1997
Proportion Urban	0.623	0.622	0.643	0.700	0.773	0.793	0.799

require the imputation of nonmarket consumption for nonurban residents; and (2) inter-regionally comparable price data for nonurban residents are not available (see details on price data below). In addition, some robustness testing is performed using both urban and nonurban residents of Canada.

In order to emphasise the difference that price dependence makes, I use only one household characteristic in equivalence scales — the number of people in the household. Thus, let z denote the number of household members[8] rather than a vector of household characteristics.

I estimate demand systems in nine commodities on a sample of families satisfying the following restrictions: (1) they rent their accomodation[9]; (2) they live in cities with 30,000 or more residents; (3) all household members were full-year members; and (4) they are households or spending units comprised of a single family[10]. The nine commodities are: (1) food purchased from stores; (2) restaurant food; (3) (rented) shelter; (4) household operation (including child care); (5) household furnishings and equipment; (6) clothing; (7) private transportation operation; (8) public transportation; and (9) personal care. These commodities are chosen because regional price data are available for them (see below) and because they are relatively nondurable given the one-year time span for the expenditure data. Table 1 gives summary statistics on these data. These nine expenditure categories account for approximately 80% of net income for households in the sample.

To estimate demand systems, price deflators and equivalence scales, a data set with price variation is essential. To this end, I use commodity- and region-specific price data compiled by Browning and Thomas (1998, 1999)[11] and rental price data from CMHC (1997). With the exception of the price of rental accomodation, inter-regionally comparable price indices for the above commodities for 1969 to 1996 are taken from Browning and Thomas. These price data are compiled for five regions of Canada: (1) Atlantic Canada (Newfoundland, Nova Scotia, New Brunswick and Prince Edward Island); (2) Quebec; (3) Ontario; (4) the Prairies (Saskatchewan, Manitoba and Alberta); and (5) British Columbia.

The Browning and Thomas series are based on a variety of sources, but the

[8]Household size can vary over the year if members leave or enter the household. I take household size to be the maximum size attained by the household during the year.

[9]The exclusion of home-owners, who typically have higher expenditures than renters, may introduce selection bias into the demand estimation. One alternative strategy is to estimate a conditional demand system, conditioning on home ownership and shelter expenditures (rather than on their prices). However, this estimation strategy does not permit the estimation of equivalence scales.

[10]The 1969, 1978, 1982 and 1986 surveys use the ‘‘spending unit’’ as the unit of analysis and the 1992, 1996 and 1997 surveys use the household as the unit of analysis. To create a consistent sample across these units of analysis, I use only spending units and households comprised of a single family, which is defined as an unattached individual, or a group of people, related by blood, marriage or adoption, who live together in a household.

[11]I am grateful to Martin Browning and Irene Thomas for their permission to use these price data.

Table 1
Demand data

Number of families: 19,526	Mean	Std Dev	Min	Max
Total expenditure	14,431	8588	503	110,555
Food at home share	0.211	0.100	0.015	0.089
Restaurant food share	0.060	0.063	0.000	0.642
Rent share	0.354	0.132	0.009	0.949
Household operation	0.080	0.046	0.000	0.635
Household furn & eq share	0.047	0.056	0.000	0.602
Clothing share	0.094	0.062	0.000	0.585
Transport operation share	0.088	0.086	0.000	0.591
Public transportation share	0.030	0.039	0.000	0.451
Personal care share	0.035	0.022	0.000	0.247
Number of family members	2.17	1.31	1	11

spirit is as follows: absolute commodity prices for various cities in Canada collected by Statistics Canada (catalogue #62-010) are aggregated (with population weighted geometric means) into regional commodity price vectors in a base year. These regional price vectors are back- and forward-dated using city- and region-specific commodity price indices. I update regional commodity prices to 1997 with population-weighted provincial commodity price indices (CANSIM, 2000)[12]. Because the basic price data are for cities, the Browning and Thomas price series more closely reflect urban than non-urban prices.

The prices of rental accomodation in each region in 1996 are weighted averages of rental prices reported in CMHC (1997) for bachelor, one-, two- and three-bedroom rentals in 66 Canadian cities[13]. Weights by number of rooms and region of residence are drawn from 1996 Public Use Census Family microdata (CENSUS, 1996). These inter-regionally comparable rental accomodation price indices are back- and forward-dated using population weighted provincial rental shelter price indices (CANSIM, 2000).

All prices are normalised so that residents of Ontario in 1982 face the base price vector $\bar{p} = 1_m \cdot 100$. Appendix Table A1 gives a complete listing of prices used in demand estimation.

The following parametric specifications for a, b and q are used:

[12]As of 1992, Statistics Canada's city-specific commodity price indices are unavailable, so I use province-specific commodity price indices to generate region-specific commodity price indices.

[13]CMHC reports rents only for cities with 50,000 or more residents. However, the public use expenditure data files only allow the identification of residents of cities with 30,000 or more residents in 1996 and 1997. Expenditure data for the period 1969 to 1992 allow the identification of residents of Canada's 15 largest cities, all of which have more than 100,000 residents. Re-estimation of the demand system and inequality measures with rental prices differing between cities with 30,000–99,999 residents and cities with 100,000 or more residents does not change the spirit of the results presented below.

$$\ln a(p,z) = \left(a_0^R + a_0^z \ln z\right) + \sum_{k=1}^{m} \left(a_k^R + a_k^z \ln z\right) \ln p_k + \frac{1}{2} \sum_{k=1}^{m} \sum_{l=1}^{m} \left(a_{kl}^R + a_{kl}^z \ln z\right) \ln p_k \ln p_l, \tag{10}$$

$$b(p) = \prod_{k=1}^{m} p_k^{b_k}, \tag{11}$$

and

$$q(p) = \sum_{k=1}^{m} q_k \ln p_k. \tag{12}$$

Homogeneity and symmetry require: $\Sigma_{k=1}^m a_k^z = \Sigma_{k=1}^m b_k = \Sigma_{k=1}^m q_k = 0$; $\Sigma_{k=1}^m a_k^R = 1$; $\Sigma_{l=1}^m a_{kl}^R = \Sigma_{l=1}^m a_{kl}^z = 0 \ \forall\ k$; $a_{kl}^R = a_{lk}^R\ \forall\ k, l$; and $a_{kl}^z = a_{lk}^z\ \forall\ k, l$. Here, $b(\bar{p}) = 1$ and $q(\bar{p}) = 0$.

