AN EXPLORATION OF AUSTRALIAN PETROL DEMAND: UNOBSERVABLE HABITS, IRREVERSIBILITY, AND SOME UPDATED ESTIMATES

Robert Breunig* and Carol Gisz

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* Research School of Social Sciences, Australian National University

ABSTRACT

We estimate a demand equation for petrol in Australia. We explore a methodological improvement to the standard dynamic demand model–a more general model which allows for slowly evolving, unobservable habits. If this habit formation model with unobserved stocks is correct, then standard estimation techniques produce inconsistent estimates. We find a short-run price elasticity of -0.1 to -0.14 and a long-run price elasticity of -0.2 to -0.3. Importantly, we find that standard techniques are misleading about the precision of elasticity estimates and that the confidence interval around the long-run price elasticity is quite wide, with a 90% confidence interval of -0.02 to -0.38. Results are very sensitive to the inclusion of time trends, which appear to be appropriate. We test for price irreversibility and find, in contrast to the U.S., almost no evidence that petrol responds differently to price increases and decreases.

JEL Classification Numbers: C2, Q41

Keywords: Single-equation demand analysis; lagged adjustment; habit formation; price irreversibility; petrol elasticities

1 Introduction

In this paper we estimate a single-equation, dynamic petrol demand model for Australia. We use quarterly data from 1966 to 2006. We provide updated estimates of short- and long-run price and income elasticities after dealing with several methodological issues, one of which, the inclusion of slowly-evolving, unobservable habits, has been largely ignored in the literature. Elasticity estimates are important in a wide range of settings. Our updated estimates will be of use to policy makers and others thinking about macroeconomic stability, petrol price instability, the effect of excise taxes on petrol demand, and responses to administratively imposed changes on petrol prices such as carbon taxes.

Single-equation, dynamic demand estimation remains a commonly employed empirical tool in the determination of price and income elasticities using aggregate data. We specifically address two methodological issues: the modeling of unobservable habits and potential asymmetry in the response of demand to price increases and decreases.

Our results have important implications for the estimation of petrol demand using aggregate data. Accounting for unobservable, slowly-changing habits introduces moving average terms into the error structure of the model. This means that standard approaches which use ordinary least squares will produce inconsistent estimates. Maximum likelihood, readily available in any modern software package, will provide consistent estimates.

Our results also have important empirical implications. Since our approach is primarily focused on determinants of long-run adjustment, we find little effect of our methodological changes on short-run price elasticities. We find point estimates for the short-run price elasticity of petrol of around -.13, in line with other studies. For the long-run price elasticity, we find point estimates ranging between -.2 and -.3. While this is roughly similar to other studies, we find much larger standard errors once we account for unobservable habits. A model which explicitly incorporates uncertainty should give estimates which are more uncertain, so this is not surprising. However, it does imply that the confidence intervals for the long-run price elasticity from standard estimation approaches are spuriously narrow. We need to accept that our estimates of the long-run price elasticity are more uncertain than we thought.

Finally, of empirical interest, we find almost no evidence that demand responsiveness to price differs for increases and decreases. This is in contrast with the U.S. literature which does find such asymmetry. In what follows, we first present the basic model and then continue with a presentation of our modeling innovations, followed by a brief description of our data. Our results are presented in section 4. In section 5 we compare our results to other studies and conclude.

2 Modeling approaches

Many aggregate studies of petrol adopt a single-equation approach of dynamic demand estimation. The single-equation approach is simple to implement and requires only a limited amount of data, but does not allow for the imposition of cross-equation restrictions and realistic modeling of the inter-relationship between consumption of different goods.[1] Despite these shortcomings, we focus on the single-equation approach as it continues to be the workhorse of applied studies of petrol demand.

Single-equation studies have a common ancestor in the flow adjustment models of Houthakker and Taylor (1966) and Balestra and Nerlove (1966). Demand for capital goods and consumer durables are often modeled using a partial adjustment mechanism where the change in the stock of a durable good is proportional to the gap between current desired and past actual stock level. A speed of adjustment parameter determines how quickly the actual stock moves toward the desired level.

A common formulation of this model (e.g. Baltagi and Griffin (1997)) for desired gasoline consumption per person, based upon the idea of partial adjustment driven by a fixed (in the short-term) vehicle stock is

$$v_t^* \equiv \left(\frac{V^*}{N}\right)_t = \kappa_1 \left(\frac{P_g}{P}\right)_t^{\kappa_2} \left(\frac{Y}{N}\right)_t^{\kappa_3} \left(\frac{Car}{N}\right)_{t-1}^{\kappa_4} \epsilon_t \tag{1}$$

where V^* is the desired volume of gasoline, P_g is the price of gasoline, P is the overall price level, Y is real income, N is population, and Car is the stock of cars.[2]

To capture the fact that adjustment may not be instantaneous, the following relationship is posited

$$v_t = \left(v_t^*\right)^{\theta} \left(v_{t-1}\right)^{1-\theta} \tag{2}$$

v_t is the actual consumption of gasoline per person at time t. If $\theta = 1$, then $v_t^* = v_t$, or adjustment of actual to desired consumption happens instantly. This is unlikely for the case of petrol as many factors which determine demand such as commuting habits and physical dispersion of housing are fixed in the short term.

Re-arrangement and combining these two equations leads to the following model

[1] Intrilligator et al. (1996, Chapter 7) contains an excellent discussion of the trade-offs of these alternative approaches.

[2] A more complicated version of this model could incorporate features of the current stock of cars, such as age, size, and fuel efficiency as well as information on vehicle utilization rates, all of which relate to the amount of petrol consumed. Such an engineering type approach was pioneered by Sweeney (1979).

for observed gasoline consumption per person

$$
\begin{aligned}
v_t &= \kappa_1^{\theta} \left(\frac{P_g}{P}\right)_t^{\kappa_2\theta} \left(\frac{Y}{N}\right)_t^{\kappa_3\theta} \left(\frac{Car}{N}\right)_{t-1}^{\kappa_4\theta} v_{t-1}^{1-\theta} \epsilon_t^{\theta} \\
&= \kappa_1^{\theta} \left(p_t\right)^{\kappa_2\theta} \left(y_t\right)^{\kappa_3\theta} \left(car_{t-1}\right)^{\kappa_4\theta} v_{t-1}^{1-\theta} \epsilon_t^{\theta} \qquad (3)
\end{aligned}
$$

where the last line defines the notation for real and per-person variables.

Taking logs provides the equation estimated in much of the literature

$$
\begin{aligned}
ln(v_t) &= \theta ln(\kappa_1) + \kappa_2 \theta ln\left(p_t\right) + \kappa_3 \theta ln\left(y_t\right) + \kappa_4 \theta ln\left(car_{t-1}\right) \\
&\qquad + \left(1-\theta\right) ln(v_{t-1}) + \theta ln\left(\epsilon_t\right) \\
&= \beta_0 + \beta_p ln\left(p_t\right) + \beta_y ln\left(y_t\right) + \beta_v ln(v_{t-1}) + \beta_c ln\left(car_{t-1}\right) + \varepsilon_t \qquad (4)
\end{aligned}
$$

Studies which use (4) include Baltagi and Griffin (1983), Drollas (1984), Baltagi and Griffin (1997) and some of the studies reviewed in Graham and Glaister (2002). Other studies, such as Gately (1991), Dargay (1992), Jones (1993), Dargay and Gately (1997) and most Australian studies such as Donnelly (1982) have estimated models without information about the stock of cars. This provides a model for estimation which is a restricted version (with β_c (κ_4) set to zero)[3] of equation (4) above

$$
ln(v_t) = \beta_0 + \beta_p ln\left(p_t\right) + \beta_y ln\left(y_t\right) + \beta_v ln(v_{t-1}) + \varepsilon_t \qquad (5)
$$

We will follow the approach of estimating this model without vehicle stocks.[4] To account for the fact that adjustment of actual to desired consumption may take more than one period, many papers extend the model by including additional lagged values of v_t. The lagged values to include are usually chosen by standard time series criteria– removing residual correlation or maximizing some fit criteria. This provides a model for estimation as

$$
ln(v_t) = \beta_0 + \beta_p ln\left(p_t\right) + \beta_y ln\left(y_t\right) + \sum_{j=1}^{q} \beta_{vq} ln(v_{t-q}) + \varepsilon_t \qquad (6)
$$

We first examine a methodological issue relating to this basic model. We criticize 2 and 5 for their inability to adequately capture the role of unobserved factors (physical or psychological) which may lock people into their current consumption habits. We propose an alternative formulation of the model. Our proposed preferred approach implies that moving average terms should be included in the model. This implies that standard estimation by ordinary least squares is inconsistent. This proposal has important implications for any single-equation, dynamic demand estimation problem.

