The valuation of labelling attributes in a wine market

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Paper to be presented at the the 2002 AAEA-WAEA Annual Meeting
Long Beach, California, July 29

This draft: 15 may 2002

Abstract

The values which market participants place on labelling information in the British wine retail market are investigated using a hedonic framework. The results suggest a near asymmetric evaluation of labelling attributes between wines from the 'New World' (Australia) and wines from the 'Old World' (France). The benefits of studying the valuation of attribute information within the hedonic framework are demonstrated by considering the revenue impact of shifts in attributes at the retail level.

Key words: labelling, wine, product quality, hedonic price analysis

JEL classification: L150, D12, C21

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1 Introduction

Consider a heterogeneous good, such as wine, for which the analyst could attempt to reconstruct a consumer's hierarchy of attitudes towards product attributes from the consumer's stated behavioural intentions. Instead of relying on stated intentions as multi-attribute attitude models and conjoint analyses do, we could employ revealed preference analyses, which obtain predictions by combining observations of realised choices with assumptions about underlying decision processes (McFadden (1974); Rosen (1974)). Using the example of observed consumer choices of heterogeneous bundles of labelling attributes in a retail market for wine, this paper employs hedonic price analysis to explore the implicit valuation that market participants make of those components of heterogeneous attribute bundles.

Frederick Waugh (1928) relied on observed consumer choices for asparagus to pioneer the development of hedonic price analysis in agricultural economics. His analysis of vegetable prices is based on the hypothesis that quality of vegetables is related to measurable specification variables. Court (1939), in a study on automobile demand, essentially incorporated the hedonic hypothesis that heterogeneous goods are aggregations of attributes (in today's Gorman (1980) - Lancaster (1966) sense), and that economic behaviour relates to these attributes.[1] He was first to attribute the constructed price indices as 'hedonic price indices'. However, the fact that until today hedonic analysis has been applied to a large field of quality-related issues is largely due to the work of Zvi Griliches and Sherwin Rosen. The foundations were laid by the characteristics approach of Griliches ((Griliches 1961) and (Griliches 1971)) to the construction of price indices and his subsequent work, as well as by the unifying approach of Rosen (1974), in which varying marginal implicit prices are derived from both a distribution of marginal rates of substitution and marginal rates of transformation.[2] Hedonic studies have been motivated by two main concerns. First, to identify implicit prices of attributes. And second, to investigate welfare impacts by analysing the structure of demand for attributes (Follain and Jimenez (1985), Bresnahan and Gordon (1997)).

Hedonic price analysis has found its application in several recent studies on wine, among them Golan and Shalit (1993), Oczkowski (1994), Nerlove (1995) and Combris, Lecocq, and Visser (1997). In Golan and Shalit's (1993) study on hedonic grape and wine pricing, the authors aim to identify and evaluate the wine quality characteristics of Israeli grapes. By assuming that the Californian wine market is perfectly competitive, wine prices are presumed to reflect both consumer preferences and the value of grape quality attributes. If, therefore, Californian and Israeli wine consumers have the same preferences, the competitiveness assumption can be used to derive hedonic prices for the Israeli market. By estimating the relative contribution of grape characteristics to wine quality, and using

[1] Though Gorman's paper was written in 1956, it was not published until 1980.

[2] The generalised commodity approach to demand analysis (Houthakker (1952) was the first one to present the hedonic function as a market phenomenon. Existing literature on hedonic quality measurement before Lancaster (1971) had already proved that the analysis of consumption at the level of characteristics is more powerful than the traditional analysis (Triplett 1971). A study of Gorman (1980)'s theory of linear consumption activities shows that Lancaster (1971) followed Gorman in specifying hedonic contours.

the monetary values from the Californian market, the authors are able to value the individual grape characteristics so as to provide a producer pricing schedule for Israel. This quality based pricing schedule could then serve to reduce the production of poor-quality wines, by giving Israeli farmers an appropriate incentive to supply higher quality grapes.

Oczkowski (1994) identifies the implicit valuation of table wine attributes for consumers and retailers from recommended retail prices for Australian premium table wine. On the producer side, the author suggests that the hedonic functions estimated provide important information upon which longer-term investment decisions may be made. Oczkowski includes dummy variables for producer size in the hedonic regression and argues that this allows for two effects. First, for possible price-making strategies and second, he argues in favour of viewing producer size as measuring the characteristic of 'exclusiveness'. That is, some consumers desire particular wines from small producers because of their limited availability, rarity and 'trendiness'. The author's innovative approach to the underlying dummy variable model permits explicit estimation of coefficients for all dummy variables.

Due to state intervention in the pricing of Swedish wines, Nerlove (1995) does not follow a standard hedonic regression, but assumes that variety prices are exogeneously determined and consumer preferences are expressed by the quantities of each variety they buy. Therefore, variety supplies are taken as perfectly elastic for the group of consumers being considered and the quantities of each variety consumed are regressed on the unit variety price and on the measures of quality attributes which characterise that variety. Nerlove (1995) builds on a generalisation of the 'pure repackaging' case, which Fisher and Shell (1971) label the 'variable repackaging' case of quality differences, and in which the amount of repackaging is allowed to depend on the quantity of the good. Using Swedish data from 1989-91, the price elasticity is estimated to be about - 1.65, which suggests that Swedish consumers are highly sensitive to price. Estimates of the implicit valuations of quality attributes are shown to differ greatly from those obtained from the classical hedonic regression with price as the dependent variable.

Whilst studying wine prices for the Bordeaux region, Combris *et al.* (1997) apply a stepwise regression procedure to investigate whether quality "matters" in explaining market prices. The authors suggest that for their data set, quality as measured by a jury grade assigned by professional wine tasters, is mainly explained by the 'subjective' sensory characteristics of the wine, which are unobservable when consumers choose the wine. Implicit price estimates are derived from data of a wine tasting panel that is unable to observe any of the 'objective characteristics' (grape variety, vintage year etc.), including price, of the wines they judge. By contrasting the results from this regression of market prices of Bordeaux wine with characteristics appearing on the label of the bottle with the results from an analysis of jury grades, the authors conclude that many variables which are important in explaining quality do not play a role in the determination of market prices. The authors explain their findings with taste differences between wine tasters and consumers and imperfect information on the wine consumers' behalf.

Several papers have recently addressed labelling issues explicitly when product attribute information is imperfect and asymmetric. In the context of international trade and economic growth, Basu, Chau, and Grote (2002) examine the effectiveness of eco-labels in providing a market-based solution to the under-consumption of eco-friendly products in developing and developed countries. Nimon and Beghin (1999) examine the implications of eco-labelling schemes on consumer choice sets and product quality in the trade of textile and apparel. Mahé (1997) and Bureau, Marette, and Schiavina (1998) investigate the role of information on quality attributes and the role of quality labelling in the process of agricultural trade liberalisation and in determining welfare effects from such de-regulation. Marette, Crespi, and Schiavina (1999) analyse the impact of certified quality labelling on welfare when common labelling schemes matter and asymmetric information is present. Bureau, Gozlan, and Marette (2001) investigate the informational role of quality labelling for trade policy and welfare when adverse selection matters due to the presence of risks of food hazards. In a vertical differentiation model, Ibanez and Stenger (2000) investigate the efficiency of labelling mentioning food safety as a means to reducing negative production externalities and raising consumer welfare. By expanding an AIDS model to include information effects and demographic characteristics, Teisl and Levy (1997) show that nutrient labelling can affect consumer purchase behaviour in significant ways. Van der Lans, van Ittersum, de Cicco and Loseby (2001) employ a conjoint analysis to show that PDO (Protected Designation of Origin) labels have no direct effect on consumer preferences in the case of olive oil. Bonnet and Simioni (2001) use a random-coefficients logit model of demand to recover the distribution of consumers' willingness-to-pay for labelled cheese, and to demonstrate that consumers do not value the quality signal provided by PDO labels for these French cheeses (Camembert).