Substituting (10), (11) and (12) into (7) and applying Roy's identity yields expenditure share equations w_j that are quadratic in the natural logarithm of expenditure:

$$w_j(p,x,z) = \left(a_k^R + a_k^z \ln z\right) + \sum_{k=1}^{m} \left(a_{jk}^R + a_{jk}^z \ln z\right) \ln p_j + b_j(\ln x - \ln a(p,z)) + \frac{q_j}{b(p)} (\ln x - \ln a(p,z))^2. \tag{13}$$

Adding an error term to the right hand side produces an estimable demand system. Estimation is by nonlinear least squares using the exact value of $a(p, z)$ at each iteration, rather than by linearised least squares using the Stone or other approximation of $a(p, z)$.

Substituting (10) into (8) gives the flexible equivalence scale:

$$\ln S_F(p,z) = a_0^z \ln z + \sum_{k=1}^{m} a_k^z \ln z \ln p_k + \frac{1}{2} \sum_{k=1}^{m} \sum_{l=1}^{m} a_{kl}^z \ln z \ln p_k \ln p_l, \tag{14}$$

$$\text{or } S_F(p,z) = z^{\left(a_0^z + \sum_{k=1}^{m} a_k^z \ln p_k + \frac{1}{2} \sum_{k=1}^{m} \sum_{l=1}^{m} a_{kl}^z \ln p_k \ln p_l\right)}. \tag{15}$$

Note that the exponent on z gives the elasticity of the equivalence scale — and therefore expenditures — with respect to household size.

Substituting (10), (11) and (12) into (9) gives the flexible price deflator:

$$\ln D_F(p,u) = \sum_{k=1}^{m} a_k^R \ln p_k + \frac{1}{2} \sum_{k=1}^{m} \sum_{l=1}^{m} a_{kl}^R \ln p_k \ln p_l + \frac{\left(\prod_{k=1}^{m} p_k^{b_k}\right) u}{1 + \left(\sum_{k=1}^{m} q_k \ln p_k\right) u} - u. \tag{16}$$

Manipulation of the expression for indirect utility, (7), gives

$$V(p, x, z) = \frac{\ln x - \ln a(p, z)}{b(p) - q(p)(\ln x - \ln a(p, z))}. \tag{17}$$

Thus, if V is defined, utility is zero if and only if $x = a(p, z)$. D_F is a complex function of prices and utility, but it simplifies considerably if $u = 0$: $\ln D_F(p, 0) = \ln a(p, z^R) - \ln a(\bar{p}, z^R)$. The elasticity of D_F with respect to x also simplifies if $u = 0$: $\partial \ln D_F(p, 0)/\partial \ln x = (b(p) - 1)/b(p)$. These features will be used to assess the estimated models.

To facilitate interpretation of estimated coefficients, I set the parameter a_0^R so that $u = 0$ for a household in the middle of the distribution of utility[14]. In particular, I set $a_0^R = \ln \bar{x} - \ln (100)$, where $\bar{x}$ is the average total expenditure of households with reference characteristics in Ontario 1982 (these households face the base price vector). Thus, any household with the same utility level as the average reference household in the base period has a utility level of zero, and has the simple flexible price deflator and price deflator elasticity given above.

Equivalence scales which depend on utility cannot in general be estimated from demand data (see Pollak and Wales, 1979)[15]. However, if the equivalence scale is assumed independent of utility, then Blackorby and Donaldson (1993) show that S_F can be estimated uniquely from demand data in the following sense: there is only one equivalence scale that is independent of u and is consistent with demand behaviour, and it can be estimated from the definition of S, which for the QAI is given by (15)[16].

Price deflators which depend on utility can be identified from demand data because z-specific monotonic transformations of utility which leave demands unaffected also leave price deflators unaffected. This is because the deflator is calculated only for the reference household, and all of the interpersonal comparisons are carried by the equivalence scale.

The QAI expenditure function can be restricted to satisfy the restrictions required for naive equivalence scales and naive price deflators. In particular, S is independent of prices and utility if and only if $a_k^z = 0$ for all k and $a_{kl}^z = 0$ for all k, l. In this case,

$$S_N(z) = z^{a_0^z}. \tag{18}$$

[14] Setting rather than estimating the parameter a_0^R is consistent with previous research using the QAI demand system (e.g., Banks et al., 1997 and Pashardes, 1995). In addition, Banks et al. (1997) note that estimated expenditure share equations do not vary much with the choice of a_0^R.

[15] For an important class of exceptions to this rule, see Donaldson and Pendakur (1999). They show that equivalence scales which depend on utility and are iso-elastic in expenditure are identifiable.

[16] Blackorby and Donaldson require that expenditure share equations are not linear in the log of expenditure for their result. Given QAI, expenditure shares are linear if and only if $q(p) = 0$. This restriction is rejected at conventional levels of significance for all of the QAI models estimated in this paper.