[3] One could alternately view this restriction as assuming that vehicle stocks are constant.

[4] Data on the stock of automobiles in Australia is not available. New motor vehicle registrations and new motor vehicle sales data are available, but neither series covers our whole sample period. We do not know of any Australia-wide source for vehicle usage data broken down by vehicle efficiency.

Secondly, we look at whether the demand response to price increases is the same as the demand response to price decreases. Equation (6) imposes that these are equal (reversibility). Price irreversible models allow the response to price increases to differ from the response to price decreases. This is less novel, but price irreversible models have not previously been estimated using Australian data.

The next two sub-sections are devoted to these issues.

2.1 Controlling for unobserved stock effects

Papers estimating equation (6) typically claim that the lagged dependent variable provides a mechanism "by which the unobservable composition of the vehicle stock and utilization rates can be included in the demand function."[5] It is not clear to us that this model successfully achieves this objective. This formulation omits vehicle stock and utilization rates from the determinants of desired petrol consumption, equation (1), and the hope is that these things have been put back into the model for actual petrol consumption through the ad-hoc adjustment process of equation (2). Although models with this basic structure are commonly estimated with similar justification (see, for example, Jones (1993), Gately (1991) and Dargay (1992)), these models fail to capture fixed costs of past decisions on current petrol consumption.

Consider an alternative model for petrol consumption at time t which depends upon relative prices and real income, as above, and also depends upon an unobservable stock variable, s^*_{t-1} which captures simultaneously physical (efficiency of current vehicle stock, physical dispersion of commuters) and psychological (automobile utilization rates, habits) aspects of fuel consumption

$$ln(v_t) = \alpha_1 + \alpha_2 ln\,(p_t) + \alpha_3 ln\,(y_t) + \alpha_4 s^*_{t-1} + w_t \tag{7}$$

Even in the presence of data on vehicle stocks, one might want to augment equation (1) with an unobserved stock variable, as the raw number of automobiles would fail to capture efficiency differences across types of cars and heterogeneous vehicle usage.

Habits, the vehicle stock, and city layouts evolve according to

$$\Delta s^*_t \equiv s^*_t - s^*_{t-1} = ln(v_t) - \delta s^*_{t-1} \tag{8}$$

If this were a standard capital equipment model, δ would be the depreciation rate. Here it captures how quickly the determinants of fuel consumption change. Were $\delta = 1$, adjustment would be instantaneous and hence the current stock of vehicles and habits have no impact on next period's fuel consumption. In this respect, $\delta = 1$ is akin to $\theta = 1$ in equation (2) above. Note further that if $\delta = 1$, equation (7) collapses to

[5] Donnelly (1982, page 320)

equation (5). Estimating equation (5) is thus imposing an assumption that there are no 'stock effects' (either physical or psychological) related to past consumption which affect current consumption. For petrol consumption in particular, as noted above, this assumption is far too strong.

The approach we prefer, of including an unobservable stock variable, following Houthakker and Taylor (1966), looks quite similar to the model of equations (1) and (2) above, however, it produces a very different estimating equation. Using equation (7) to write $\Delta ln(v_t)$, substituting equation (8) into that expression, and using equation (7) lagged one period to replace s^*_{t-2} in the expression for $\Delta ln(v_t)$ provides, after some re-arrangement,

$$\begin{aligned} ln(v_t) = &\alpha_1\delta + \alpha_2\delta ln\left(p_{t-1}\right) + \alpha_2\Delta ln\left(p_t\right) + \alpha_3\delta ln\left(y_{t-1}\right) + \alpha_3\Delta ln\left(y_t\right) \\ &+ \left(1 + \alpha_4 - \delta\right) ln(v_{t-1}) + w_t - \left(1 - \delta\right) w_{t-1} \end{aligned} \tag{9}$$

or,

$$\begin{aligned} ln(v_t) = &\gamma_0 + \gamma_p ln\left(p_{t-1}\right) + \gamma_{\Delta p}\Delta ln\left(p_t\right) + \gamma_y ln\left(y_{t-1}\right) + \gamma_{\Delta y}\Delta ln\left(y_t\right) \\ &+ \gamma_v ln(v_{t-1}) + u_t \end{aligned} \tag{10}$$

Importantly, the error term in (10) is a moving average. The presence of the lagged dependent variable implies that this equation can not be estimated consistently by ordinary least squares,[6] yet the vast majority of applied studies use least squares with no mention of this issue.

Stocks may take more than one period to adjust in which case we would replace equation (8) with

$$s^*_t = ln(v_t) + \left(1 - \delta_1\right) s^*_{t-1} + \sum_{j=2}^{q} \delta_j s^*_{t-j} \tag{11}$$

Combining equations (7) and (11), we would estimate

$$\begin{aligned} ln(v_t) = &\gamma_0 + \sum_{j=1}^{q} \gamma_{pj} ln\left(p_{t-j}\right) + \gamma_{\Delta p}\Delta ln\left(p_t\right) + \sum_{j=1}^{q} \gamma_{yj} ln\left(y_{t-j}\right) \\ &+ \gamma_{\Delta y}\Delta ln\left(y_t\right) + \sum_{j=1}^{q} \gamma_{vj} ln(v_{t-j}) + u_t \end{aligned} \tag{12}$$

u_t is now a qth-order moving average. Maximum likelihood estimation will be the most efficient approach to estimation.

An alternative approach to single equation estimation building upon these simple models, inspired by the modern time series literature, has been to eschew a theoretical model and directly specify an auto-regressive, distributed lag version of equation (5)

[6] The one exception to this statement would be if w_t followed an AR(1) process $w_t = \left(1 - \delta\right) w_{t-1} + \nu_t$ where ν_t were i.i.d.

where multiple lags of each variable are included, and to employ general to specific testing to reduce the autoregressive distributed lag model to some parsimonious representation with white noise residuals (see Jones (1993)). Other papers, such as Hunt and Ninomiya (2003) have pursued a co-integration approach.

In the absence of strong theoretical views of the adjustment process, it might be preferable to begin such model building with the more general model of equation (12) rather than its restricted counterpart, equation (6). Equation (12) is a more flexible specification in that the parameters which determine the long- and short- run elasticities are estimated separately, unlike the parameters of (6) which require that the long-run elasticity be a scaled up (or down) value of the short-run elasticity. It is straightforward to test whether the restrictions of equation (6) hold after estimating equation (12).

2.2 Allowing for price irreversibility

The models of (6) and (12) both assume that price rises and drops have symmetric effects on consumption. This reversibility assumption has come under some attack (see Gately (1991), Dargay (1992), and Dargay (2004) for example) with the justification that consumers react quite differently to prices rises and decreases. Furthermore, petrol prices over the last 40 years appear to be characterized by long periods of stable nominal (and gently declining real) prices and short periods of large price increases. It seems plausible that these might provoke asymmetric reactions.

