International Agricultural Trade and Policy Center

IMPORTS VERSUS DOMESTIC PRODUCTION: A DEMAND SYSTEM ANALYSIS OF THE U.S. RED WINE MARKET

BY

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IMPORTS VERSUS DOMESTIC PRODUCTION: A DEMAND SYSTEM ANALYSIS OF THE U.S. RED WINE MARKET

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Abstract: This research estimates price and expenditure elasticities of U.S. red wine imports from five countries--Italy, France, Spain, Australia, and Chile—which are compared to elasticities of domestically produced red wine using the first-difference version of the almost ideal demand system (AIDS). Expenditure elasticity results indicate that if U.S. total expenditures on red wine increase, domestic producers would gain most. Empirical results for conditional own-price elasticities of demand indicate that U.S. and Chilean red wines are elastic while U.S. demand for red wines from other countries are highly inelastic. Due to the magnitude of consumption of U.S. domestic red wines relative to imports, an increase in the price of U.S. wine results in a decline in quantity demanded that is six times larger than that for French and Italian red wines and over 20 times larger than that of other import countries. Results suggest that U.S. red-wine producers could increase their total revenue by decreasing prices, while Italian and French producers can increase total revenues by increasing prices.

Keywords: imports, red wines, Almost Ideal Demand System, AIDS

IMPORTS VERSUS DOMESTIC PRODUCTION: A DEMAND SYSTEM ANALYSIS OF THE U.S. RED WINE MARKET

Introduction

The overall goal of this research is to estimate U.S. demand for red wine in order to obtain price and expenditure elasticities using a difference version of the almost ideal demand system (AIDS). This analysis examines both import demand as well as demand for domestically produced red wines. U.S. imports of red wines increased over 330 percent in the last decade, and red wines account for 56 percent of total wine imports by quantity (U.S. Department of Commerce-Bureau of the Census). Despite its importance, few studies were identified in the literature that examine U.S. wine trade and, more specifically, U.S. import demand for red wine. This research seeks to fill this void by estimating demand elasticities for U.S. red wine imports and assessing the impacts of these imports on consumption of domestically produced red wines.

Trends in the U.S. Wine Industry

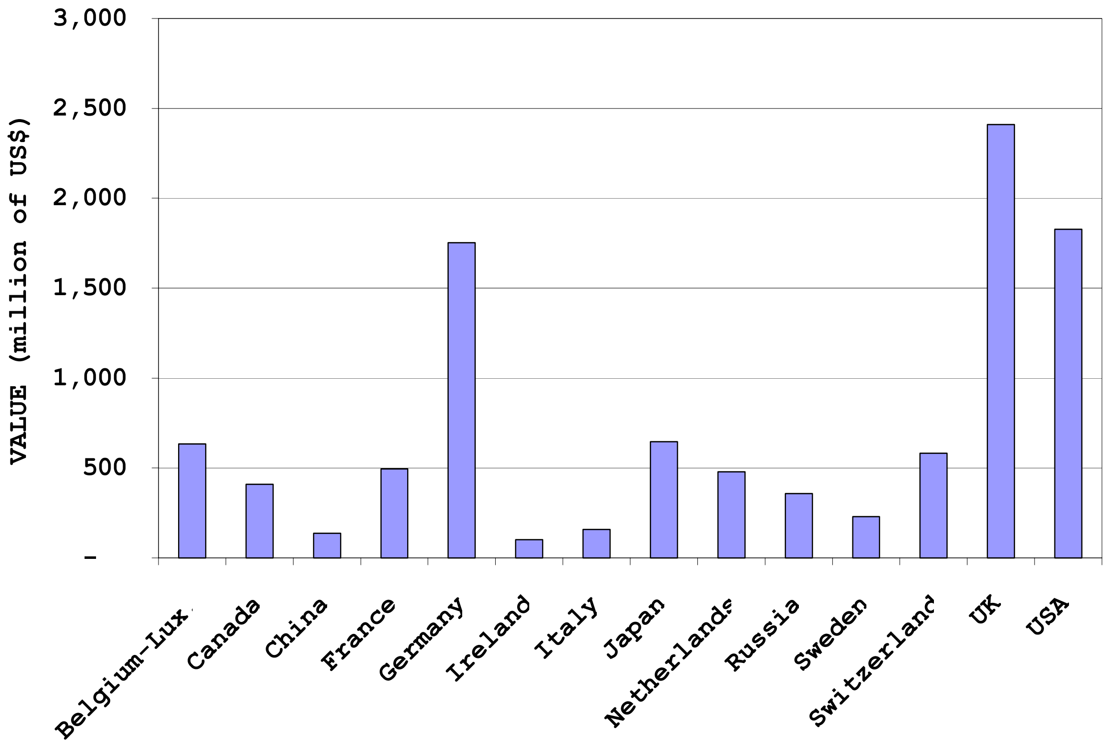
In 1997, total U.S. wine production was 17.6 million hectoliters, had a corresponding value of 19 billion dollars (Wine Institute), and accounted for almost 8% of total U.S. agricultural output (USDA-FAS). The value of U.S. wine production increased dramatically in the last decade due to considerable improvement in overall quality, product refinement, and an appreciation of American wines by consumers (Wine Institute).