The price deflator D is independent of utility if and only if $b_k = q_k = 0$ for all k, so that $b(p) = 1$ and $q(p) = 0$. In this case,

$$\ln D_N(p) = \sum_{k=1}^{m} a_k^R \ln p_k + \frac{1}{2} \sum_{k=1}^{m} \sum_{l=1}^{m} a_{kl}^R \ln p_k \ln p_l. \tag{19}$$

Table 2 gives selected parameter estimates for three specifications of the QAI demand system which use either naive price deflators or naive equivalence scales[17]. Table 3 gives all parameter estimates for a fourth specification of the QAI

Table 2
Selected parameter estimates: Models with D_N or S_N

Parameter	Model 1: $x/D_N S_N$		Model 2: $x/D_N S_F$		Model 3: $x/D_F S_N$	
	Coef	Std Err	Coef	Std Err	Coef	Std Err
a_0^R	4.540	@	4.540	@	4.540	@
a_0^z	0.441	@	0.455	@	0.441	*0.006*
a_{fh}^R	0.213	*0.001*	0.181	*0.002*	0.203	*0.001*
a_{fr}^R	0.069	*0.001*	0.078	*0.001*	0.074	*0.001*
a_{ra}^R	0.336	*0.001*	0.392	*0.001*	0.328	*0.001*
a_{op}^R	0.077	*0.001*	0.072	*0.001*	0.075	*0.001*
a_{fe}^R	0.048	*0.001*	0.040	*0.001*	0.048	*0.001*
a_{cl}^R	0.102	*0.001*	0.093	*0.001*	0.105	*0.001*
a_{to}^R	0.085	*0.001*	0.072	*0.001*	0.097	*0.001*
a_{pt}^R	0.035	*0.001*	0.039	*0.001*	0.035	*0.001*
b_{fh}					−0.118	*0.002*
b_{fr}					0.061	*0.001*
b_{ra}					−0.116	*0.002*
b_{op}					−0.001	*0.001*
b_{fe}					0.043	*0.001*
b_{cl}					0.062	*0.001*
b_{to}					0.060	*0.002*
b_{pt}					0.007	*0.001*
q_{fh}					0.007	*0.003*
q_{fr}					0.013	*0.002*
q_{ra}					−0.025	*0.004*
q_{op}					0.008	*0.001*
q_{fe}					0.017	*0.002*
q_{cl}					0.016	*0.002*
q_{to}					−0.042	*0.002*
q_{pt}					0.005	*0.001*
Cases	19,526		19,526		19,526	
LLF	240,730		243,910		246,794	
df	44		88		61	

[17] A full set of parameter estimates is available on request from the author.

Table 3
Model 4: $x/D_F S_F$

	Coef	Std Err		Coef	Std Err		Coef	Std Err
a_0^R	4.540	*@*	a_0^z	0.455	*0.020*			
a_{fh}^R	0.163	*0.002*	a_{fh}^z	0.059	*0.003*	b_{fh}	−0.117	*0.002*
a_{fr}^R	0.084	*0.001*	a_{fr}^z	−0.015	*0.002*	b_{fr}	0.062	*0.001*
a_{ra}^R	0.387	*0.001*	a_{ra}^z	−0.089	*0.003*	b_{ra}	−0.118	*0.002*
a_{op}^R	0.070	*0.001*	a_{op}^z	0.008	*0.002*	b_{op}	−0.001	*0.001*
a_{fe}^R	0.041	*0.001*	a_{fe}^z	0.010	*0.002*	b_{fe}	0.043	*0.001*
a_{cl}^R	0.096	*0.001*	a_{cl}^z	0.013	*0.002*	b_{cl}	0.063	*0.001*
a_{to}^R	0.085	*0.001*	a_{to}^z	0.018	*0.002*	b_{to}	0.060	*0.002*
a_{pt}^R	0.040	*0.001*	a_{pt}^z	−0.009	*0.001*	b_{pt}	0.006	*0.001*
$a_{fh,fh}^R$	−0.030	*0.014*	$a_{fh,fh}^z$	0.007	*0.017*	q_{fh}	0.022	*0.002*
$a_{fh,fr}^R$	0.032	*0.010*	$a_{fh,fr}^z$	−0.038	*0.012*	q_{fr}	0.011	*0.002*
$a_{fh,ra}^R$	−0.084	*0.005*	$a_{fh,ra}^z$	0.014	*0.006*	q_{ra}	−0.048	*0.003*
$a_{fh,op}^R$	0.018	*0.011*	$a_{fh,op}^z$	0.002	*0.013*	q_{op}	0.010	*0.001*
$a_{fh,fe}^R$	0.050	*0.010*	$a_{fh,fe}^z$	0.019	*0.012*	q_{fe}	0.019	*0.002*
$a_{fh,cl}^R$	0.015	*0.010*	$a_{fh,cl}^z$	0.004	*0.012*	q_{cl}	0.020	*0.002*
$a_{fh,to}^R$	0.013	*0.006*	$a_{fh,to}^z$	0.016	*0.007*	q_{to}	−0.037	*0.002*
$a_{fh,pt}^R$	−0.028	*0.005*	$a_{fh,pt}^z$	−0.016	*0.006*	q_{pt}	0.003	*0.001*
$a_{fr,fr}^R$	−0.014	*0.009*	$a_{fr,fr}^z$	0.001	*0.011*			
$a_{fr,ra}^R$	0.048	*0.004*	$a_{fr,ra}^z$	−0.010	*0.005*	Cases	19,526	
$a_{fr,op}^R$	0.006	*0.008*	$a_{fr,op}^z$	0.019	*0.009*	LLF	250,076	
$a_{fr,fe}^R$	0.005	*0.009*	$a_{fr,fe}^z$	−0.003	*0.010*	df	105	
$a_{fr,to}^R$	−0.057	*0.008*	$a_{fr,to}^z$	0.003	*0.010*			
$a_{fr,pt}^R$	0.008	*0.004*	$a_{fr,pt}^z$	−0.002	*0.005*			
$a_{fr,pt}^R$	−0.025	*0.004*	$a_{fr,pt}^z$	0.019	*0.005*			
$a_{ra,ra}^R$	0.056	*0.007*	$a_{ra,ra}^z$	0.047	*0.009*			
$a_{ra,op}^R$	−0.001	*0.003*	$a_{ra,op}^z$	0.004	*0.004*			
$a_{ra,fe}^R$	−0.027	*0.004*	$a_{ra,fe}^z$	−0.021	*0.004*			
$a_{ra,cl}^R$	−0.028	*0.004*	$a_{ra,cl}^z$	−0.020	*0.005*			
$a_{ra,to}^R$	−0.002	*0.005*	$a_{ra,to}^z$	0.002	*0.005*			
$a_{ra,pt}^R$	0.044	*0.003*	$a_{ra,pt}^z$	−0.007	*0.003*			
$a_{op,op}^R$	0.026	*0.013*	$a_{op,op}^z$	−0.007	*0.016*			
$a_{op,fe}^R$	−0.034	*0.011*	$a_{op,fe}^z$	−0.005	*0.013*			
$a_{op,cl}^R$	0.006	*0.010*	$a_{op,cl}^z$	−0.012	*0.012*			
$a_{op,to}^R$	−0.017	*0.004*	$a_{op,to}^z$	−0.021	*0.005*			
$a_{op,pt}^R$	0.020	*0.004*	$a_{op,pt}^z$	0.006	*0.005*			
$a_{fe,fe}^R$	−0.032	*0.017*	$a_{fe,fe}^z$	−0.029	*0.020*			
$a_{fe,cl}^R$	0.051	*0.015*	$a_{fe,cl}^z$	0.025	*0.018*			
$a_{fe,to}^R$	−0.017	*0.004*	$a_{fe,to}^z$	0.000	*0.005*			
$a_{fe,pt}^R$	−0.017	*0.004*	$a_{fe,pt}^z$	−0.018	*0.005*			
$a_{cl,cl}^R$	0.035	*0.015*	$a_{cl,cl}^z$	0.024	*0.018*			
$a_{cl,to}^R$	0.004	*0.004*	$a_{cl,to}^z$	−0.007	*0.005*			
$a_{cl,pt}^R$	0.000	*0.004*	$a_{cl,pt}^z$	0.004	*0.005*			
$a_{to,to}^R$	0.006	*0.006*	$a_{to,to}^z$	0.014	*0.007*			
$a_{to,pt}^R$	0.004	*0.003*	$a_{to,pt}^z$	0.009	*0.004*			
$a_{pt,pt}^R$	0.000	*0.003*	$a_{pt,pt}^z$	−0.004	*0.004*			