Several models exist which allow for asymmetric effects of price increases and decreases. Such models were introduced by Wolffram (1971) and Traill et al. (1978). We follow the formulation of Dargay (1992) where price decreases and increases are allowed to have different effects on consumption. Furthermore, we allow any portion of a price increase above the historical maximum (to that point in time) to have a separate effect. This can be achieved by replacing price in the above equations with three constructed price series

$$\begin{aligned}
ln(p)_t^+ &= \sum_{s=2}^{t} \left[(ln(p)_s - ln(p)_{s-1}) - \left(ln(p)_s^{max} - ln(p)_{s-1}^{max} \right) \right] 1\left(ln(p)_s > ln(p)_{s-1} \right) \\
ln(p)_t^- &= \sum_{s=2}^{t} \left[ln(p)_s - ln(p)_{s-1} \right] 1\left(ln(p)_s < ln(p)_{s-1} \right) \\
ln(p)_t^{max} &= \begin{array}{l} ln(p)_1 \text{ if } t = 1 \\ ln(p)_t \text{ if } ln(p)_t > ln(p)_s \text{ for all } s < t \\ ln(p)_{t-1} \text{ if } ln(p)_t = ln(p)_{t-1} \end{array}
\end{aligned} \tag{13}$$

where $1(\cdot)$ is an indicator function equal to one if the expression in brackets is true, zero otherwise. $ln(p)_t^+$ is the cumulative series of sub-maximum price increases, $ln(p)_t^-$

is the cumulative series of price decreases, and $ln(p)_t^{max}$ is the series of maximum historical prices. It's easy to show that

$$\begin{aligned} ln(p_t) &= ln(p)_t^{max} + ln(p)_t^+ + ln(p)_t^- \qquad \text{and} \\ \Delta ln(p_t) &= \Delta ln(p)_t^{max} + \Delta ln(p)_t^+ + \Delta ln(p)_t^- \end{aligned} \tag{14}$$

and we can plug these into equations (5) and (10) to generate price irreversible versions of those models

$$\begin{aligned} ln(v_t) =&\beta_0 + \beta_{pmax} ln(p)_t^{max} + \beta_{p+} ln(p)_t^+ + \beta_{p-} ln(p)_t^- \\ &+ \beta_y ln\,(y_t) + \beta_v ln(v_{t-1}) + \varepsilon_t \end{aligned} \tag{15}$$

and

$$\begin{aligned} ln(v_t) =&\gamma_0 + \gamma_{pmax} ln\,(p)_{t-1}^{max} + \gamma_{p+} ln(p)_{t-1}^+ + \gamma_{p-} ln(p)_{t-1}^- \\ &+ \gamma_{\Delta pmax} \Delta ln(p)_t^{max} + \gamma_{\Delta p+} \Delta ln(p)_t^+ + \gamma_{\Delta p-} \Delta ln(p)_t^- \\ &+ \gamma_y ln\,(y_{t-1}) + \gamma_{\Delta y} \Delta ln\,(y_t) + \gamma_v ln(v_{t-1}) + u_t \end{aligned} \tag{16}$$

The extension to models with q lags is straightforward. Before we present our estimates of the base models–equations (5) and (10)–and their extensions, we will briefly discuss the data.

3 Data

All of the data we use are publicly available and the analysis data are available on Robert Breunig's web page.[7] In the appendix, we provide details on data sources and detailed univariate ARIMA analysis of each variable. Our sample estimation period is September, 1966 to September, 2006 and we use quarterly data. Reliable petrol price data are not available prior to September, 1966.

Litres of volume of petrol consumed per person is shown in figure 1 for the period September, 1966 to September, 2006. The series shows steady growth prior to 1979 and then is slowly declining since 1979. This may be evidence of some kind of transition to a steady-state level of petrol consumption or may be evidence of a permanent change in the properties of the series due to the oil price shocks of the 1970s. In general, our modeling approach will be to seek a consistent model which describes the data over the entire sample period. We will look at both the consequences of modeling apparent breaks in the data and of simply attempting to ignore those breaks. If we assume that

[7] `http://econrsss.anu.edu.au/Staff/breunig/contact_bb.htm`

there is no break in the data, our preferred univariate model for the log volume data is an AR(8). The pre-1979 and post-1978 data are individually best described using a deterministic time trend. If we assume that a break occurred in 1979, our preferred univariate model is an AR(8) with separate time trends before and after 1979.

The switch to smaller, more fuel efficient vehicles in Australia occurred later than it did in Europe, only after 1978 when domestic petrol prices were fully exposed to world oil prices.[8] The first oil price shocks of 1973-1974 had only a minor impact as import parity pricing of oil was introduced in Australia in 1978. Thus the 'observed' break coincides with what one might expect based upon this history.

Our population and price level data are from the Australian Bureau of Statistics (ABS). Using population over age 15, instead of total population as we do, makes no difference to our results. We use a consumer price index (CPI) which includes petrol price. One may argue that it would be better to use a price index net of petrol, but the weighting of petrol price in the overall CPI is never more than six per cent during our sample period. Removing the effect of petrol price from the CPI has almost no consequences for the movements of CPI.

We use petrol price data from the ABS going back to 1972. For the September, 1966 to 1972 period, we use data published by Donnelly (1981) which he obtained from the ABS. The ABS no longer publishes petrol data prior to 1972. Donnelly (1981) provides petrol data going back to 1960, but in his paper he indicates that the ABS data he uses began in September, 1966 and the data prior to that date is generated by a back-casting model that he created. We thus begin our sample period with the first available observation from the ABS in September, 1966. Real petrol price is graphed in figure 1 alongside the volume series. Its log appears to follow a pure unit root.

For real, per-person income we use household gross disposable income divided by population and the price level, as measured by CPI. This measure has the advantage over wage measures such as average weekly earnings in that it includes other sources of household income such as investment income. Considering the entire series, we find that it is best described by a unit root with drift. Income growth, as has been well-documented, decreases from the mid-1970s. If we consider the data only in the post-1974 period, we also find a unit root with drift. We do not have confidence in separate unit root tests over the 1966-1974 period due to the small number of observations. Our preferred univariate model for income growth is an AR(2) with a dummy intercept shift in the post-1974 period.

[8] We are grateful to an anonymous referee for pointing this out and for suggesting a test of a break in the relationship between price and volume–see section 4 below.

4 Estimation results

We estimate a range of models to compare and contrast the modeling approaches described in section 2. Given our results from the univariate analysis of petrol volume, we estimate each of these models with and without time trends. For the time trend models, we allow for separate time trends before the first quarter of 1979 and after the last quarter of 1978. We also estimate separate models for the pre-1979 and post-1978 periods.[9]

Our general approach will follow this outline: we estimate models with multiple lags of volume (or volume, price, and income) and with or without moving average terms consistent with the specific modeling approach under consideration. Given the univariate results, we explore the consequences of adding time trends to these models. We will select the 'best' models (with and without time trends) using standard selection criteria (such as Akaiki Information Criteria (AIC)) and requiring that the model produce white noise residuals. We will then test irreversibility by introducing the appropriate terms into both the base model and our 'best' model.

4.1 Models for full sample period: 1966 to 2006

In Table 1 we present a summary of the price and income elasticities from the model results for the full sample period from 1966 to 2006. Complete results are in appendix tables A1 and A3. The standard approach would be to estimate the base model with lags and keep significant lags. This would produce the estimates of column (a) in Table 1. As can be seen from the table, the inclusion of separate time trends (column (b)) for the pre-1979 and post-1978 periods–which are strongly statistically significant–have a large impact on the elasticity estimates. When we include time trends, the long-term price elasticity falls from unity to -0.25. The long-run income elasticity on the other hand goes from being insignificant to being about 0.34 and significantly different than zero.

The models of columns (c) and (d) introduce price-irreversibility into the models of columns (a) and (b). For (c), we are unable to reject that the effect of new maximums is zero and we impose that restriction. For column (d), we can not reject that the effect of new maximums, price increases, and price decreases are all identical. For the model without time trends, we prefer the irreversible model of column (c) to the reversible model of column (a). Once we introduce the time trends, however, we find that we prefer the reversible model of column (b) to the irreversible one of column (d) on standard model selection criteria.

[9] We only present the full sample and post-1978 estimates in this paper. Please contact the authors for the estimates for the 1966 to 1978 period.