This paper aims to examine wine labelling attributes by estimating hedonic price functions for still light wine that was on offer in the British off-licence market.[3] Since the following empirical analysis relies on data from 1994, we will briefly introduce developments on the supply and demand side around that period. The wine market in the United Kingdom (UK) was and is dominated by a large variety of foreign still light wine imports (more than 90 percent, value 1994). English and Welsh wine, produced from fresh grapes, accounts for only 0.3 percent (value, 1997) of domestic consumption. Two types of licences give the right to sell alcoholic beverages in the UK. The "off-licence", where the product is consumed outside the premises in which it was purchased (e.g. retail outlets), and the "on-licence" where alcohol is consumed *in situ* (e.g. pubs, clubs and restaurants). With more than 45,000 points of sale and 70 percent of total wine sales in 1993 (value), the off-licence sector dominates the wine market in the UK. Regarding the evolution of sales by country of origin, the big four traditional suppliers, France, Germany, Italy and Spain, continue to dominate but, collectively, if not in all cases individually, have seen their share eroded. Their combined share declined from 89 percent of volume of imported wine from fresh grapes in 1983 to 78 percent in 1993 and 71.5 percent in 2000 (DWI (2002)). Most countries depend heavily upon off-licence sales, with France and, to a lesser extent, Germany depending disproportionally upon the on-licence trade.

With the exception of Northern Ireland, Great Britain is the EU member which is characterised by the lowest level of per capita consumption. With 64,5 litres annual per capita wine consumption in 1992, France was leading worldwide consumption, whereas in the

[3]Still light wine is defined as the product obtained exclusively from the total or partial alcoholic fermentation of fresh grapes or fresh musts, with a total alcoholic strength usually not exceeding 15 percent volume.

UK only 12,4 litres were consumed in the same year (Robinson 1994). Considering the consumption pattern according to colour, the sales shares in 1993 by volume of imported still light wine in the UK were 63.7 percent for white, 33.2 percent for red and 2.9 percent for rose (EIU 1994).

The following analysis relies on a survey that covered 3940 bottles of still wine that were uniquely identified by objective labelling attributes (region of origin, vintage etc.). The article contributes to and distinguishes itself from the existing hedonic price literature on wine markets in several ways. First, we expand the dummy variable approach that was pioneered by Kennedy (1986) and Oczkowski (1994) to obtain a distinct and comparable contribution for each attribute to the variation of goods prices. The econometric approach addresses heteroscedasticity explicitly by using a General least squares (GLS) estimator. Second, in contrast to previous papers we do not rely on sensory characteristics. Rather, we have two sets of variables upon which we place our hypotheses. We consider objective attributes, which can be observed by consumers from the label and are thus assumed to determine the use value and tasting qualities of the wine. However, we also consider retailer traits as an additional choice variable that does not impact on the tasting qualities directly. Third, in this study, more than 14,000 observations are used that reflect the significance of 3940 uniquely identified attribute bundles (bottles of wine) in the sample of retailers. Fourth, in contrast to previous hedonic studies related to wine, we do not rely on recommended retailer prices, but rather on actual retail prices. Finally, we demonstrate the usefulness of studying the valuation of attribute information within the hedonic framework in two steps. Firstly, the revenue impact of shifts in attributes is examined at the retail level. Secondly, the welfare impact of changes in the attribute choice set facing consumers is considered.

In section (2) of the paper, the theoretical framework for describing agents' valuation of wine attributes is briefly developed from previous models of product differentiation. This is followed by a statement of objectives and hypotheses. Section (3) begins with a description of the survey data employed and provides an empirical assessment of postulates from the above. Section (4) explores marketing implications of shifts in attributes at the retailer level.

2 A hedonic price analysis

2.1 Methodological issues

Houthakker (1952) and Theil (1952) proposed independently a model of consumer choice based on product characteristics. Houthakker (1952), who assumes a continuous spectrum of product qualities, was the first to develop a market notion of hedonic prices. This contrasts with the Lancaster (1971) model and its variants, which consider hedonic price functions as a reflection of consumer behaviour only and assume a discrete spectrum of alternative qualities. In Rosen's (1974) model of product differentiation, upon which this paper relies, this market notion is developed further. Market clearing conditions determine the set of hedonic prices, where hedonic prices are defined as the implicit prices of attributes as they are revealed to economic agents from observed prices and specific amounts of those characteristics which are associated with them. What is being estimated in Rosen's (1974) description of a competitive equilibrium is the locus of intersections of the demand curves of different consumers with varying tastes and the supply functions of different producers with possibly varying technologies of production. The implicit estimated prices for quality give us, therefore, the implicit marginal valuation that consumers and producers place on a vector of attributes. Consider a vector of wine attributes $(z_1, ..., z_n)$, and a composite good with vector $\mathbf{x}$. When consumers choose one unit of wine, the maximisation of utility $U(\mathbf{x}, \mathbf{z})$ subject to the consumer's budget constraint,

$$y \geq p(\mathbf{z}) + \mathbf{x}, \tag{1}$$

where y denotes consumer income and $p(x)$ reflects the per unit price, satisfies the first-order conditions,

$$\frac{\partial p}{\partial z_i} = p_i = \frac{\partial U / \partial z_i}{\partial U / \partial \mathbf{x}}, \quad \forall\ i. \tag{2}$$

The marginal rate of substitution between wine attribute z_i and $\mathbf{x}$ equals, therefore, the marginal price of wine attribute z_i.

Following Rosen (1974), we consider a one-period model of wine consumers' choice behaviour, in which the agent chooses one wine attribute bundle at a time from among a number of different wine attribute bundles. We assume that under perfect competition, market equilibrium conditions are reflected in the valuation of the attributes. Although it is assumed that only certain attribute combinations can be selected in a reshuffled form (the consumer finds a Merlot 1993, *either* of French *or* Chilean origin), we assume that any quantity can be supplied to match consumer demand. Hence, we conjecture perfect divisibility.

2.2 Objectives and hypotheses

We aim to examine implicit prices for labelling attributes through the estimation of hedonic price functions. It is assumed that when consumers are confronted with the labels of the bottles on the shelf, a first group of categories of attributes (quality designation, grape variety, vintage, region and country of origin) determines the use value of the wine. Another category group, the originating retailer, is deemed to have no bearing on this use value and is therefore assumed as not entering the consumer's utility function for tasting qualities. Consumers' willingness-to-pay should, therefore, be determined by variables from the first group of categories only, unless retailer traits enter the utility function in an indirect way.[4] Since we have no information on individual retailer traits, we assume that the valuation which the consumer places on the name of the retailer reflects the

[4] Although Bliss (1988) does not refer to retailer traits explicitly, his use of indirect utility functions in a model of a multiproduct monopolist allows to distinguish some retailers by offering "better value" for money to the consumer.

aggregate valuation of relevant retailer traits to the consumer.[5]

Since our implicit prices are assumed to reflect an equilibrium price relationship, they can be given both a user value and a resource cost interpretation. Hence, we assume that retailers themselves incur costs to build a reputation based upon their own traits. They regard reputation as an asset, as they receive a competitive return on their reputation investment.[6] In a market where reputation effects are likely to be important, we assume that the degree of information which the consumer possesses about the wines will be reflected in his or her degree of product involvement. This degree of product involvement can be identified by the analyst from the willingness of wine consumers to differentiate between, and pay for, different attributes within the total attribute bundle. We assume, therefore, that the further down their decision trees consumers are willing to proceed, the more distinct attributes they are willing to pay for, and the higher must be their level of information about the attributes which they are comparing.