demand system which uses both a flexible price deflator and a flexible equivalence scale. In the Tables, I use the following shorthand for price effects: *fh* denotes food-at-home; *fr* restaurant food; *ra* rental accomodation; *op* household operation; *fe* household furnishing and equipment; *cl* clothing; *to* private transportation operation; *pt* public transportation. Personal care price effects are all determined by the adding-up restrictions noted above. In the tables, the symbol '@' indicates a parameter that is set a priori rather than estimated.

In cases with a naive price deflator which requires homothetic preferences, equivalence scales cannot be identified from demand behaviour (see Blackorby and Donaldson, 1993). In this case, only the price dependence of the equivalence scale can be identified — the dependence of the equivalence scale on household size z cannot be identified. In homothetic models, the dependence of equivalence scales on household size, a_0^z, will be taken to equal its estimated value from the corresponding non-homothetic model.

Not surprisingly, models (1) and (2) which are restricted to naive price deflators — and therefore homotheticity — fit the data poorly. The likelihood ratio test statistics for $b_k = q_k = 0$ for all k are 6064 with naive equivalence scales and 6166 with flexible equivalence scales, both of which exceed the 1% critical value of 32. Models (1) and (3) which are restricted to naive equivalence scales also fit the data comparatively poorly. The likelihood ratio test statistics for the restrictions $a_k^z = 0$ for all k and $a_{kl}^z = 0$ for all k,l are 3180 with naive price deflators and 3282 with flexible price deflators, both of which exceed the 1% critical value of 68. Thus the flexible model with D_F and S_F fits the data much better than the models with D_N or S_N. The estimates for the flexible model are similar to Pashardes' (1995) estimates of a QAI model using American quarterly household expenditure data[18].

The naive price deflator associated with Model 1 is translog in prices, and its values for the seven survey years and five regions are given in Table 4. Table 5 gives the values of the flexible price deflator for a household with $x = a(p, z)$ — that is, with $u = 0$ — calculated from the Model 4 parameter estimates using Eq. (16). The flexible price deflator depends on expenditure, and its elasticities with respect to total expenditure for a family with $x = a(p, z)$ are given in Table 6. Notably, the values of the flexible price deflator at $x = a(p, z)$ are quite close to the

[18]Pashardes (1995) estimates a price-dependent equivalence scale using a QAI specification for reference preferences, but requires the equivalence scale to be Cobb–Douglas rather than translog in prices. (Unlike the current paper, Pashardes (1995) estimates the equivalence scale but does not assess its impact on inequality measurement).
I estimate a model restricted to give a Cobb–Douglas equivalence scale with $a_{kl}^z = 0$ for all k, l. The base price household size elasticity, a_0^z, is equal to 0.489, which is very close to the model (4) estimate of a_0^z. The likelihood ratio test statistic for the restrictions $a_{kl}^z = 0$ for all k, l is 388, which is larger than the χ_{36}^2 1% critical value of 59. As in the case with a translog equivalence scale, use of a Cobb–Douglas equivalence scale instead of a price-independent equivalence scale makes essentially no difference to patterns in measured inequality.

Table 4
Naive price deflator

	1969	1978	1982	1986	1992	1996	1997
Atlantic	0.352	0.671	0.938	1.115	1.345	1.461	1.489
Quebec	0.346	0.632	0.912	1.095	1.328	1.384	1.404
Ontario	0.375	0.693	1	1.214	1.525	1.637	1.668
Prairies	0.335	0.653	0.952	1.074	1.318	1.380	1.409
BC	0.384	0.739	1.081	1.224	1.547	1.699	1.730

Table 5
Flexible price deflator at $x = a(p, z)$

	1969	1978	1982	1986	1992	1996	1997
Atlantic	0.352	0.670	0.930	1.114	1.346	1.458	1.485
Quebec	0.348	0.627	0.897	1.083	1.321	1.378	1.397
Ontario	0.383	0.701	1	1.219	1.543	1.659	1.692
Prairies	0.336	0.654	0.949	1.068	1.315	1.379	1.406
BC	0.390	0.747	1.087	1.231	1.565	1.718	1.748

Table 6
Elasticity of D_F with respect to x at $x = a(p, z)$

	1969	1978	1982	1986	1992	1996	1997
Atlantic	0.022	−0.023	0.007	−0.002	0.014	0.014	0.016
Quebec	0.052	0.029	0.047	0.024	0.042	0.035	0.035
Ontario	0.001	−0.032	0	−0.002	0.010	0.009	0.012
Prairies	0.025	−0.017	0.009	0.007	0.026	0.021	0.021
BC	0.014	−0.030	0.006	−0.015	0.004	0.011	0.010

values of the naive price deflator. This means that any differences that emerge between using naive and flexible price deflators in inequality measurement have to do with the expenditure-dependence of the price deflator rather than with the overall level of the price deflator.