The models of columns (e) and (f) incorporate unobservable stocks into the standard model. The model with time trends (column (f)) is preferred to the one without (column (e)). While the time trends are individually insignificant, they are jointly significant. We reject irreversibility in both of these models. Importantly, note that the standard errors of the long-run price elasticity in the models with unobserved stocks are twice those of the models without unobserved stocks. We can reject the specification of column (b) in favor of column (f) and likewise for column (a) with respect to column (e). This leads us to conclude that the standard approach is providing spurious precision in the estimates of the long-run price elasticity.

Table 1: Elasticity estimates for Australian petrol demand
Full sample period 1966q3 to 2006q3

	(a)	(b)	(c)	(d)	(e)	(f)
Equation number in text	(6)	(6)	(15)	(15)	(12)	(12)
Estimation Technique	OLS	OLS	OLS	OLS	MLE	MLE
Includes time trends	No	Yes	No	Yes	No	Yes
Lags of v_t	1,2,4	4	1,2,4	4	1,3,4	1,3,4
Short-run price elasticity	$-.082^{**}$ (0.015)	$-.13^{**}$ (0.021)			$-.12^{**}$ (0.027)	$-.13^{**}$ (0.030)
Long-run price elasticity	-1.01^{**} (0.37)	-0.25^{**} (0.051)			-1.15^{*} (0.69)	-0.20^{*} (0.11)
Short-run price elasticity (to new maximums)				$-.077^{*}$ (0.043)		
Short-run price elasticity (to increases)			$-.14^{**}$ (0.023)	$-.14^{**}$ (0.030)		
Short-run price elasticity (to decreases)			$-.096^{**}$ (0.020)	$-.14^{**}$ (0.035)		
Long-run price elasticity (to new maximums)				-0.27^{**} (0.074)		
Long-run price elasticity (to increases)			-0.78^{**} (0.15)	-0.27^{**} (0.069)		
Long-run price elasticity (to decreases)			-0.55^{**} (0.14)	-0.15^{*} (0.086)		
Short-run income elasticity	0.01 (0.02)	0.18^{**} (0.062)	0.16^{**} (0.055)	0.20^{**} (0.076)	.12 (0.085)	$.16^{**}$ (0.093)
Long-run income elasticity	0.12 (0.22)	0.34^{**} (0.13)	0.94^{**} (0.22)	0.39^{**} (0.15)	$-.33$ (0.54)	.27 (0.28)

Appendix Tables A1 and A3 contain the full model estimates.
** and * indicate significance at the .05 and .10 levels, respectively.

Since the six models of Table 1 span the break in the volume data in the late 1970's (described in section 3 above) and since this break is related to a change in petrol pricing, it may not be reasonable to impose a constant relationship between price and volume across the entire sample period. We re-estimated the six models of Table 1, allowing the coefficients on all current and lagged price variables in each model to differ in the pre-1979 and post-1978 periods. In five of the six models, we fail to reject that there is a break in the short-run or long-run price elasticities[10]. The one exception was the model of column (b) where we did find a break (significant at the six per cent level) in the price elasticities across the full sample period. If we estimate the model with the break, we find a large decrease in the price elasticity after 1978. The short-run elasticity pre-1979 is -0.28 while after 1978, it falls to -0.12. The long-run elasticity pre-1979 is -0.57 and it decreases, after 1978, to -0.23. Below, we consider the late sample period in isolation and, for the same specification, find an identical short-run price elasticity for that period and a slightly higher long-run elasticity.

4.2 Models for the post-1978 period

To further explore the impact of the break in the late 1970s, we estimate all models on the data post-1978. In Table 2 we present a summary of the price and income elasticities from these results. Complete results are in appendix tables A2 and A4.

Table 2: Elasticity estimates for Australian petrol demand
Late sample period 1979q1 to 2006q3

	(a)	(b)	(c)	(d)
Equation number in text	(6)	(6)	(12)	(12)
Estimation Technique	OLS	OLS	MLE	MLE
Includes time trends	No	Yes	No	Yes
Lags of v_t	1,4	1,4	1,3,4	1,3,4
Short-run price elasticity	$-.091^{**}$ (0.023)	$-.12^{**}$ (0.024)	$-.11^{**}$ (0.041)	$-.14^{**}$ (0.045)
Long-run price elasticity	-0.28^{**} (0.11)	-0.30^{**} (0.093)	-0.15 (0.15)	-0.23 (0.19)
Short-run income elasticity	-0.055^{*} (0.032)	0.21^{**} (0.099)	0.22 (0.13)	0.28^{**} (0.14)
Long-run income elasticity	-0.17^{**} (0.078)	0.53^{**} (0.029)	-0.26^{**} (0.11)	0.61 (0.76)

Appendix Tables A2 and A4 contain the full model estimates.

Columns (a) and (b) present models without unobservable stocks while (c) and (d) are the results from the specification which includes unobservable stocks. We present both models without and with time trends. For the standard model without a time trend, income perversely has a significant, negative influence on consumption. This is due to the fact that consumption is trending downward while income is trending upward. Unsurprisingly, this effect disappears when the time trend is added. We test for, and reject, irreversibility in the models of columns (b) through (d). We prefer the

[10] Evidence for a break was very weak–the p-values on these tests ranged from .5 to .9. Full details of these estimates and tests are available from the authors.

irreversible version of column (a), but given the importance of time trends in general, we reject this specification as inappropriate. (See appendix Table A2, column (b), for the details of the irreversible model.)

4.3 Methodological implications: some conclusions

4.3.1 Models with and without unobservable stocks

In both the estimates on the full sample and those using only the post-1978 data we find that we can reject the standard specification in favor of the model with unobservable stocks. The moving average terms are jointly significantly different than zero in all cases. We can also reject that the short- and long-run elasticities are scaled versions of each other in all of the unobservable stock models that we estimate. These conclusions are robust to the exclusion or inclusion of time trends. We feel this provides a strong argument for the adoption of the model of equation (12) over that of equation (6). This is a major departure from the established literature.

The main consequence of the adoption of the unobservable stock model is a large decrease in the precision of the estimates of the long-run elasticity. We discuss this more in detail in section 5 below. There is not much effect on short-run elasticities.

4.3.2 The importance of time trends

The inclusion of time trends in the models appears to be justified from the univariate analysis of the volume data. The time trends are statistically significant and improve model fit, providing another justification for their inclusion. When we estimate the models over the entire sample period, the inclusion of the time trends has a large impact on the results. Failing to include time trends produces very large own-price, long-run elasticities (around -1). Including the time trend produces long-run elasticities in the range of -.20 to -.25.

In the post-1978 sample, the time trend is also statistically significant, but price elasticities appear to be fairly insensitive to its omission. However, excluding the time trend produces negative and statistically significant income elasticities which hardly seems intuitive. Including the time trends results in positive (although not always statistically significant) short- and long-run income elasticities.

One could view the post-1978 time trend as picking up technological change. In this period, there seems to be little effect of income on demand. It could be that income may lead to increasing demand for travel and/or for more expensive cars, but this need not equate to more petrol consumption, particularly if more expensive cars are using more fuel-efficient technology. The model, while hardly proof of this kind of relationship, is consistent with this type of story. As previously noted, the positive pre-1979 time trend is consistent with a transition to a steady-state level of petrol consumption.

Overall, the time trends seem important both in producing sensible elasticity estimates and in model fit.

4.3.3 Price irreversibility

The only models for which we find any evidence of price irreversibility are the models without the unobservable stock variable and without time trends. As argued above, we think that both of these additions to the standard specification are justified. Therefore, we find the evidence for price irreversibility very weak. This is in contrast to research on the U.S. but similar to results from the U.K. (see Dargay and Gately (1997) and Dargay (2004). In our data, price irreversibility only arises when the model is misspecified.

4.4 How should the effect of income be modeled?

The literature on dynamic demand estimation is split when it comes to dealing with non-stationarity. One strand of the literature (e.g. Baltagi and Griffin (1997) or Dargay (2004)) completely ignores the stationarity properties of the data. Another strand models demand and income as being co-integrated (e.g. Samimi (1995) or Hunt and Ninomiya (2003)). We have, to this point, adopted the first approach. The second approach is clearly inappropriate given our univariate analysis. Volume is trend stationary (with a break) and decreasing on average after 1979 whereas income is integrated of order one. It is not theoretically possible for two such series to be co-integrated.