Additionally, the U.S. is the second largest importer of wine in the world by value (Figure 1; FAO). U.S wine imports increased 200% between 1989 and 1998 (USDA-FATUS) while that of U.S. imported red wines more than tripled (U.S. Department of

Commerce-Bureau of the Census). In 1998, wine imports totaled 1.9 billion dollars, for a total volume of 4.1 million hectoliters, and were 5% of total U.S. agricultural imports (USDA-FATUS). Additionally, U.S. wine imports equal 23% of total U.S. wine consumption (Wine Institute), and red wine imports account for 56% of total U.S. wine imports by quantity. This compares to 35% for white wines and 9% for sparkling wines (U.S. Department of Commerce-Bureau of the Census). Since red wines constitute the majority of wine imports, this research focuses on red wines.

Tremendous growth in U.S. red wine imports has occurred. Total U.S. imports of red wines increased 350% in terms of value and 330% in terms of quantity in the period 1989-1998 (U.S. Department of Commerce-Bureau of the Census). In 1998, the U.S. imported 1.9 million hectoliters of red wines, valued at 848 million dollars. Italy, France, Spain, Australia and Chile are the major red wine exporters to the United States. These countries accounted for more than 94% of total U.S. imports of red wines by value and

Figure 1. Wine Imports by Country, 1997

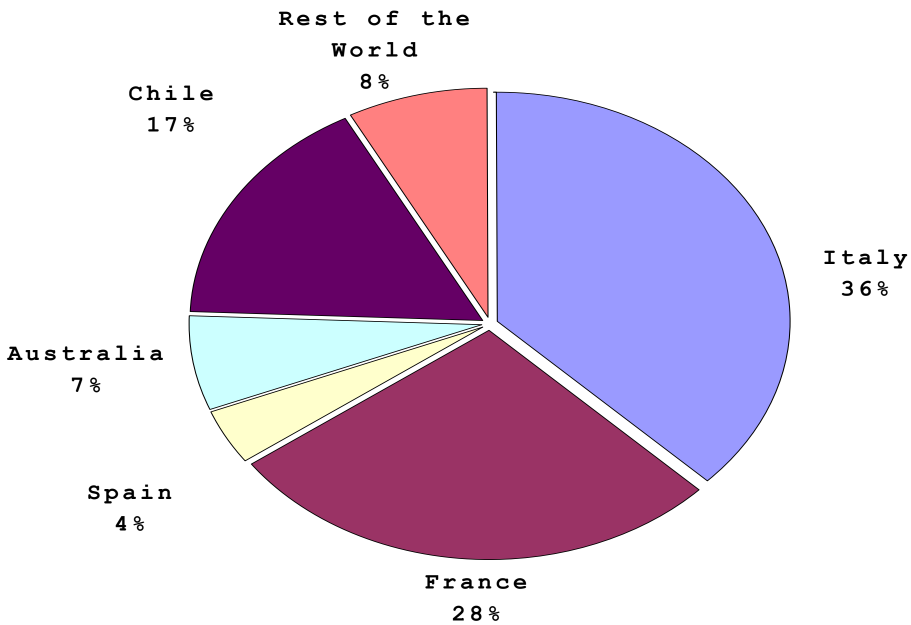


Source: Food and Agriculture Organization (FAO) of the United Nations, 1998.

more than 92% by quantity in 1997 (U.S. Department of Commerce-Bureau of the Census). Italy and France are the dominant red wine exporters to the United States, and they accounted for 78% of total U.S. red wine imports by value and 64% by quantity (Figure 2).