The naive equivalence scale may be computed from the Model 2 results in Table 2. It is given by $S_N(z) = z^{0.441}$. A flexible equivalence scale can be computed from the estimated parameters for Model 4 using (14). Table 7 shows the

Table 7
Elasticity of S_F with respect to z

Region	1969	1978	1982	1986	1992	1996	1997
Atlantic	0.457	0.451	0.467	0.453	0.454	0.458	0.459
Quebec	0.458	0.470	0.485	0.475	0.468	0.465	0.466
Ontario	0.424	0.432	0.454	0.448	0.440	0.438	0.439
Prairies	0.448	0.448	0.458	0.460	0.460	0.458	0.458
BC	0.433	0.432	0.448	0.442	0.440	0.444	0.444

estimated elasticities of the flexible equivalence scale with respect to family size. There is not much variation in the equivalence scale elasticity due to relative price variation. The main pattern is that the elasticity of expenditure with respect to household size is smaller in regions where the price of housing is higher, that is, in Ontario and British Columbia.

4. Measurement of consumption inequality

There is a growing empirical literature on the evolution of expenditure and consumption inequality in various countries, due in part to increasing awareness that income may be a poor indicator of well-being in the presence of well-functioning credit markets[19]. For example, Slesnick (1998) suggests that family expenditure may be a useful indicator for welfare analysis, and Blundell and Preston (1998) show the precise conditions under which it is an exact ordinal measure of well-being[20]. These authors highlight the importance of intertemporal decisions for families and suggest that since within-period expenditure is a choice variable for households (subject to lifetime expenditure constraints), it may be a useful indicator of well-being. Empirical work on the distribution of family and household expenditure in Australia (Barrett et al., 1999a), Canada (Pendakur, 1998), Portugal (Gouveia and Tavares, 1995), the United Kingdom (Blundell and Preston, 1998), and the United States (Cutler and Katz, 1992) suggests that the distribution of expenditure has evolved differently from the distribution of income. All of these papers use naive price deflators and naive equivalence scales and none of them take account of price differences across regions. I extend the work of Pendakur (1998) to cover the period 1969–1997, and investigate whether or not the use of regional price information, expenditure-dependent price deflators and price-dependent equivalence scales changes our qualitative assessment of the evolution of family expenditure inequality.

I assess inequality in expenditure on a bundle of the same nine commodities used in the demand estimation, but replace rented shelter expenditures with the imputed rental flow from shelter. Family expenditure inequality is measured off a full sample of families who live in cities with greater than 30,000 residents, regardless of their tenure as owners or renters. These observations are weighted at

[19]With a few exceptions (e.g., Gouveia and Tavares, 1995), the literature uses the term 'consumption' rather than 'expenditure', but they refer to the same economic concept: the level of expenditure on goods and services during a period of time, chosen subject to a contraints on saving, borrowing and (lifetime) income.

[20]Blundell and Preston (1998) show that consumption decisions are independent of age only under very restrictive conditions, which suggests that analysis of consumption inequality should condition on age. Blundell and Preston (1998) and Barrett et al. (1999b) evaluate the distribution of consumption conditional on age and find significant differences across cohorts.

the level of the family, so each individual in a family is assigned the family weight.

For rental tenure households, the shelter consumption flow is known, but for owner-occupier households, the shelter consumption flow is not known, because for these households the flow of spending includes an investment component. Since many poor households — especially the elderly — own their accomodation, it is important to account for this. Further, since some rental tenure households live in subsidized or cooperatively owned housing (CMHC, 1997), they may get a larger flow of consumption than their rental expenditures indicate. For both these reasons, I impute the value of shelter for all households.

As noted in Smeeding et al. (1993) and Katz (1983), imputed consumption flows may be based on either the *market value* of the good or the *opportunity cost* of the capital embodied in the good (see Diewert, 1974 or Yates, 1994). In the former case, the researcher assigns the market value of housing, conditional on dwelling characteristics, to the household as its flow of imputed rent. In the latter case, the researcher assigns the opportunity cost, or alternative capital market return, of the capital implicitly invested in housing to the household as its flow of imputed rent. Smeeding et al. (1993) impute consumption flows from owned accomodation based on the opportunity cost of home equity because they do not have data on the the local cost of housing. In this paper, I lack information on home equity in 1997 and have information on local housing costs, so I use the market value approach.

I estimate the market value of accomodation as the average rent for accomodation in the same year and region (35 region-years) in the same city size[21] (0–29,999 or 30,000+) with the same number of rooms (1 to 11+rooms). I then assign the imputed market value of accomodation to each household instead of actual shelter expenditure.

To check for robustness, I also estimate inequality measures using the opportunity cost approach for the period 1969 to 1996. Here, the imputed consumption flow from owner-occupied housing is equal to the Government of Canada one-year real bond yield multiplied by the estimated potential selling price of the house. This imputation assumes that the expected appreciation of owned housing equals the current rate of inflation. For owner-occupier households, the imputed consumption flow is given by the opportunity cost of capital. For renters, it is given by actual rent expenditures.