One might further argue that it is not possible for income, being non-stationary and drifting upwards, to have any long-run effect on a stationary series (volume, in this case). Also, particularly in the models with lagged income and moving average terms, the presence of the non-stationary variables on the right-hand side of the model may result in inconsistent estimates. For both of these reasons, we think that it is important to estimate models which impose no long-run relationship between income and volume in our preferred models with unobservable stocks. (Note that in the models without unobservable stocks that there is no way to impose this relationship without also imposing no short-run relationship between income and volume. Furthermore, these models do not include lagged values of the non-stationary variable.)

We provide details of this model (for the full and late sample periods, with and without time trends) in appendix Table A5. Our overall conclusions are not much affected. Dropping income increases the precision with which we estimate the time trends in those models that include them. We find insignificant short-run income elasticities in all models. The short-run price elasticities that we estimate are all in the range -0.10 to -0.12, and not significantly different from those summarized in Tables 1 and 2. In the models without time trends, we find very large long-run price elasticities as before. In the preferred specifications with time trends, we find long-run price elasticities of about -0.20 which are, as above, very imprecisely estimated.

5 Discussion and conclusions

The primary contribution of this paper is to provide updated estimates of the petrol demand function for Australia while at the same time exploring the role of unobservable habits, the lagged adjustment of desired to actual consumption, and the possibility of asymmetric demand reactions to price increases and decreases.

The inclusion of time trends appears to be important and, given the univariate analysis, would appear to be justified. Dimitropoulos et al. (2005), in a recent paper, also stress the importance of including time trends in energy demand equations. We thus focus our discussion on the models which include time trends. Empirically, we find short-run price elasticities of -.10 to -.14 irrespective of the model which we estimate. Long-run elasticities range from -.2 to -.3 but are not significantly different from -.25 in any of the models we estimate. Elasticity estimates are surprisingly robust to the choice of model. If we restrict the estimation sample to the post-1978 period, we find slightly larger (more negative) long-run price elasticities, but the differences are very small (generally around .03) and not significantly different than zero. This is consistent with the evidence presented in Espey (1998) that long-run elasticities have tended to increase over time. Short-run income elasticities across the models we estimate range from .16 to .28 while long-run income elasticities are generally about twice as large, ranging from .27 to .61. In our preferred models with unobservable stocks and time trends, the long-run income elasticity is never significantly different than zero. If we impose no long-run effect of income on volume, an implication of our univariate data analysis, our conclusions about price elasticities do not change.

Our 'preferred' model for the full sample period would be that of column (f) of Table 1, our estimate of equation (12) with time trends. When compared to the 'standard' model (column (b) of Table 1), we find identical short-run price elasticities of -.13 and very similar point estimates for long-run price elasticities, -.20 and -.25. However, the inclusion of unobservable 'stocks' (physical and psychological) causes the 90% confidence interval to go from $(-0.17, -0.33)$ to $(-0.02, -0.38)$. If we consider the post-1978 period only, the 90% confidence interval increases from $(-0.15, -0.45)$

in the standard model to $(0.08, -0.54)$ in our preferred model of column (d) of Table 2. In both cases, the width of the confidence interval doubles. One very important conclusion of our paper is that the uncertainty about the 'true' long-run price elasticity is probably much greater than people think. We suspect that researchers may be attracted to the more precise estimates of the standard model, but would suggest that these estimates are spuriously precise.

There has been a small amount of recent work in Australia on petrol demand. Our short-run price elasticity is similar to Donnelly (1982), who used a sample period from 1958 to 1981. Our long-run estimates are less elastic. Hensher and Young (1991) use Australian Bureau of Agriculture and Resource Economics (ABARE) data with a sample period from 1976 to 1988 and obtain a long run price elasticity for Australia of -0.25. They also use a panel data set which they gathered between 1981 and 1985 from a sample of households in Sydney. Using an engineering approach, they find a much higher long-run elasticity of -.66. Samimi (1995), using a co-integration approach, finds a very small long-run price elasticity of -0.12. His focus is on the road transport sector whereas our results are driven more by personal consumption. Given our univariate unit root test results, a co-integration approach would be inappropriate for this data.[11]

The recently released Australian Government Green Paper on the Carbon Pollution Reduction Scheme, Australian Government (2008), cite a short-run price elasticity, based on estimates from the Bureau of Infrastructure, Transport and Regional Economics, of -0.15, which is very similar to what we find. The Green Paper cites a long-run elasticity of -0.4, which falls within the 95% confidence interval for our preferred model.

Internationally, Graham and Glaister (2002) survey a range of papers which report short-run price elasticities between -.2 and -.3 and long-run elasticities between -.2 and -.8. Our results indicate that Australia falls on the low end of the scale. Espey (1998) reports that studies from Australia, New Zealand, and Canada are on the low end of the scale for estimates of long-run elasticities.[12] The similarity between Australia and Canada is not universally found. Interestingly, Sterner and Dahl (1992) report much higher elasticity estimates for Canada. In fact, looking at their survey of OECD countries, it is difficult to discern much pattern across types of countries. In a recent paper, Hughes et al. (2006), using U.S. data, find a dramatic drop in short-run elasticities of petrol in recent time periods. Our data does not show a similar effect for Australia.

Lastly, we examine asymmetry in demand responses to price increases and de-

[11] If we restrict our sample to the time period used by Samimi (1995), we also reject co-integration in our data.

[12] Her Australian estimates come from Donnelly (1982) and Donnelly (1981).

creases. In contrast with the series of papers by Gately (1991), Dargay (1992), and Dargay and Gately (1997) using U.S. data, we find little evidence of price irreversibility. For one model only, that with time trends and no unobservable stocks, we find some evidence that the short-run response to price decreases is significantly smaller than the short-run response to price increases. This is similar to Dargay (2004) who finds some, but only small, evidence of irreversibility in U.K. data. Irreversibility may be primarily a U.S. phenomena.

While our results provide support for point estimates found in other studies, the approach of including unobservable, slowly evolving unobserved stocks/habits produces estimates which are much less precise. While it may be natural for researchers to be attracted to the more precise estimates, the model which produces those estimates is rejected in favor of the unobservable stocks model. Point estimates are important, but standard errors and confidence intervals are of equal importance. We conclude that the correct ninety per cent confidence interval for the long-run elasticity of petrol is $(-0.02, -0.38)$, which is twice as wide as other studies have found. Useful answers to empirical questions must, of necessity, include information about their precision. In this case, that information has been misleading in previous studies.

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Appendix One: Detailed Model Estimation Results

Table A1: Estimates from basic model without stocks
Reversible and irreversible models with and without time trends
Full sample period 1966q3 to 2006q3

	(a)	(b)	(c)	(d)	(e)
	Models without time trends			**Models with time trends**	
Equation number in text	(6)	(15)	Alternative (15)	(6)	(15)
Table (column) in text	1(a)	n/a	1(c)	1(b)	1(d)
Coefficient					
β_0	0.42^{**} (0.11)	0.38^{**} (0.14)	0.38^{**} (0.13)	2.08^{**} (0.38)	1.99^{**} (0.40)
Post-1978 dummy variable				0.18^{**} (0.032)	0.17^{**} (0.034)
Pre-1979 time trend				0.0023^{**} (0.00075)	0.0021^{**} (0.0011)
Post-1978 time trend				-0.00074^{**} (0.00018)	-0.00081 (0.00089)
β_p	-0.082^{**} (0.015)			-0.13^{**} (0.021)	
β_{pmax}		-0.0062 (0.030)			-0.077^{*} (0.043)
β_{p+}		-0.13^{**} (0.024)	-0.14^{**} (0.023)		-0.14^{**} (0.030)
β_{p-}		-0.096^{**} (0.021)	-0.096^{**} (0.020)		-0.14^{**} (0.035)
β_y	0.0097 (0.02)	0.16^{**} (0.055)	0.16^{**} (0.050)	0.18^{**} (0.062)	0.20^{**} (0.076)
β_{v1}	0.17^{**} (0.06)	0.13^{**} (0.060)	0.13^{**} (0.060)		
β_{v2}	0.12^{**} (0.061)	0.088 (0.061)	0.087 (0.060)		
β_{v4}	0.63^{**} (0.058)	0.61^{**} (0.057)	0.61^{**} (0.056)	0.49^{**} (0.066)	0.49^{**} (0.066)
Long-run income elasticity	0.12 (0.22)	0.94^{**} (0.22)	0.95^{*} (0.21)	0.34^{**} (0.13)	0.39^{**} (0.15)
Long-run price elasticity	-1.01^{**} (0.37)			-0.25^{**} (0.051)	
Long-run price elasticity (to new maximums)		-0.037 (0.18)			-0.15^{*} (0.086)
Long-run price elasticity (to increases)		-0.79^{**} (0.16)	-0.78^{**} (0.15)		-0.27^{**} (0.069)
Long-run price elasticity (to decreases)		-0.57^{**} (0.16)	-0.55^{**} (0.14)		-0.27^{**} (0.074)
AIC	-747.5	-754.1	-756.0	-766.9	-764.8