2.3 Model specification

Our variables have to undergo a modification that alters the interpretation of the estimates only. This is due to the nature of the data (dummy variables) and due to the necessity to retain comparability across attributes. The modification does not alter the underlying meaning of the implicit price estimates as 'missing prices' in a hypothetical market where both consumers and producers are asked to attribute their valuation to the existence of a particular wine attribute, *ceteris paribus*. As a result of this modification, and after adjusting the coefficient estimates with the estimated variances, the final interpretation is that the coefficient estimates measure the relative impact on the dependent variable (the unit price evaluated at the sample means) of the presence of the attribute *ceteris paribus*.

Economic theory suggests that non-linear functional forms could frequently provide a more appropriate alternative, although the choice of the functional form for the hedonic price function should remain an empirical matter. Also, on pragmatic grounds, with respect to heteroscedasticity, a non-linear form such as the semilogarithmic (log-lin) model could be preferable. In this instance, the coefficient of a dummy variable measures the percentage effect on the dependent variable of the presence of the factor represented by the dummy variable. However, Kennedy (1981) objects to the interpretation of Halvorsen and Palmquist (1980) of estimating the percentage effect on asymptotic grounds.[7] Kennedy (1981) argues that their suggested procedure leads to a biased esti-

[5] Betancourt and Malanoski (1995) provide empirical evidence of the mechanisms through which retail distribution services (cleanliness, short wait for checkout, unit pricing on shelves, convenient store location) affect demand, costs and retail competition. The authors demonstrate that for their sample of 616 supermarkets across the United States, distribution services have a positive effect on the demand for product.

[6] Shapiro (1983) demonstrates that the introduction of reputation as an asset that must initially built up allows the construction of an equilibrium model that includes perfect competition, free entry, and quality choices by firms under imperfect information.

[7] Halvorsen and Palmquist (1980) use the general form of a log-lin equation, $lnY = a + \sum_i b_i X_i +$

mator for the dummy variable. Instead of estimating g by

$$\hat{g} = exp(\hat{c}) - 1, \tag{3}$$

he suggests to follow Goldberger (1968) and to estimate g by

$$g^{\star} = exp\left(\hat{c} - \frac{1}{2}\hat{V}(\hat{c})\right) - 1, \tag{4}$$

(where $\hat{V}(\hat{c})$ is an estimate of the variance of $\hat{c}$), which is assumed to have less bias than $\hat{g}$. A procedure for adjusting dummy variable coefficient estimates which does not require to discard variables from the equation was put forward by Suits (1984). He interprets the estimates as deviations from average behaviour.[8] Following Suits (1984), we impose identifying restrictions, but instead of employing Kennedy's (1986) laborious extension of Suits (1984), we expand on Oczkowski (1994), and substitute the full constraint into the original equation. Following symmetrical estimations, it is possible to obtain all coefficient estimates. If, for example, the objective was to get coefficient estimates for wine colours (red, white, rose: C_1, C_2, C_3) and, say, three producer regions of a given county (R_1, R_2, R_3), the following constraints (5) and (6) could be substituted into the original equation (9) as,

$$\begin{aligned} \alpha_1 Pc_1 + \alpha_2 Pc_2 + \alpha_3 Pc_3 &= 0 \\ \alpha_1 &= [-(\alpha_2 Pc_2)/Pc_1 - (\alpha_3 Pc_3)/Pc_1] \end{aligned} \tag{5}$$

where Pc indicates the mean, hence the proportion of non-zero's, in the colour categories for each bottle of wine. And,

$$\begin{aligned} \beta_1 Pr_1 + \beta_2 Pr_2 + \beta_3 Pr_3 &= 0 \\ \beta_2 &= [-(\beta_1 Pr_1)/Pr_2 - (\beta_3 Pr_3)/Pr_2], \end{aligned} \tag{6}$$

where Pr reflects the proportion of non-zero's in the region categories for each bottle of wine. This, substituted into the original equation, gives

$$\begin{aligned} P = {} & [-(\alpha_2 Pc_2)/Pc_1 - (\alpha_3 Pc_3)/Pc_1]C_1 + \alpha_2 C_2 + \alpha_3 C_3 + \beta_1 R_1 \\ & + [-(\beta_1 Pr_1)/Pr_2 - (\beta_3 Pr_3)/Pr_2]R_2 + \beta_3 R_3 \end{aligned} \tag{7}$$

and,

$$\begin{aligned} P = {} & \alpha_2[C_2 - (Pc_2/Pc_1)C_1] + \alpha_3[C_3 - (Pc_3/Pc_1)C_1] \\ & + \beta_1[R_1 - (Pr_1/Pr_2)R_2] + \beta_3[R_3 - (Pr_3/Pr_2)R_2] \end{aligned} \tag{8}$$

$\sum_j c_j D_j$, where X_i denote continuous variables and D_j represent the dummy variables. When considering a single dummy variable, the interpretation of the coefficient of the dummy variable becomes more transparent when transforming the above equation to $Y = (1+g)^D exp(a + \sum_i b_i X_i)$, where $g = (Y_1 - Y_0)/Y_0$. Y_1 and Y_0 denote the values of the dependent variable when the dummy is equal to one and zero, respectively. The coefficient of the dummy variable is thus $c = ln(1+g)$, and the relative effect on Y of the presence of the factor represented by the dummy is given by $g = exp(c) - 1$. The percentage effect of the dummy variable on Y, in units of Y, is found by applying the antilog function, $100 \cdot g = 100 \cdot exp(c) - 1$, which is the percentage difference associated with being in group 1 rather than being in the reference group.

[8]Instead of forcing one of the coefficients of the dummy variables to be zero, all of them could be restricted to zero and the resulting intercept can be interpreted as the average of the intercepts of all observations in the sample.

The corresponding hedonic model assumes therefore,

$$p = \alpha_2[X_{a2}] + \alpha_3[X_{a3}] + \beta_1[X_{b1}] + \beta_3[X_{b3}] + \varepsilon. \quad (9)$$

where p is a $N \times 1$ vector of transformed observations on the dependent variable, price per bottle P, there are four $N \times 1$ vectors of $\mathbf{X}$ of observations, α and β define the unknown parameters, and ε is a $N \times 1$ vector of unknown stochastic disturbances. A symmetrical substitution generates estimates for the remaining coefficients α_1 and β_2 (symmetrical regressions). Importantly, this specification would embody an equivalence effect, if we were to apply the traditional way of dropping one category to avoid perfect multicollinearity. The effect of grape variety, for example, i.e., the estimated implicit price differences between Cabernet Sauvignon and Shiraz, would be assumed to be the same across all regions. Therefore, a model should be specified that provides sufficient flexibility to allow differential effects to show. Interaction terms will be introduced, which enable us to test for these differential effects.[9]

3 Empirical analysis

3.1 The data

The paper uses data on prices and attributes of foreign still wines from a survey that was undertaken in August 1994 in 94 retail outlets of different commercial forms in England and Scotland (see Appendix A). Retailers were selected according to market share to give a representative sample of foreign still wines sold off-licence in those regions. Each price for a bottle of wine is, where appropriate, described by a combination of the following dimensions:

country of origin	category (e.g. AOC)[10]	importer
appelation(e.g. Chianti)	brand (e.g. Gallo)	producer
region of origin	place of bottling	vintage
volume	grape variety	colour.

The survey collected thus all information that appears on the label of the bottles, except for the degree of alcohol. It reveals in how many outlets per company a uniquely identified bottle was found. We employ this information as quantity proxy. In total, the survey includes 14,440 bottles (prices) from 13 countries of origin that are identified by 575 attributes. This large number of bottles is due to the fact that there are 3940 uniquely identified bottles of still wines that appear on average in 3.7 retail outlets of the same commercial form.

[9]The interaction terms of primary interest are those for region/variety. The coefficient estimates for those product variables estimate then the differential effect of region by variety. For example, the interaction term for grape variety and region estimates the extent to which, say, the effect of being Chardonnay differs for Hunter Valley versus Napa Valley.