Table 8 gives the number of observations and the mean nominal family expenditure in each year. Table 9 gives the mean nominal family expenditure for each region in each year. The key feature to note from Table 9 is that Ontario and BC have consistently higher family expenditure levels than the other regions. Examination of the regional price deflators in Tables 4 and 5 suggests that some of

[21] Most of the empirical work focusses on residents of cities with 30,000 or more residents, but Table 11 presents some results for all residents of Canada.

Table 8
Nominal consumption data

Year	Cases	Mean	Std Dev	Min	Max
1969	7759	5227	2387	722	65,861
1978	5532	10,666	4568	1671	42,526
1982	7951	14,951	6636	2787	88,637
1986	7635	18,671	8772	3318	79,597
1992	6649	22,295	9737	4951	88,729
1996	6874	23,846	10,603	4267	112,281
1997	7941	23,986	10,337	4339	101,300

Table 9
Mean nominal consumption by region

Region	1969	1978	1982	1986	1992	1996	1997
Atlantic	4573	9660	13,545	16,402	19,752	20,600	20,786
Quebec	4878	10,357	13,783	17,230	20,057	21,158	21,073
Ontario	5725	10,957	15,511	20,156	24,328	25,886	25,923
Prairies	4910	10,703	15,283	18,217	21,016	22,697	23,362
BC	4936	10,761	15,901	18,889	23,600	25,270	25,436

this inter-regional inequality may be undone by the countervailing effect of higher prices in Ontario and BC relative to the rest of Canada.

5. Results

Table 10 presents estimated Gini coefficients for five functions of family expenditure, x, defined above. The asymptotic standard error for all of these estimated Gini coefficients is 0.0015[22]. As noted above the base price vector is

Table 10
Gini coefficients for family consumption

Model	1969	1978	1982	1986	1992	1996	1997
S_N only	0.179	0.175	0.181	0.192	0.184	0.184	0.180
D_N and S_N	0.175	0.175	0.178	0.190	0.179	0.180	0.177
D_N and S_F	0.176	0.174	0.178	0.190	0.178	0.180	0.176
D_F and S_N	0.172	0.177	0.175	0.189	0.175	0.177	0.173
D_F and S_F	0.172	0.177	0.175	0.189	0.175	0.176	0.173

[22]The estimated asymptotic standard errors are calculated following Barrett and Pendakur (1995), and range from 0.0012 to 0.0017.

that faced by residents of Ontario in 1982[23]. Relative inequality indices (e.g., the Gini, Theil and Atkinson indices) are independent of scalings of the distribution. So, if regional price differences are ignored and all observations within a year are assumed to face the same prices, users of naive price deflators need not even deflate x. In this case, measured inequality is the same whether or not x is deflated by the (naive) price index. Thus, I include estimated Gini coefficients for family expenditure divided by the naive equivalence scale only.

The five measures presented in Table 10 are: x/S_N; $x/D_N S_N$; $x/D_N S_F$; $x/D_F S_N$; and $x/D_F S_F$. I note that the estimated Gini coefficients for x/S_N are quite similar to Pendakur's (1998) estimates for family consumption in Canada over 1978–92, and somewhat similar to Osberg's (1997) estimates for after-tax money income in Canada over 1975–84. I note also that Gini coefficients are quite sluggish: in 1994, the Gini coefficient for disposable family income was 0.39 in the United States and 0.27 in Sweden, so that 12 percentage points separated the most and least equal income distributions in the Luxembourg Income Study (De Nardi et al., 2000).

Several results emerge from Table 10. First, allowing for price dependence in the equivalence scale does not seem to affect the level of or trend in measured inequality. The difference between the estimated Gini coefficients using flexible versus naive equivalence scales never exceeds 0.1 percentage points. This finding is somewhat surprising given previous research showing that measured inequality does respond to the form of the equivalence scale (e.g., Lancaster et al., 1999; Pendakur, 1999; Phipps, 1993; Buhmann et al., 1988)[24]. However, that research assessed how inequality measures change when the dependence of the equivalence scale on family size is changed, rather than when the dependence of the equivalence scale on prices is changed. Here, measuring inequality with price-dependent scales does not change results in comparison to measurement using price-independent scales.

Second, measured inequality responds to whether or not variation in prices across regions is accounted for. Fig. 1 plots three rows of Table 10: x/S_N; $x/D_N S_N$; and $x/D_F S_F$. Consider the difference between undeflated measures and naively deflated measures, using a naive equivalence scale in both cases, shown in Fig. 1 with the dotted and thin lines. Here, we can see that taking regional price differences into account changes the estimated level of inequality, and pushes down the Gini coefficient in each year by as much as 0.5 percentage points. This is

[23] Donaldson (1992) shows that social evaluation on the basis of adjusted expenditure functions (extended money metrics) is not in general independent of the base price vector. I re-estimated all the elements of Table 10 with base prices set to those in Ontario 1969 and Ontario 1997, and found no substantive departures from the main features in Table 10.

[24] A similar finding to the current paper is Idson and Miller (1999). They find that allowing for different price deflators for families with and without children — effectively allowing for a price dependent equivalence scale — does not affect the measured poverty level in the United States over 1968 to 1987.