All models estimated by OLS. The lag length of columns (a), (b), and (c) are chosen on the basis of AIC and residuals which appeared to be close to white noise.
We fail to reject (p-value 0.40) that β_{pmax}, β_{p+}, and β_{p-} are the same in column (e).
Column (c) would be preferred to columns (a) or (b) while column (d) is preferred to column (e) on standard model selection criteria.
In sum, we would reject the irreversible model in favor of the standard reversible model once we incorporate time trends in the model.

Table A2: Estimates from basic model without stocks
Reversible and irreversible models with and without time trends
Late sample period 1979q1 to 2006q3

	(a)	(b)	(c)	(d)
	Models without time trends		**Models with time trends**	
Equation number in text	(6)	(15)	(6)	(15)
Table (column) in text	2(a)	n/a	2(b)	n/a
Coefficient				
β_0	1.99^{**} (0.66)	1.67^{**} (0.65)	1.45^{**} (0.67)	1.06 (0.72)
Time trend			-0.00076^{**} (0.00027)	-0.0014 (0.0011)
β_p	-0.091^{**} (0.023)		$-.12^{**}$ (0.024)	
β_{pmax}		-0.12^{*} (0.026)		-0.057 (0.047)
β_{p+}		-0.12^{*} (0.026)		-0.13^{**} (0.034)
β_{p-}		-0.091^{**} (0.023)		-0.15^{**} (0.043)
β_y	-0.055^{*} (0.032)	0.12 (0.078)	0.21^{**} (0.099)	0.31^{**} (0.13)
β_{v1}	0.10 (0.076)	0.062 (0.076)	0.056 (0.076)	0.054 (0.076)
β_{v4}	0.57^{**} (0.077)	0.56^{**} (0.076)	0.56^{**} (0.075)	0.56^{**} (0.075)
Long-run income elasticity	-0.17^{**} (0.078)	0.33 (0.23)	0.53^{**} (0.029)	0.82^{**} (0.40)
Long-run price elasticity	$-.28^{**}$ (0.11)		$-.30^{**}$ (0.093)	
Long-run price elasticity (to new maximums)				$-.15$ (0.12)
Long-run price elasticity (to increases)		$-.33^{**}$ (0.10)		-0.34^{**} (0.12)
Long-run price elasticity (to decreases)		$-.24^{**}$ (0.085)		-0.39^{**} (0.14)
AIC	-521.8	-526.2	-527.7	-526.2

All models estimated by ordinary least squares.
Column (b) incorporates the restriction $\beta_{pmax} = \beta_{p+}$.
In column (d), we fail to reject the restriction that the model is reversible (with p-value of 0.32). Like in Table A1, we would reject the irreversible model in favor of the standard reversible model once we incorporate time trends in the model.

Table A3: Estimates from model with unobservable stocks
Reversible models with and without time trends
Full sample period 1966q3 to 2006q3

	(a)	(b)	(c)	(d)
Equation number in text	(12)	(12)	(12)	(12)
Table (column) in text	1(e)	1(f)	n/a	n/a
Coefficient				
γ_0	0.37** (0.13)	1.58* (0.82)	1.95** (0.96)	1.34** (0.66)
Post-1978 dummy variable		0.13 * (0.072)	0.16 ** (0.083)	0.11 ** (0.056)
Pre-1979 time trend		0.0020 (0.0015)	0.0026 (0.0017)	0.0018 (0.0013)
Post-1978 time trend		−0.00050 (0.00033)	−0.00060** (0.00027)	−0.00045 (0.00029)
$\gamma_{\Delta p}$	− 0.12 * (0.027)	− 0.13 * (0.030)	− 0.17 ** (0.028)	− 0.14 * (0.031)
γ_{p1}	− 0.16 ** (0.032)	− 0.16 ** (0.041)	− 0.17 ** (0.033)	− 0.20 ** (0.040)
γ_{p2}			−0.041** (0.017)	
γ_{p3}	− 0.12 ** (0.042)	− 0.12 ** (0.043)		
γ_{p4}	0.22 ** (0.033)	0.21 ** (0.037)	0.14 ** (0.029)	0.15 ** (0.034)
$\gamma_{\Delta y}$	0.12 (0.085)	0.16 * (0.093)	0.19 ** (0.098)	0.22 ** (0.089)
γ_{y1}	0.23 ** (0.083)	0.32** (0.10)	0.31** (0.11)	0.34 ** (0.091)
γ_{y2}			0.030 (0.078)	
γ_{y3}	−0.020 (0.089)	−0.025 (0.087)		
γ_{y4}	− 0.23 ** (0.086)	−0.20* (0.11)	−0.22** (0.10)	− 0.26 ** (0.097)
γ_{v1}	0.0042 (0.042)	−0.079 (0.075)	−0.0053 (0.019)	− 0.11 (0.077)
γ_{v2}			− 0.24 ** (0.094)	
γ_{v3}	−0.0060 (0.040)	−0.081 (0.078)		
γ_{v4}	0.95 ** (0.017)	0.79 ** (0.085)	0.78 ** (0.092)	0.80 ** (0.073)
θ_1	0.31* (0.17)	0.36 (0.25)	0.053 (0.16)	0.26** (0.11)
θ_2			0.59 (10.3)	
θ_3	0.24 (0.15)	0.26 (0.23)		
θ_4	− 0.47 ** (0.086)	−0.38** (0.17)	−0.41** (0.11)	−0.37** (0.12)
Long-run income elasticity	−0.33 (0.54)	0.27 (0.28)	0.27 (0.23)	0.28 (0.32)
Long-run price elasticity	−1.15* (0.69)	−0.20* (0.11)	− 0.17 ** (0.076)	− 0.17 * (0.088)
AIC	-791.8	-798.1	-797.0	-781.3

Notes to Table A3 appear after Table A5.