3.2 The functional form

Regarding the functional form in hedonic regressions, there is little theoretical guidance. Our initial objective would be to include all forms that theory shows are plausible. However, as all our explanatory variables are dummy variables, the choice of the functional form is limited to the linear and the log-lin, i.e., semilog, specification. Nevertheless, the use of interaction terms allows us to gain additional flexibility. When we employ a log-lin hedonic price function, we assume nonconstant marginal Engel prices (the prices paid for incremental units of characteristics when purchased as part of the same bundle) and constancy of relative prices with respect to changes in proportions of characteristics (Triplett 1975). This log-lin specification assumes therefore homotheticity of the utility function, hence homogeneity of degree zero of the demand equations for attributes. Since only relative prices matter, the imputed price is independent of the level of the characteristic, which appears to be a realistic and convenient assumption, since only dummy variables are used as explanatory variables in the present model. Also, since the log-lin form allows each marginal implicit price to be a nonlinear function of the entire set of characteristics, it appears as an attractive alternative hypothesis, since it accommodates the idea that bundling constraints are present for wine attributes in a bottle of wine.

3.3 Data analysis and specification search

To estimate the above functional relationship, the following modeling strategy borrows from several methodologies, namely from those frequently associated with David Hendry and Edward Leamer. The present analysis follows Leamer's (1990) 'classical' references to sensitivity analysis, and subsequent attempts to simplify the models by incorporating the insights gained from specification uncertainty diagnostics and measurement error diagnostics. Although the Hendry methodology is time series based, Hendry's 'general-to-specific' approach and the related steps, are thought to be appropriate in the present cross-sectional context (Hendry (2000)).[11] The evaluation of the resulting model by extensive analysis of residuals and predictive performance is borrowed from the final step of Hendry's analysis. We expand the above approach by applying the diagnostic framework suggested by Belsley, Kuh, and Welsch (1980), and Belsley (1986), to uncover statistical problems in an OLS framework. By proceeding in this fashion, it is hoped that the strengths of the above approaches can be applied together, so as to ensure a robust estimation procedure that provides stable implicit price estimates. We follow Leamer (1990) in distinguishing three phases in data analysis: (1) estimation, (2) sensitivity analysis and (3) simplification.

[11] See Hansen (1996) for a discussion of Hendry's specification searches and his 'general-to-specific' approach.

3.3.1 Estimation

Model selection

The following estimation and testing procedure is rigorously pursued, as theory does not provide further guidance to the inclusion of variables in the present application (it is assumed that all pre-selected variables have a resource cost/user value interpretation). For the initial hedonic regression, we make a subjective pre-selection of attribute categories that is based on information from magazines widely available in the UK for wine marketers and consumers, such as "Decanter" and "The Sainsbury's Magazine" and the daily newspapers' weekly magazines and supplements of three newspapers (Observer, Independent, Financial Times, 1994/95). Hence in the initial regressions, we include country of origin, region of origin, category, brand, importer, grape variety, colour and vintage, jointly with a subset of interaction terms: interactions for colour/country of origin, colour/region of origin, category/country of origin, grape variety/region, and grape variety/country of origin. Following this pre-selection of regressors, the subsequent selection procedure, based on the single equation hedonic approach, does not follow a purely mechanical procedure - such as stepwise regression - as the dangers of doing so are well established (e.g. Wallace and Ashar (1972); Judge and Bock (1983); Leamer (1983); Greene (2000)).

Specification tests

We begin by testing for equality of implicit price contributions. This is implemented in two ways, while relying most heavily on the second. First, we follow Berndt *et al.* (1993) and compensate for the large sample size by choosing very tight significance levels for the standard F-tests (.01 significance level). Second, we follow Ohta and Griliches (1975) and Ohta and Griliches (1986), who suggest specifically for hedonic models to consider the difference in fit between the unconstrained and constrained regressions, and not to reject the simpler hypothesis unless they are very different. Hence, we compare the standard errors (SER) of both regressions. However, we consider the null hypothesis of parameter equality only as relevant, if it is based on economic significance rather than on statistical significance. If the difference in SER of the regression is smaller than or equal to .01 in the system under the test, the null hypothesis will not be rejected on practical grounds. As the regression is semilogarithmic, an increase in SER by .01 implies an increase in the standard deviation of the unexplained component of price of about 1 percent.[12] In searching for the most parsimonious specification, we follow Berndt *et al.* (1993) in rejecting the null-hypothesis when the root mean squared errors under the alternative results in a reduction of more than 5 percent in the standard deviation of the unexplained variation of log prices. The following specification tests were applied:

[12] Consider a difference in the standard errors in the constrained and unconstrained regressions of .01 and a SER of the constrained regressions that was .1. The implication is that the lack of fit of the constrained regression is increased by 10 percent compared with that of the unconstrained regression (.01/.1 = .1). Equally, if the SER was .2, the .01 criterion implies the willingness to accept up to a 5 percent deterioration in the fit of the model as measured by the standard error of its residuals.

(a) Tests for Heteroscedasticity

The Breusch-Pagan test (Breusch and Pagan 1979) and its extension by Koenker (1981) is used. We apply weighted regressions as this has a double advantage. First, it permits us to correct for heteroscedasticity by transforming the error terms.[13] However, it also satisfies hedonic theory, as each attribute should be accounted for in terms of its market significance. Hence, using weighted regressions, where the weights reflect a proxy for the quantity demanded, should provide meaningful results.

(b) Specification tests for collinearity

Multicollinearity may give rise to two serious problems in hedonic models (Atkinson and Crocker 1987). First, the mean squared error of the estimator may cause substantial instabilities in coefficient signs and magnitudes as independent variables are added or removed from the model. Second, measurement error bias may be transferred in part to collinear variables measured without error and may alter their signs.

(*i*) As in standard analysis, we consider F-values, t-values and corrected R-square together, and ask whether there is a lack of individual significance despite overall significance and high corrected R-square. Furthermore, the Akaike information criterion (AIC) is selected here in order to attempt a judgment about the trade-off between model complexity and goodness of fit.[14]

(*ii*) We run auxiliary regressions, as collinearity can appear both in the form of linear dependence between variables, and as a lack of variation in the values of a control variable about its mean. Thus, both auxiliary regression R square *and* the sum of squared least squares residuals from the auxiliary regression are considered together (Berndt and Griliches 1993).

(*iii*) Finally, the condition number of the data matrix is examined (Belsley, *et al.* 1980). Judge, Griffith, Hill, Lütkepohl, and Lee (1985) suggest that moderate to strong near exact linear dependencies are associated with condition indices between 30 and 100.

3.3.2 Sensitivity Analysis

We aim to perform a robust estimation procedure that is able to produce estimates which are insensitive to model misspecifications. Thus, we follow Leamer (1990) in his 'classical approach' to sensitivity analysis by investigating whether inference is fragile and not believable. We apply techniques for discovering influential observations, as developed by

[13] Since the present analysis employs GLS, only one form of heteroscedasticity is tested for. Given the weights in the present study, is assumed that the error variance varies with the expected price. The consequence is that White's (White 1980) heteroscedastic-consistent covariance matrix estimation cannot be employed. The Goldfeld-Quandt test is not used as it may lack power if an error variance is present that is related to more than one variable.

[14] We prefer the AIC to the Schwarz criterion in the present context of a large number of potential variables, as the latter penalises model complexity much more heavily.

Belsley *et al.* (1980). These techniques are complemented by applying the trimmed least squares estimation method as performed by SHAZAM.[15]

Three means for deletion diagnostics are examined (Belsley *et al.* 1980). First, we consider single-row diagnostics. We investigate the change in the estimated regression coefficients that would occur if the i-th observation were deleted. This diagnostic measure (DFFIT) has the advantage of being independent from the particular co-ordinate system used to form the regression model. Scaling this measure with the standard deviation of the fit displays a scaled row-deleted change in fit (DFFITS). Second, we examine the hat matrix by studying the diagonal elements of the least-squares projection.[16] Finally, we are also running a Lagrange-Multiplier test for normality (Jarque-Bera). We exploit the link between the hat matrix and the residual variance by investigating the standardised residual (studentised residual). If the observation conforms to the model that is estimated with other observations, this standardised residual should be small (the calculation is repeated for each observation). Absolute values less than two are acceptable in terms of the model specification. Others are regarded as outliers. Since some of the most influential data points can have relatively small studentised residuals, row deletion and the analysis of residuals are studied together and on an equal footing (Belsley *et al.*, 1980: 21).