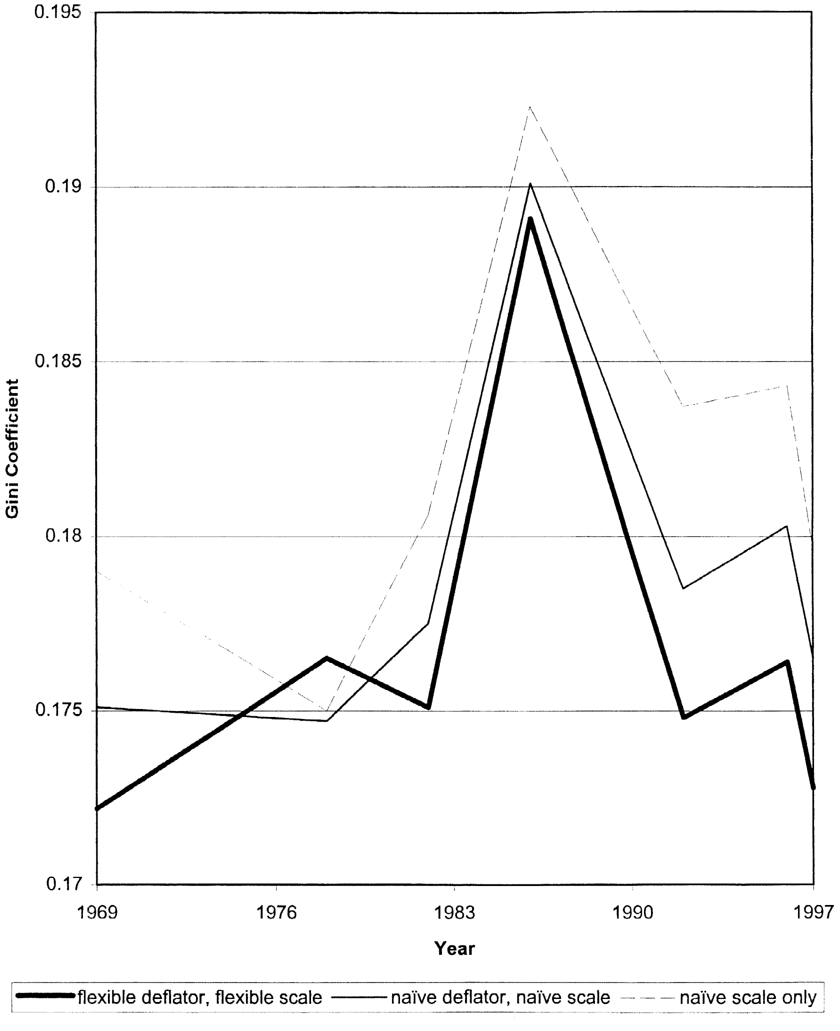

Fig. 1. Family expenditure inequality, urban residents, 1969–1997.

because regions with higher average expenditure also have higher prices, so that some inter-regional expenditure inequality is undone by inter-regional price variation. The difference between measures seems to be larger in the 1990s than in the 1980s. This is due to the greater variation in price deflators across regions in the 1990s versus the 1980s.

Taking regional price differences into account also changes the trend in measured inequality. Between 1969 and 1978, the Gini coefficient using x/S_N declined significantly by 0.4 percentage points, but that for the measure using

$x/D_N S_N$ was unchanged. This difference is due to disequalising price changes across regions during this period. In particular, the naive price deflator rose by 91% and 95%, respectively, in the Atlantic and Prairies, but only by 85% in wealthy Ontario.

Third, measured inequality responds to whether or not an expenditure-dependent price deflator is used. Consider the difference between the Gini coefficient using $x/D_N S_N$ and that using $x/D_F S_F$, shown in Fig. 1 with the thin and thick lines. Recall that there is little difference between measures using $x/D_N S_N$ and measures using $x/D_N S_F$. Thus, it is the flexible price deflator drives the difference between measures using $x/D_N S_N$ and measures using $x/D_F S_F$. Between 1969 and 1978, the Gini coefficient using the naive deflator was unchanged, but that using the flexible deflator and scale increased significantly by 0.5 percentage points. Over the next four years, the naive measure increased (marginally significantly) by 0.3 percentage points, but the flexible measure decreased (insignificantly) by 0.2 percentage points. These differences can be understood by looking at the expenditure dependence of the flexible price deflator (Table 6) during these years. Between 1969 and 1978, the elasticity of the price deflator with respect to total expenditure decreased. The elasticity decreased because the relative price of food at home rose by about one-quarter over this period (see Appendix A). Since the expenditure share of food at home declines strongly with total expenditure ($b_{fh} = -0.117 < 0$), price changes over this period hurt poor families more than rich families, which pushed up the level of inequality.

In contrast, between 1978 and 1982, the elasticity of the price deflator increased due to a large decrease in the relative price of shelter — the price of rented accomodation rose by 20–30% over this period compared to 40–60% for most other commodities. Since the expenditure share of shelter declines with total expenditure ($b_{ra} = -0.118 < 0$), price changes over this period hurt poor families less than rich families, which pushed down the level of inequality.

Between 1982 and 1986, the elasticity of the price deflator declined. This resulted in a small difference between the naive and flexible measures over this period: they increased by 1.2 and 1.4 percentage points, respectively. Between 1986 and 1992, the elasticity of the price deflator increased in all regions. This resulted in another small difference between the naive and flexible measures. They decreased by 1.1 and 1.4 percentage points, respectively.

Between 1992 and 1996, inequality increased slightly for both measures, and between 1996 and 1997, inequality decreased moderately. For both measures, the level of inequality in 1992 is insignificantly different from that in 1996 or that in 1997 — inequality was fairly stable during the 1990s. This finding stands in contrast to the significant decline of 0.4 percentage points between 1992 and 1997 observed for measured inequality using expenditure divided by the naive equivalence scale only. In sum, using an expenditure-dependent price deflator and regional price information reverses trends in family expenditure inequality in the 1970s, early 1980s and possibly even the 1990s.

Table 11
Gini coefficients: Alternate imputation, all residents

	Model	1969	1978	1982	1986	1992	1996	1997
Opp. Cost Imp.	S_N only	0.207	0.206	0.209	0.223	0.215	0.211	N/A
Opp. Cost Imp.	D_N and S_N	0.204	0.208	0.208	0.220	0.212	0.205	N/A
Opp. Cost Imp.	D_F and S_F	0.200	0.210	0.205	0.219	0.207	0.202	N/A
All Residents	S_N only	0.194	0.185	0.187	0.196	0.188	0.190	0.183
All Residents	D_N and S_N	0.190	0.184	0.183	0.193	0.183	0.184	0.178
All Residents	D_F and S_F	0.186	0.186	0.181	0.192	0.179	0.181	0.175

Table 11 presents results using an alternate imputation strategy and using all residents of Canada. Rows labelled 'Opp. Cost Imp.' use the opportunity cost method to impute consumption flows for shelter. This imputation is only available for 1969 to 1996 due to data limitions. Rows labelled 'All Residents' show Gini coefficients computed for the entire population of Canada in each year (using the market value imputation), rather than just families living in cities with 30,000 or more residents.