Table A4: Estimates from model with unobservable stocks
Reversible and irreversible models with and without time trends
Late sample period 1979q1 to 2006q3

Equation number in text	$\frac{(a)}{(12)}$	$\frac{(b)}{(12)}$	$\frac{(c)}{(12)}$
Table (column) in text	2(c)	2(d)	n/a
Coefficient			
γ_0	1.99^{**} (1.06)	0.99 (1.19)	-1.31^{**} (0.50)
Time trend		-0.00067 (0.00048)	-0.00081^{**} (0.00026)
$\gamma_{\Delta p}$	-0.11^{*} (0.041)	-0.14^{**} (0.045)	-0.19^{*} (0.031)
γ_{p1}	-0.12^{**} (0.046)	-0.16^{**} (0.054)	-0.20^{**} (0.038)
γ_{p3}	-0.13^{**} (0.051)	-0.13^{**} (0.050)	
γ_{p4}	0.20^{**} (0.047)	0.22^{**} (0.049)	0.18^{**} (0.030)
$\gamma_{\Delta y}$	0.22 (0.13)	0.28^{*} (0.14)	0.39^{**} (0.14)
γ_{y1}	0.33^{**} (0.15)	0.46^{**} (0.19)	0.57^{**} (0.14)
γ_{y3}	-0.093 (0.12)	-0.036 (0.13)	
γ_{y4}	-0.31^{**} (0.14)	-0.24 (0.16)	-0.26^{**} (0.13)
γ_{v1}	-0.054 (0.078)	-0.049 (0.075)	0.031 (0.036)
γ_{v3}	-0.080 (0.080)	-0.078 (0.076)	
γ_{v4}	0.83^{**} (0.087)	0.83^{**} (0.083)	1.01^{**} (0.030)
θ_1	0.42 (0.36)	0.41 (0.35)	0.22 (0.14)
θ_3	0.24 (0.23)	0.21 (0.23)	
θ_4	-0.35 (0.18)	-0.39^{**} (0.27)	-0.99^{**} (0.13)
Long-run income elasticity	-0.26^{**} (0.12)	0.61 (0.76)	0.41 (0.75)
Long-run price elasticity	-0.15 (0.15)	-0.23 (0.19)	-7.43 (9.58)
AIC	-553.1	-554.5	-550.0

All models estimated by maximum likelihood.
If we test whether the third lag can be eliminated from (a), we reject the restriction at the 10 per cent level (p-value is .08). If we impose this restriction, the estimated long-run elasticities do not change.
If we test whether the third lag can be eliminated from (b), we reject the restriction at the 10 per cent level (p-value is .09). If we impose this restriction, we get the model of column (c) which does not have white noise residuals.
For (a) we fail to reject reversibility in all long and short-run coefficients (p-value of .34).
For (b) we fail to reject reversibility in all long and short-run coefficients (p-value of .56).

Table A5: Estimates from model with unobservable stocks
Reversible models with and without time trends
All of these models impose zero relationship between non-stationary income and trend stationary volume

Equation number in text	(a) 1966q3 to 2006q3 (12)	(b) 1966q3 to 2006q3 (12)	(c) 1979q1 to 2006q3 (12)	(d) 1979q1 to 2006q3 (12)
Coefficient				
γ_0	0.51^{**} (0.13)	2.1^{*} (1.0)	1.07 (0.88)	2.24^{**} (1.12)
Post-1978 dummy variable		0.15^{*} (0.085)		
Pre-1979 time trend		0.0028^{*} (0.0016)		
Post-1978 time trend		-0.00024^{**} (0.00012)		-0.00024^{**} (0.00011)
$\gamma_{\Delta p}$	-0.12^{*} (0.029)	-0.12^{*} (0.032)	-0.11^{**} (0.039)	-0.10^{*} (0.042)
γ_{p1}	-0.16^{**} (0.035)	-0.15^{**} (0.043)	-0.13^{**} (0.045)	-0.13^{**} (0.048)
γ_{p3}	-0.13^{**} (0.039)	-0.14^{**} (0.040)	-0.16^{**} (0.051)	-0.16^{**} (0.051)
γ_{p4}	0.21^{**} (0.036)	0.20^{**} (0.040)	0.22^{**} (0.048)	0.21^{**} (0.048)
$\gamma_{\Delta y}$	-0.073 (0.047)	-0.070 (0.045)	-0.033 (0.095)	0.042 (0.086)
γ_{v1}	-0.0016 (0.047)	-0.082 (0.082)	-0.013 (0.063)	-0.068 (0.083)
γ_{v3}	-0.021 (0.043)	-0.098 (0.088)	-0.065 (0.079)	-0.11 (0.10)
γ_{v4}	0.93^{**} (0.015)	0.78^{**} (0.098)	0.88^{**} (0.081)	0.78^{**} (0.10)
θ_1	0.32^{*} (0.18)	0.39 (0.27)	0.43^{**} (0.26)	0.46 (0.35)
θ_3	0.25 (0.16)	0.30 (0.24)	0.28 (0.20)	0.29 (0.25)
θ_4	-0.45^{**} (0.083)	-0.32^{**} (0.16)	-0.32^{**} (0.15)	-0.26^{*} (0.15)
Long-run price elasticity	-0.80^{**} (0.33)	-0.21^{*} (0.11)	-0.36 (0.31)	-0.19 (0.13)
AIC	-787.2	-789.7	-545.9	-549.5

All models estimated by maximum likelihood.
Long-run income elasticity is restricted to equal zero in all of these models.

Notes to Table A3

In column (a), if we set $\gamma_{p3} = \gamma_{y3} = \gamma_{v3} = \theta_3 = 0$ the long-term elasticities are roughly the same. However, we reject this restriction on the model (p-value is 0.01).
The p-value on the test of joint significance of the time variables in column (b) is .19, however, a likelihood ratio test comparing (a) to (b) leads us to reject the restrictions of column (a). The p-value is .006.
The p-value on the test of joint significance of the time variables in column (c) is .05. The restrictions of column (d) are rejected in favor of column (b) or column (c). Columns (b) and (c) are both preferred to a model with all four lags included. They are both acceptable models using standard selection criteria.
When we add irreversibility of the type described in equation (16) to the models of table 11, we do not find convincing evidence of irreversibility. For all four columns we fail to reject reversibility in the long-run and short-run coefficients at the 10 per cent level or higher.

Appendix Two: Data sources and univariate analysis

Table A6: Data definitions and sources

Variable	Symbol	Explanation
Volume of petrol	V	Department of Industry, Tourism and Resources Australian Petroleum Statistics Publication Measured in thousands of litres 1958q3 to 2006q3
Population	N	Unpublished ABS data Measured in thousands of persons 1966q2 to 2006q3
Price of petrol	P_g	1966q3 to 1972q3 from Donnelly (1981) 1972q4 to 2006q3 from ABS (cat. no. 6401.0) Price index 1989/90=100
Overall price level	P	CPI: All groups, weighted average of eight capital cities; ABS (cat. no. 6401.0) Price Index 1989/90 = 100
Income	Y	Seasonally adjusted, total household gross disposable income Measured in millions of dollars from ABS (cat. no. 5206.0) 1959q3 to 2006q3 Divided by CPI: All groups to give real income

Details of univariate data analysis

Volume

We use the log of the per-person volume measure, $ln(v) = ln\left(\frac{V}{N}\right)$. Volume in levels is shown in figure 1 for the sample period 1966q3 to 2006q3. From the graph, we can rule out a time trend or drift in the full series. We conduct the augmented Dickey-Fuller test using a regression with eight lags (the number necessary to generate white noise residuals).

$$\Delta ln(v_t) = \alpha_0 + \rho ln(v_{t-1}) + \alpha_1 \Delta ln(v_{t-1}) + \ldots + \alpha_8 \Delta ln(v_{t-8}) + \epsilon_t \tag{A1}$$

The p-value for the test that $\rho = 0$ is .0049 and we clearly reject a unit root in this data. This result is not sensitive to the number of lags included–if we use only 4 lags, we get a p-value of .0009.

The series looks quite different before 1979 and after 1979. This may be evidence of some type of structural change in the underlying determinants of the consumption of petrol. We analyze the properties of this series before and after this break.

Pre-1979 volume data

We use the 50 observations from 1966q3 to 1978q4. The steady increase in the series must be driven by either a stochastic or a deterministic trend and we estimate the following model to test these competing hypotheses

$$\Delta ln(v_t) = \alpha_0 + \rho ln(v_{t-1}) + \gamma_1 t + \epsilon_t \tag{A2}$$

We then test

H_0	Test value	5% critical value
$\rho = 0$	-6.69	-3.504
$\rho = \gamma_1 = 0$	22.99	6.73
$\gamma_1 = 0$	6.47	2.79

Although the sample is small, the evidence for a time trend, and against a unit root, is strong. We reject the joint hypothesis of unit root and no time trend, we reject the unit root hypothesis, but we find a significant time trend. Thus we conclude that the pre-1979 data are characterized by a deterministic trend.