We follow Belsley *et al.* (1980: 22) to perform row deletions. The authors suggest to employ the COVRAT statistic and to compare the covariance matrix using all data with the covariance matrix that results when the i-th row has been deleted. Since this magnitude is a ratio of the estimated generalised variances of the regression coefficients with and without the i-th observation deleted from the data, it can be interpreted as a measure of the effect of the i-th observation on the efficiency of coefficient estimation (Belsley *et al.*, 1980: 48). As the two matrices differ only by the inclusion of the i-th row in the sum of squares and cross products, values of this ratio near unity can be taken to indicate that the two covariance matrices are close, or that the covariance matrix is insensitive to the deletion of row i. A value of COVRAT greater than one indicates, therefore, that the absence of the associated observation impairs efficiency.

External scaling is applied, where cut-off values are determined by recourse to statistical theory. Belsley *et al.* (1980) suggest that this procedure permits us to discover which observations are most strongly influential. If observations have a high leverage and a significant influence on the estimated parameters, enough evidence exists to view them as presenting potentially serious problems. Accounting for the above measures, we consider about 2.4 percent of the observations as occasionally influential. However, the results from the trimmed least squares estimation also suggest that parameter estimates are sufficiently stable so as to continue with weighted least squares regressions.

[15] All regressions were performed by using SHAZAM, version 7.0..

[16] This hat matrix (equation 2.15 in Besley *et al.*, 1980) determines the fitted values. Since the diagonal elements of the hat matrix have a distance interpretation, they provide a basic starting point for revealing 'multivariate outliers' which would not be revealed by scatter plots when $p > 2$.

3.4 Discussion of the empirical results

Summary statistics are presented in Appendix B. The hedonic price functions are estimated by employing a General least squares (GLS) estimator.[17] The resulting GLS regressions were performed for two reasons. First, employing GLS rather than OLS as an estimation rule is pursued on the basis that each attribute (and its price) in the context of hedonic market studies is important only to the extent that it captures some relevant fraction of the market (Griliches 1961). Here, the weights applied in the GLS regressions reflect in how many retail outlets of each retailer type (e.g. Marks and Spencer) a uniquely identified bottle was found. It is therefore implicitly assumed that the sample fractions are directly proportionate to the number of bottles sold. Second, the implementation of GLS allows us to account for heteroscedasticity due to omitted variables and/or due to misspecification.

The linear specification was rejected in favour of the log-lin model. It was suggested that certain categories of attributes (quality designation, grape variety, region and country of origin, vintage) determine the use value of the wine, and enter, therefore, the utility function of the consumer. Another category was assumed not to have any bearing on this use value (the retailer). The willingness-to-pay of the consumer would therefore be determined by variables from the first group of categories. However, the results suggest that the retailer in which the bottle is chosen (and thus the retailer traits) affect consumer choice in significant ways. Although it was not possible to compare exact attribute bundles across 'non-taste attributes' (namely the retailers), distinct and significant valuation of retailers were identified. The results indicate that consumers attach a high value to the information provided on the label. In all cases where conditional effects between attributes were found to have a significant impact on price, consumers are viewed as regarding these attribute bundles as imperfect substitutes. In these instances of more than overall impacts, outstanding grape varieties are shown to have a strongly positive or negative regional impact on price just as outstanding regions have a similar grape varietal impact.

The estimation results of the log-lin hedonic model are given in Appendix *C*. The estimates are interpreted as follows. The valuation which a consumer is assumed to place on the colour of wine is as anticipated, as the parameter estimates for red (+2.2), white (-1.6) and rosé (-11.2) take the expected signs.[18]

When we consider the countries of origin, being French achieves the greatest impact on price (+12.3 percent), whereas Romania shows the greatest negative impact (-48.4 percent). Perhaps surprising is the only moderate positive impact of Australian origin (+5.3 percent), especially when compared to France, and a rather high negative impact on price of wines of Chilean origin (-10.5 percent). As for Chilean wines, the sluggish expansion of imports and the increase in popularity that is only recent, following the introduction of new wine-making technologies, may be part of the explanation. Not

[17] The regressions were implemented as weighted least squares regressions, where ordinary least squares (OLS) are applied to a transformed model.

[18] Since the majority of top quality wines is red rather than white, the relative valuation is as expected (assuming we regard red and whites not as complements).

necessarily surprising is the highly different impact of being of New Zealand origin, as compared to average (+25.5 percent). This valuation could be explained by the fact that Chardonnay as well as Sauvignon Blanc produce probably the top quality whites from the Island (+15.3 percent and +11.7 percent respectively). Equally expected is the impact on price of German (-28 percent), Bulgarian (-39.2 percent), Hungarian (-34.4 percent) and Romanian origin (-48.4 percent).

The valuation for two mediterranean competitors, Italy (-3.9 percent) and Spain (-20.3 percent) meets also our expectations. Italy represents thus the most typical of average prices amongst all countries of origin. Although Italian Merlot, Cabernet Sauvignon and Chardonnay is sold in the British market, these grape varieties do not represent the qualities for which Italy has long been known. The expected negative price contribution therefore comes about because Italy is only recently increasing its supply of the world's most favourite grape varieties. Spain's reducing impact on price could be partly explained by the (low) performance and high importance (in volume terms) of its wines from La Mancha (-26.9 percent), where nearly half of Spain's production originates, as well as from Valencia (-25.7 percent). When we consider the impact of regions on price, our expectations are thus met for La Mancha, for Rioja (+18.9 percent), as well as for Provence (-29.8 percent), the Côte Chalonnaise (+45.7 percent) and Veneto (-25.96 percent). However, surprising is both the negative impact of Sonoma Valley (-16.74 percent) as well as that of the Douro (-12.2 percent).[19] While wines from the Douro valley have recently become more highly valued, it is hard to explain the poor performance of Sonoma. As for French regions, an implicit valuation of the AOC system seems to be revealed, since the impact of Libourne (+43 percent), Medoc (+52.3 percent) and Sauternes (+133.3 percent), all in the heart of Bordeaux, is distinctly higher than that of generic Bordeaux wine (-17.3 percent). Unexpected, however, may appear the magnitude of impact for Chablis and Côte de Beaune (+48 percent and +148.4 percent respectively). However, consider the high impact of wines from the Côte de Nuit (+152.1 percent) jointly. The fact that top quality reds come from both Côte de Beaune and Côte de Nuit, whereas Chablis is highly regarded for its Chardonnay, seems to be also reflected in the relative contribution of white versus red wines.