The results are similar in spirit to those presented in Table 10. Consider the results using the opportunity cost rental imputation method. Here, naive methods show no change in measured inequality between 1969 and 1978, but flexible methods show a significant increase of 1.0 percentage point in the Gini coefficient. Between 1978 and 1982, the Gini coefficient using naive methods were unchanged or increased slightly, but that using flexible methods decreased significantly by 0.5 percentage points. However, although the results using the market value imputation show a qualitative difference between measures during the 1990s, results using the opportunity cost imputation show a significant decrease in inequality regardless of which deflator or scale is used.

Inequality measures computed for all residents of Canada tell a similar story: between 1969 and 1978, the Gini coefficient using naive deflators and scales declined greatly, but that using the flexible deflator and scale was unchanged. Between 1978 and 1982, naive methods show little change in the Gini coefficient, but more flexible methods show a significant decline. Although the level of and trend in measured inequality is affected by choice of rental imputation method (market value or opportunity cost) and choice of population (urban or all), these choices do not seem to change conclusions about whether or not it is important to take prices seriously in the measurement of inequality. The use of regional price variation and expenditure-dependent price deflators makes the distributional change in the 1970s less equalising, and the distributional change over the early 1980s more equalising. It may also make the 1990s appear less equalising. On the other hand, it seems that the use of a price-dependent equivalence scale has essentially no effect on measured inequality, regardless of imputation method or target population.

6. Conclusions

The measurement of inequality is typically characterised by very simple adjustments for price differences across regions and time periods and for differences in the demographic characteristics of families. Adjustment for price differences is typically made through division of expenditure by an expenditure-independent price deflator and adjustment for demographic differences is typically made through division of expenditure by a price-independent equivalence scale. These strategies are associated with very severe restrictions on consumer demand. I use demand estimation to recover expenditure-dependent price deflators and price-dependent equivalence scales, which are then used to estimate Gini coefficients for Canadian expenditure (consumption) inequality over 1969 to 1997. The use of more flexible price-dependent equivalence scales does not seem to affect the level of or trend in measured inequality. However, the use of regional price information and expenditure-dependent price deflators affects both the level of and year-to-year changes in family expenditure inequality.

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Appendix A. Price data

Region	Year	Pfh	Pfr	Pra	Pop	Pfe	Pcl	Pto	Ppt	Ppc
Atlantic	1969	31.3	29.5	36.2	32.2	50.5	47.7	29.1	27.9	40.7
	1978	71.9	71.3	70.4	59.8	70.2	71.0	50.7	50.9	65.1
	1982	101.3	99.8	86.2	100.1	104.8	96.8	87.3	91.1	93.8
	1986	115.8	122.0	109.5	111.9	121.4	111.8	94.7	125.6	106.3
	1992	137.0	164.7	129.0	128.6	139.8	142.1	119.4	156.9	131.8
	1996	153.0	175.1	135.0	144.7	140.9	149.2	139.2	212.2	145.0
	1997	155.8	174.4	136.1	146.0	138.3	156.2	145.3	237.7	149.3
Quebec	1969	29.1	30.0	33.6	31.6	47.2	47.1	36.9	38.3	41.3
	1978	65.2	70.3	57.5	60.1	72.7	71.8	65.8	51.6	67.2
	1982	100.5	103.9	72.8	96.1	97.2	98.6	119.7	101.1	94.4
	1986	122.1	126.3	93.7	109.6	109.9	109.4	127.5	127.3	111.9
	1992	136.4	173.1	115.5	128.0	135.6	146.9	143.0	170.4	145.9
	1996	143.4	181.6	121.2	137.7	134.9	144.2	144.7	204.1	152.7
	1997	146.2	185.4	122.1	140.9	136.0	146.4	146.3	216.0	154.1

Ontario	1969	29.6	32.0	45.7	30.5	45.9	45.9	29.1	35.9	39.8
	1978	65.2	67.9	82.8	56.6	67.5	70.0	54.6	54.0	66.5
	1982	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
	1986	119.5	126.6	125.2	111.0	116.6	115.4	123.7	135.4	115.1
	1992	139.5	177.3	160.5	135.0	135.4	147.3	159.6	179.1	147.2
	1996	147.0	186.3	173.5	146.0	136.1	155.1	176.9	222.8	145.3
	1997	147.1	188.2	176.1	149.0	136.2	156.6	188.1	241.3	149.3
Prairies	1969	29.2	31.5	35.1	27.5	45.6	45.2	26.7	29.5	37.8
	1978	69.1	74.0	68.1	55.7	68.9	73.1	45.7	54.2	65.9
	1982	100.3	107.2	90.8	86.2	98.5	106.4	84.6	101.3	95.2
	1986	117.2	127.1	99.8	99.5	113.1	118.6	89.7	123.5	109.9
	1992	137.5	167.2	120.5	120.0	133.2	151.4	121.0	169.3	134.8
	1996	145.0	179.6	126.8	127.7	130.7	154.9	124.4	193.5	142.5
	1997	149.4	184.4	128.6	129.3	132.3	157.0	127.0	211.6	145.2
BC	1969	30.9	36.5	44.9	32.2	49.6	48.1	30.8	28.5	43.8
	1978	71.8	83.3	85.6	63.6	71.3	75.3	51.7	58.5	70.6
	1982	105.5	125.7	110.6	99.0	102.2	107.4	108.5	102.1	102.8
	1986	123.3	137.7	129.0	112.8	116.8	120.3	99.5	133.8	116.3
	1992	146.0	185.2	163.7	131.9	138.3	146.9	156.1	177.4	142.0
	1996	159.5	204.3	176.4	140.5	136.3	152.3	209.7	206.2	145.0
	1997	163.8	208.4	178.4	143.6	136.5	157.1	208.1	225.4	147.4

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