Our preferred model for the per-1979 period is a simple time trend

$$\widehat{ln(v_t)} = \underset{(0.0056)}{5.14} + \underset{(0.00018)}{0.0085}\, t \tag{A3}$$

The numbers in parentheses are standard errors. The coefficient on $ln(v_{t-1})$, if added to the above equation, is insignificant.

Post-1978 volume data

Use of petrol is gently declining over the post-1978 period. We estimate a model with eight lags to eliminate auto-correlation in the residuals

$$\Delta ln(v_t) = \alpha_0 + \rho ln(v_{t-1}) + \alpha_1 \Delta ln(v_{t-1}) + \ldots + \alpha_8 \Delta ln(v_{t-8}) + \gamma t + \epsilon_t \tag{A4}$$

and test for unit root and time trend as above.

H_0	Test value	5% critical value
$\rho = 0$	-3.546	-3.504
$\rho = \gamma_1 = 0$	7.04	6.49
$\gamma_1 = 0$	-3.14	2.79

This combination of test results would seem to indicate that there is no unit root in the post-1978 sample, but that the data does have a (negative) deterministic trend.

Our preferred model for the 1979q1 to 2006q3 period is an AR(8) with a time trend

$$\begin{aligned}\widehat{ln(v_t)} = {} & \underset{(0.014)}{5.57} - \underset{(0.00012)}{0.00057}t + \underset{(0.078)}{0.25}\, ln(v_{t-1}) + \underset{(0.12)}{0.13}\, ln(v_{t-2}) + \underset{(0.13)}{0.11}\, ln(v_{t-3}) \\ & + \underset{(0.10)}{0.45}\, ln(v_{t-4}) - \underset{(0.10)}{0.22}\, ln(v_{t-5}) - \underset{(0.12)}{0.22}\, ln(v_{t-6}) \\ & - \underset{(0.12)}{0.27}\, ln(v_{t-7}) + \underset{(0.09)}{0.21}\, ln(v_{t-8})\end{aligned} \tag{A5}$$

The second and third lags are not significant, but the model performs better with their inclusion.

Volume data for full sample period

Our preferred model for the full sample period is an AR(8) with a dummy for the post-1978 period and separate time trends for the pre-1979 and post-1978 periods. This model produces white noise residuals and is preferred to any other auto-regressive model in terms of the Akaike (AIC) and Bayes Information Criteria (BIC).

$$\begin{aligned}\widehat{ln(v_t)} = &\underset{(0.014)}{5.14} + \underset{(0.02)}{0.44}\, D_{post1978} + \underset{(0.00043)}{0.0082}\, t * D_{pre1979} - \underset{(0.00012)}{0.00065} t * D_{post1978} \\ &+ \underset{(0.069)}{0.16}\, ln(v_{t-1}) + \underset{(0.09)}{0.13}\, ln(v_{t-2}) + \underset{(0.09)}{0.12}\, ln(v_{t-3}) + \underset{(0.08)}{0.43}\, ln(v_{t-4}) \\ &- \underset{(0.09)}{0.14}\, ln(v_{t-5}) - \underset{(0.09)}{0.18}\, ln(v_{t-6}) - \underset{(0.09)}{0.20}\, ln(v_{t-7}) + \underset{(0.07)}{0.15}\, ln(v_{t-8}) \end{aligned} \tag{A6}$$

There is some quarterly seasonality in the data and we experimented with more parsimonious models which directly incorporate this seasonality. Of these, the best was a seasonal auto-regression model with a non-seasonal AR(2) and a multiplicative seasonal auto-regressive term at the 4th lag. This model, using the lag operator L, may be expressed as

$$\left(1 - \rho_1 L - \rho_2 L^2\right)\left(1 - \rho_{4,1} L\right)\left(\widetilde{ln(v_t)}\right) = \varepsilon_t \tag{A7}$$

where

$$\widetilde{ln(v_t)} = \left(ln(v_t) - \alpha_0 - \beta_1 D_{post1978} - \beta_2 t * D_{pre1979} - \beta_3 t * D_{post1978}\right)$$

Our estimates from this model were significant with values $\rho_1 = 0.19$, $\rho_2 = 0.15$, and $\rho_{4,1} = 0.54$, but unlike the model of equation (A6), the residuals were not equivalent to white noise. Furthermore, the less parsimonious AR(8) of equation (A6) was preferred to the model of (A7) by the AIC and BIC.

Price

We create a real price for petrol by dividing the petrol price index, P_g by the CPI, P, to get $p = \frac{P_g}{P}$.

To test for a stochastic trend we estimate

$$\Delta ln(p_t) = \alpha_0 + \rho ln(p_{t-1}) + \gamma_1 t + \epsilon_t \tag{A8}$$

We then test

H_0	Test value	5% critical value
$\rho = 0$	-2.301	-3.442
$\rho = \gamma_1 = 0$	3.03	6.43
$\gamma_1 = 0$	2.12	2.79

As we fail to reject a unit root or the hypothesis of unit root and no time trend and we reject the time trend, this series would clearly appear to be a unit root. The drift term is not significant, and the unit root tests conducted under the assumption that the drift is zero (or suppressing the drift term) also provide strong evidence of a unit root. Our preferred univariate model is simply

$$\Delta ln(p_t) = \epsilon_t \tag{A9}$$

and, in fact, $\Delta ln(p_t)$ is indistinguishable from white noise.

Real per-person income

We transform our measure of gross domestic household income into real income per-person as $y = \frac{1000*Y}{P*N}$.

The rate at which the series increases appears to change in the mid-1970s. If we are willing to assume that real per-person income follows a unit root with a break in the drift parameter and we conduct a QLR-type break test (see Stock and Watson (2003)) on the series $\Delta ln(y_t)$, we find a break point at 1974q2. Alternatively, we can use the approach of Perron (1989) to test whether the series is best represented by a unit root with a break in the drift parameter or two time trends with different slopes before and after 1974q2.

Ignoring the apparent break and testing the entire series, we begin by estimating

$$\Delta ln(y_t) = \alpha_0 + \rho ln(y_{t-1}) + \sum_{k=1}^{4} \alpha_k \Delta ln(y_{t-k}) + \gamma_1 t + \epsilon_t \tag{A10}$$

We then test

H_0	Test value	5% critical value
$\rho = 0$	-3.411	-3.443
$\rho = \gamma_1 = 0$	6.95	6.43
$\gamma_1 = 0$	2.83	2.79

The second through fourth lags are insignificant, however, so we repeat the test for the model of (A10) with only one lag of $\Delta ln(y_t)$, we get

H_0	Test value	5% critical value
$\rho = 0$	-3.068	-3.442
$\rho = \gamma_1 = 0$	5.52	6.43
$\gamma_1 = 0$	2.56	2.79

While the model with four lags produces test results that are fairly borderline between a stochastic and a deterministic time trend, the model with one lag points more clearly towards a stochastic trend model. This is an interesting example of how adding too many lags to the augmented Dickey-Fuller test can affect the results.

The above approach naively ignores what looks like a break in the mid-1970s. If we consider only the post 1974q2 period and we perform the unit root tests, we have

H_0	Test value	5% critical value
$\rho = 0$	-3.05	-3.446
$\rho = \gamma_1 = 0$	5.43	6.49
$\gamma_1 = 0$	3.13	2.79

The model with one lag is preferred by standard selection criteria to those with more lags and additional lags were insignificant. We fail to reject the joint hypothesis of unit root and no time trend.

Our preferred models are either

$$\Delta y_t = \underset{(0.0032)}{0.013} - \underset{(0.076)}{0.299} \Delta y_{t-1} - \underset{(0.0035)}{0.0098} D_{post1974} \quad \text{or}$$

$$\Delta y_t = \underset{(0.0034)}{0.0144} - \underset{(0.080)}{0.341} \Delta y_{t-1} - \underset{(0.079)}{0.143} \Delta y_{t-2} - \underset{(0.0036)}{0.0107} D_{post1974} \tag{A11}$$

The second lag is significant at the 10% level, but not at the 5% level. The AR(1) is preferred by the AIC and BIC, but the AR(2) produces residuals which look closer to white noise.

Figure 1: Volume consumed and price of petrol. Australia: 1966 to 2006