In the case of grape varieties, the coefficient estimates are striking in that consumers appear to value the price premium associated with Chardonnay from any origin more than twice as highly as they do Cabernet Sauvignon (+15.3 percent and +7.3 percent respectively), and this on the background of a reverse valuation in terms of colour and the fact that both grapes take the largest proportion amongst the red and white wines in the sample, respectively. Comparing the grape varieties according to colour, the high valuation of Riesling relative to Chardonnay seems also somewhat surprising, in particular if the highly negative impact on price of Riesling from Australia (-34.9 percent) is considered. However, since Riesling is a rather classical grape for France and Germany, the high valuation might be associated with those countries, whereas Australia is more valued for its Chardonnay. In line with our expectations is the highly positive impact of Sauvignon Blanc (+11.7 percent) relative to Semillon and Sangiovese, especially when taking account of its classic background from the Loire Valley, Bordeaux and its rising

[19]Classic regions that have a good reputation for their quality are Rioja, Côte Chalonnaise, and Sonoma Valley.

success in the New World. However, all the more surprising is that the national impact of Sauvignon Blanc is highly negative, for both France (-22.2 percent) as well as for Chile (-20 percent). Perhaps surprising is the impact that Chardonnay has in the case of Spain (+21.2 percent) and Italy (-18.7 percent). Given its classic roots in Burgundy and its success in Australia, the significant impact in the case of Spain appears to be particularly unexpected. The negative valuation of Italian Chardonnay, however, seems to support the above suggestion that consumers may not consider Italy as a classic source for 'quality Chardonnay'. When we consider red varietals, the highly positive impact on price of Pinot Noir (+25.7 percent) relative to Cabernet Sauvignon (+7.3 percent) is not too surprising, accounting for the impact of the Côte de Nuit, the heartland of Pinot Noir. It was therefore expected to find that Pinot Noir shows more than just an overall impact, and that consumers would value it regionally, as reflected in its interaction term. This, however, was not the case.

The valuation of the different vintages should be regarded with caution. First, a relatively high level of aggregation in this all-country model makes the interpretation difficult. Second, if unmeasured quality attributes make certain vintages survive in the market, the vintage coefficients could reflect these unmeasured quality differences among the surviving wines.[20] However, a rather consistent pattern emerges, whereby the increasing valuation of older vintages reflects both interest rate differentials as well as cost of storage. Nevertheless, the 1986 and 1988 vintages stand out as being particularly valued (+52.4 percent and +28.8 percent, respectively).

Estimates for the retailers suggest that retailer traits are valued as expected in the case of Asda (-11.3 percent) and Marks and Spencer (+23.1 percent).[21] However, the rather high impact of Co-op on price is somewhat surprising (+12 percent), though it may be partly explained by consumers valuing its long opening hours and may reflect that the retailer acts as a monopolist in its local area.

4 Marketing implications

4.1 The revenue impact of shifts in attributes

Bearing in mind the interpretation of our estimation results, we can demonstrate that implicit price estimates could be usefully employed, even if we are lacking the necessary information to analyse the structure of demand for attributes. Consider that a labelling attribute is found to explain a positive or negative deviation from the unit price evaluated at the sample means. In this case, a retailer could investigate the revenue impact of altering a particular range of labelling attributes on display. Supposing the retailer intends to shift the available attributes on display from French Sauvignon Blanc (FSB)

[20] See Berndt *et al.* (1993) for a discussion of age coefficients among microcomputers.

[21] Asda is a large grocery and non-food retailer, known for high volume and value for money. Marks and Spencer, in contrast, whose reputation is built on quality, dependability and good value, is a traditional retailer that is tailored towards consumer groups with higher income.

to Chilean Sauvignon Blanc (CSB), a proportionate adjustment to the mean price can be found in three steps. First, we need to identify the proportionate loss for the type of wine that is replaced, and the standard errors involved. Second, we need to account for the market share of wine to be removed from the overall sample. Third, we collect the adjusted premium for the affected (grape) variety, and weight this pivot variable by the results from the first and second step. The following box (Box 1) aims to demonstrate the implementation of this procedure.

Box 1: *The revenue impact of a shift in labelling attributes*

Step 1: It is necessary to identify the proportionate loss for the pivotal attributes that we wish to replace. Therefore, all the attributes involved for which explicit coefficients have been estimated have to be identified first. As shown below, we should also account for the certainty of the joint effects, as derived from the variance covariance matrix of the estimated coefficients. However, we first proceed by computing the proportionate loss and the corresponding standard errors in three sub-steps.

(a) Find the total sum of the relevant estimated coefficients:

Chile	France	CSB	FSB	TOTAL
-.1103	-.1156	-.2234	+25.05	= -.1988

(b) Compute the corresponding joint standard error, assuming initially that all the parameters have zero covariances:

SE Chile	SE France	SE CSB	SE FSB	TOTAL SE
$[(.0217)^2$	$+(.003)^2$	$+(.0238)^2$	$+(.0282)^2]^{1/2}$	=.0429

(c) Find the proportionate loss or gain from the log-lin model, considering both all relevant estimated coefficients and the corresponding certainty of the joint effects, as in equation (4),

$$g^{\star} = exp\left(\hat{c} - \tfrac{1}{2}\hat{V}(\hat{c})\right) - 1,$$

where $\hat{V}(\hat{c})$ is an estimate of the variance of $\hat{c}$, the coefficient of the dummy. Therefore,

$$g^{\star} = exp\left[(-.1988) - \tfrac{1}{2}(.0429)^2\right] - 1 = -.181$$

If we want to take an estimate of the variance into account, this can be done by pre-multiplying the corresponding segment of the variance-covariance matrix by (1,-1,1,-1), recognising positive and negative correlation between coefficients, and then post-multiplying by the transpose of this unit vector. The corresponding standard error estimate is 11.25. We can confirm this result by considering that for any random variable x and y, $var(x+y) = var(x) + var(y) + 2cov(x,y)$.

In our example, we add the covariances of those variables that move together and consider that the variances of all those variables that move into the opposite direction subtract. This will result in the same standard error. As a result, the proportionate loss accounting for the variance estimate is 18.5 percent (18.1 percent from the above), and applies to the market share of the desired attribute bundle (French Sauvignon Blanc).

$$g^{\star} = exp\left[(-.1988) - \tfrac{1}{2}(.1125)^2\right] - 1 = -.185$$

Step 2: Identify the market share of the attribute bundle to be removed from the overall sample, hence the retailer's intended stock transfer: In our example, the 184 bottles of French Sauvignon Blanc correspond to 1.27 percent of the total sample of 14.440 bottles.

Step 3: Obtain the adjusted premium for the affected attribute (Sauvignon Blanc) by applying Kennedy's (1981) adjustment (equation (4)), and weight this pivot attribute (the grape variety) by the results from step (1) and (2). The adjusted coefficient for Sauvignon Blanc is +11.69 percent (Appendix C, Table 3). As a result, the monetary impact of this stock transfer, hence the proportionate adjustment to the overall mean price, is:

$$1.1169 \times .01274 \times (-.1854) = -.00264 \ percent.$$

Given the mean price of 551 pence per bottle, the proportionate adjustment to the mean price would be - 1.45 pence (-1.42, respectively) a bottle, if a stock-transfer of French Sauvignon Blanc to Chilean Sauvignon Blanc was intended.

The proportionate adjustment to the overall mean price is derived under the assumption that the retailer can shift the attribute bundle (from Chile) without offering special discounts when more is purchased. Thus, demand is assumed perfectly price elastic.

This will be an acceptable assumption if we are considering consumer demand from an individual retailer.

However, with knowledge of implicit price estimates from the above labelling attributes, information on retailer traits and possibly consumer characteristics, we could derive more extensive predictions about the retailer's stock planning and supply decisions.[22]

5 Concluding remarks

We have employed hedonic price analysis to reveal the values which market participants place on labelling information. Estimation results deliver information on wine consumer preferences for attributes contained in the label on wine bottles. By means of a parametric approach, implicit prices for these attributes are derived from prices and quantities of wines sold in the British off-licence market.

The results suggest that consumers attach a high value to the information about those attributes, namely the retailers, that were initially assumed as having no bearing on the use value of the wines. Interaction terms are employed in order to reveal the differential effects between attributes, and where these are found to be relevant, consumers are viewed as regarding attribute bundles as imperfect substitutes. Therefore, Chardonnay from Spain, Chardonnay from Italy, Sauvignon Blanc from Chile, Sauvignon Blanc from France, Chenin Blanc from New Zealand, Semillon from France and Riesling from Australia are considered as distinctly different attribute bundles. A highly distinct valuation of grape varieties according to region of origin emerges only for Australia and France. When accounting for the relative importance both of grape varieties and of regional origins, the results suggest an asymmetry between possibly the most classical 'New World' wine producer, Australia, and the most classical 'Old World' wine producer, France. Results indicate that grape varieties are highly important in the choice of Australian wines, whereas regional origins are valued most in the case of French wines.

We demonstrate that implicit price estimates could be usefully employed, even if we are lacking the necessary information to analyse the structure of demand for attributes. The valuation of attribute information as derived from hedonic analysis permits the analyst to determine the revenue impact of shifts in attributes at a given stage in the marketing chain. Thus, both marketers and producers could achieve a more efficient tailoring of marketing and production efforts to specific consumer groups due to their knowledge of consumers' valuation of labelling attributes. Revenue implications of changes in labelling policy on the retail level could thus be considered.

However, several caveats remain. The analysis is inherently static and does not account explicitly for valuation due to repeat purchases or different advertising intensity across wines. Due to the nature of the data (dummy variables), limited functional flexibility

[22] Although parameters of attribute demand could be inferred from observed consumer choices and implicit prices of characteristics, the direct derivation of price and income elasticities from attribute demand bears several well-known problems (Murray (1983), Ohsfeldt and Smith (1985)).

may limit the validity of the estimates. However, early studies have already shown that such constraints may not be as limiting as initially considered ((Butler 1982), (Bartik and Smith 1987)). Furthermore, the question remains as to whether the attributes included as variables in the regression are proxies for other attributes, which themselves are the 'true' attributes in the eyes of the consumers. In future analysis, the hedonic framework should, therefore, be accompanied by performing a conjoint analysis. However, if conjoint analyses treat price as an attribute of the good, the relation between part-worth utility and revealed preference is not as clear as it is in hedonic analysis. Also, conjoint analysis assumes that consumers behave as though tradeoffs are being considered, yet the tradeoff model may be only a gross approximation to the actual decision rules that are employed (Payson (1994)). In contrast, hedonic pricing allows the identification of consumer preferences in the proximity of observed choices and thus avoids some of the well-known biases that arise in conjoint analysis from a survey of consumers' willingness-to-pay for hypothetical items.

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APPENDIX A

Table 1: Retail outlets distinguished by commercial forms

27 Supermarket outlets	**37 Wine specialist outlets**	**18 Hypermarket outlets**	**5 Large retailer outlets**	**7 Others**
7 Tesco	4 Wine Rack	6 Asda	2 Littlewoods	1 Coop
3 Coop	14 Victoria Wines	1 Morrisons	3 Marks and Spencer	1 Cullen's
1 Somerfield	3 Unwin's	1 Safeway		1 Europa Food
1 Kwiksave	8 Thresher	6 Sainsbury		1 Gateway
6 Safeway	2 Oddbins	1 Scotmid (Coop)		1 Independant
6 Sainsbury	2 Majestic	3 Tesco		1 Kwiksave
3 Waitrose	2 Cellar Five			1 Spar
	1 Bottom's up			
	1 Haddows			

Source: CFCE, 1994

APPENDIX B
Table 2: Summary statistics

Variable description	Number of observations	Mean***	Standard Deviation	Variance	Minimum	Maximum
PRICE (£)	14440** (3940*)	(5.51)	(4.5752)	(20.932)	1.09	99.99
RED	6933	0.49137	0.49999	0.24999	0	1
WHITE	7111	0.48655	0.49988	0.24988	0	1
ROSE	396	2.23E-02	0.14779	2.18E-02	0	1
ARGENTINIA	35	2.54E-03	5.03E-02	2.53E-03	0	1
AUSTRALIA	1495	6.95E-02	0.25441	6.47E-02	0	1
GERMANY	801	6.75E-02	0.25094	6.30E-02	0	1
BULGARIA	314	1.93E-02	0.13756	1.89E-02	0	1
CHILE	248	1.78E-02	0.13212	1.75E-02	0	1
SPAIN	1067	6.29E-02	0.24289	5.90E-02	0	1
HUNGARY	281	1.47E-02	0.12045	1.45E-02	0	1
ITALY	1240	8.83E-02	0.2838	8.05E-02	0	1
NEW ZEALAND	502	2.18E-02	0.14614	2.14E-02	0	1
PORTUGAL	485	2.92E-02	0.16835	2.83E-02	0	1
ROUMANIA	55	4.06E-03	6.36E-02	4.05E-03	0	1
SOUTH AFRICA	405	3.05E-02	0.17186	2.95E-02	0	1
FRANCE	7062	0.55838	0.49664	0.24665	0	1
CAB SAUV	841	5.00E-02	0.21797	4.75E-02	0	1
CHARDONN	1152	5.71E-02	0.23208	5.39E-02	0	1
CHENIN BLANC	73	4.57E-03	6.74E-02	4.55E-03	0	1
GEWÜRZ-TRAMINER	76	7.36E-03	8.55E-02	7.31E-03	0	1
PINOT NOIR	181	1.29E-02	0.11305	1.28E-02	0	1
RIESLING	227	2.03E-02	0.14106	1.99E-02	0	1
SANGIOVESE	14	1.52E-03	3.90E-02	1.52E-03	0	1
SEMILLON	918	5.25E-02	0.22314	4.98E-02	0	1
SAUVIGNON	712	3.43E-02	0.18193	3.31E-02	0	1
VINTAGE-83	16	2.03E-03	4.50E-02	2.03E-03	0	1
VINTAGE-85	48	4.57E-03	6.74E-02	4.55E-03	0	1
VINTAGE-86	75	9.14E-03	9.52E-02	9.06E-03	0	1
VINTAGE-87	216	1.57E-02	0.12447	1.55E-02	0	1
VINTAGE-88	304	2.79E-02	0.16476	2.71E-02	0	1
VINTAGE-89	608	5.23E-02	0.22263	4.96E-02	0	1
VINTAGE-92	2780	0.20964	0.40711	0.16574	0	1
VINTAGE-94	109	6.35E-03	7.94E-02	6.31E-03	0	1
ASDA	530	3.17E-02	0.17529	3.07E-02	0	1
CWS	118	1.17E-02	0.10743	1.15E-02	0	1
COOP	131	1.68E-02	0.12835	1.65E-02	0	1
MARKS AND SPENCER	216	2.28E-02	0.14942	2.23E-02	0	1
SAFEWAY	329	2.11E-02	0.14362	2.06E-02	0	1
BREEDE	5	1.27E-03	3.56E-02	1.27E-03	0	1
COONAWARA	162	6.35E-03	7.94E-02	6.31E-03	0	1
HUNTER VALLEY	172	5.33E-03	7.28E-02	5.30E-03	0	1
LA MANCHA	61	5.08E-03	7.11E-02	5.05E-03	0	1
RIOJA	301	1.95E-02	0.13844	1.92E-02	0	1
VALENCIA	239	1.40E-02	0.11734	1.38E-02	0	1
VENETO	399	2.84E-02	0.16621	2.76E-02	0	1
DOURO	99	6.35E-03	7.94E-02	6.31E-03	0	1
BORDEAUX	760	5.63E-02	0.23062	5.32E-02	0	1
LANGUEDOC	1617	0.10939	0.31217	9.74E-02	0	1
LIBOURNE	139	1.17E-02	0.10743	1.15E-02	0	1
MEDOC	348	3.32E-02	0.17931	3.22E-02	0	1
PROVENCE	101	9.14E-03	9.52E-02	9.06E-03	0	1
SAUTERNES	32	4.06E-03	6.36E-02	4.05E-03	0	1
CÔTE CHALONNAISE	56	5.84E-03	7.62E-02	5.81E-03	0	1
CÔTE BEAUNE	134	1.88E-02	0.13577	1.84E-02	0	1
CHABLIS	176	1.57E-02	0.12447	1.55E-02	0	1
CÔTE DE NUIT	139	1.42E-02	0.11838	1.40E-02	0	1
SONOMA VALLEY	342	5.58E-03	7.45E-02	5.55E-03	0	1

* There are 3940 unique and hence different bottles in the sample. Corresponding descriptive statistics are in brackets. Since the same unique bottle appears frequently in different outlets, the total sample size is 14440.

** The difference between the total sum of all observed prices after accounting for replicates [14440] and the sum of observations for the above attributes as they remained in the final specification, is therefore due to (a) statistically non-significant attributes and (b) the nature of the data set (some wines are specified by less attributes than others: (i) indication of the retailer's name from which the price was collected is only given if the retailer's name appears on the label of the bottle, or (ii) it is due to legal restrictions, i.e. EU or national law does not allow to indicate the region of origin or the vintage for certain wines).

*** The sample mean applies to the observations not accounting for replicates, which explains the divergence between the proportion of non-zero's of each attribute in each category (i.e. the mean value) and the number of observations.

APPENDIX C
Table 3: Estimation results of the log-lin hedonic model

Variable description	Variable name	Relative impact (percent)	Estimated coefficient	Standard Error	T-Ratio
	CONSTANT		1.5234	1.73E-02	479.90
COLOUR	*RED	2.20	2.17E-02	3.33E-03	6.52
COLOUR	WHITE	-1.63	-1.65E-02	3.39E-03	-4.86
COLOUR	ROSE	-11.23	-0.11894	1.60E-02	-7.43
COUNTRY OF ORIGIN	*ARGENTINIA	-26.19	-3.02E-01	5.29E-02	-5.71
COUNTRY OF ORIGIN	AUSTRALIA	5.32	5.18E-02	9.39E-03	5.52
COUNTRY OF ORIGIN	GERMANY	-28.04	-0.32895	1.16E-02	-28.49
COUNTRY OF ORIGIN	BULGARIA	-39.17	-0.49689	1.79E-02	-27.82
COUNTRY OF ORIGIN	CHILE	-10.47	-0.11033	2.17E-02	-5.09
COUNTRY OF ORIGIN	SPAIN	-20.29	-0.2267	1.46E-02	-15.54
COUNTRY OF ORIGIN	HUNGARY	-34.36	-0.42088	1.90E-02	-22.15
COUNTRY OF ORIGIN	ITALY	-3.99	-4.07E-02	1.07E-02	-3.82
COUNTRY OF ORIGIN	NEW ZEALAND	25.55	0.22765	1.52E-02	14.99
COUNTRY OF ORIGIN	PORTUGAL	-19.89	-0.22167	1.61E-02	-13.75
COUNTRY OF ORIGIN	ROUMANIA	-48.44	-0.66143	4.33E-02	-15.28
COUNTRY OF ORIGIN	SOUTH AFRICA	-16.31	-0.17792	1.65E-02	-10.78
COUNTRY OF ORIGIN	FRANCE	12.26	0.11566	3.00E-03	38.57
GRAPE VARIETY	CABERNET SAUVIGNON	7.26	7.01E-02	1.20E-02	5.83
GRAPE VARIETY	CHARDONNAY	15.33	0.14265	1.09E-02	13.13
GRAPE VARIETY	CHENIN BLANC	-8.91	-9.25E-02	3.84E-02	-2.41
GRAPE VARIETY	GEWÜRZTRAMINER	35.06	0.30119	3.59E-02	8.38
GRAPE VARIETY	PINOT NOIR	25.73	0.22924	2.41E-02	9.52
GRAPE VARIETY	RIESLING	36.56	0.31192	2.51E-02	12.42
GRAPE VARIETY	SANGIOVESE	-34.12	-0.41385	8.41E-02	-4.92
GRAPE VARIETY	SEMILLON	-28.82	-0.33959	2.48E-02	-13.70
GRAPE VARIETY	SAUVIGNON	11.69	0.11072	1.59E-02	6.97
INTERACTION TERM	RIESLING-AUSTRALIA	-34.93	-0.42895	3.91E-02	-10.97
INTERACTION TERM	SEMILLON-FRANCE	12.10	0.11463	2.85E-02	4.02
INTERACTION TERM	SAUVIGNON-FRANCE	-22.19	-0.25047	2.82E-02	-8.90
INTERACTION TERM	CHENIN BLANC-NZ	-31.99	-0.38472	4.05E-02	-9.51
INTERACTION TERM	SAUVIGNON-CHILE	-20.04	-0.22339	2.38E-02	-9.37
INTERACTION TERM	CHARDONNAY-SPAIN	21.23	0.19389	5.24E-02	3.70
INTERACTION TERM	CHARDONNAY-ITALY	-18.67	-0.20566	4.57E-02	-4.50
VINTAGE	1983	69.98	0.53356	7.83E-02	6.82
VINTAGE	1985	35.89	0.30769	4.51E-02	6.8
VINTAGE	1986	52.44	0.4222	3.55E-02	11.91
VINTAGE	1987	27.55	0.24358	2.14E-02	11.38
VINTAGE	*1988	28.80	0.25322	1.75E-02	14.50
VINTAGE	1989	14.60	0.13636	1.18E-02	11.61
VINTAGE	1992	-10.53	-0.11126	3.99E-03	-27.87
VINTAGE	1994	-15.35	-0.16615	3.07E-02	-5.42
RETAILER NAME	ASDA	-11.33	-0.12022	1.22E-02	-9.90
RETAILER NAME	CO-OP	12.03	0.11389	2.41E-02	4.74
RETAILER NAME	MARKS AND SPENCER	23.11	0.20807	1.83E-02	11.37
RETAILER NAME	CWS	-12.01	-0.1276	2.71E-02	-4.71
RETAILER NAME	*SAFEWAY	-6.25	-6.44E-02	1.59E-02	-4.06
REGION OF ORIGIN	HUNTER VALLEY	23.41	0.21069	2.56E-02	8.23
REGION OF ORIGIN	*BREEDE	-25.02	-0.27812	1.40E-01	-1.99
REGION OF ORIGIN	COONAWARA	39.06	0.33012	2.67E-02	12.38
REGION OF ORIGIN	LA MANCHA	-26.98	-0.3135	4.21E-02	-7.44
REGION OF ORIGIN	RIOJA	18.89	0.17332	2.28E-02	7.59
REGION OF ORIGIN	VALENCIA	-25.66	-0.29628	2.46E-02	-12.07
REGION OF ORIGIN	VENETO	-25.96	-0.30042	1.85E-02	-16.20
REGION OF ORIGIN	DOURO	-12.19	-0.12939	3.53E-02	-3.67
REGION OF ORIGIN	SONOMA VALLEY	-16.74	-0.18303	1.73E-02	-10.56
REGION OF ORIGIN	BORDEAUX	-17.33	-0.19021	1.09E-02	-17.49
REGION OF ORIGIN	LANGUEDOC	-28.33	-0.33309	7.57E-03	-44.03
REGION OF ORIGIN	LIBOURNE	43.02	0.35815	2.63E-02	13.62
REGION OF ORIGIN	MEDOC	52.29	0.42075	1.66E-02	25.36
REGION OF ORIGIN	PROVENCE	-29.79	-0.35315	3.15E-02	-11.2
REGION OF ORIGIN	SAUTERNES	133.33	0.84878	5.49E-02	15.47
REGION OF ORIGIN	CôTE CHALONNAISE	45.72	0.3774	4.13E-02	9.14
REGION OF ORIGIN	CôTE BEAUNE	148.41	0.91027	2.61E-02	34.88
REGION OF ORIGIN	CHABLIS	48.01	0.3924	2.35E-02	16.73
REGION OF ORIGIN	CôTE DE NUIT	152.10	0.92502	2.63E-02	35.16

- 14.380 degrees of freedom
- Adjusted R-square: 0.52.
- Breusch-Pagan: Chi-Square = 138 with 59 regressors [for 59 D.F., P(chi-square > 77.93) = 0.05].
- Variables preceded by a * are taken from symmetric regressions.
- The relative impact of the attribute on price is measured as in equation (4).