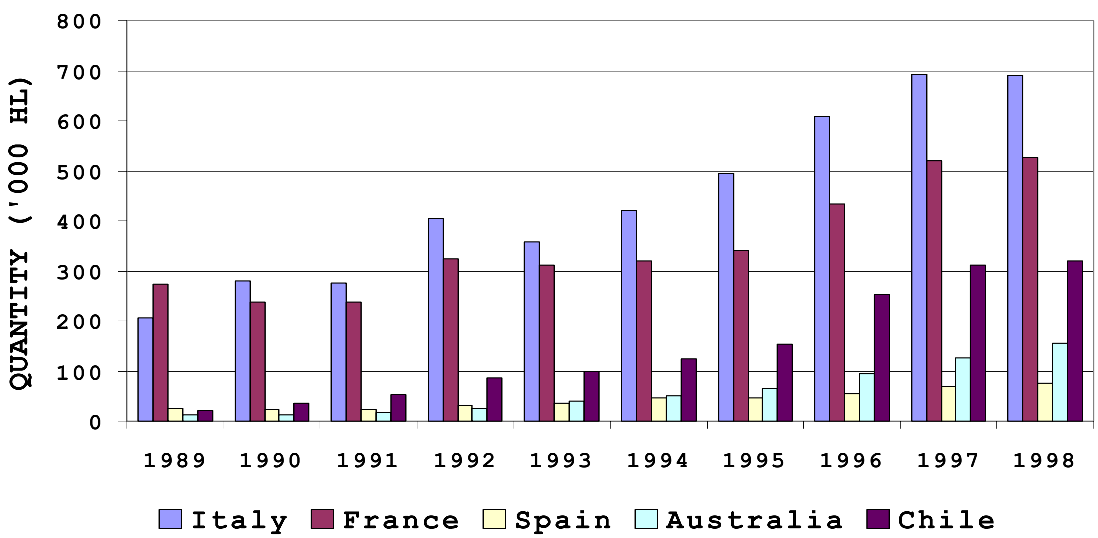
Although all five red wine-exporting countries increased their export quantity to the U.S., Australian and Chilean red wines experienced the most dramatic growth (Figure 3). Italy increased red wine exports to the U.S. nearly four times between 1989 and 1998, France 2.4 times, Spain 4.7 times, Australia nearly 14 times, and Chile almost 20 times.

Figure 2. U.S. Red Wine Imports by Country, 1997
(Total quantity equals 186,168,000 HL)



Source: U.S. Department of Commerce-Bureau of the Census, 1999.

Figure 3. U.S Red Wine Imports by Selected Countries, 1989-1998



Source: U.S. Department of Commerce-Bureau of the Census, 1999.

While consumption of red wine imports is significant, U.S. consumption of domestically produced red wines greatly exceeds total import consumption. In 1999, the share of domestic red wines consumed relative to total U.S. consumption of red wines was 63% while that of French, Italian, Spanish, Australian, Chilean, and the rest-of-world (ROW) was 14%, 11%, 2%, 5%, 3%, and 2%, respectively.

Data

Data on U.S. red wine imports from Italy, France, Spain, Australia, Chile, and the ROW were obtained from the U.S Department of Commerce-Bureau of the Census. Monthly import data from April 1990 to August 1999 of the "International Harmonized System of Commodity Classification" (HTSUSA) are used in this analysis, and the sample size includes 113 observations. The quantity of imports from each country is measured in liters, and the value of imports is defined as cost insurance freight (CIF) prices plus import taxes. Unit prices of imported red wine from each country are imputed or derived by dividing total value by total quantity of imports from the above data.

Domestic data were much more difficult to obtain. Although the U.S. Department of Agriculture (USDA) collects domestic data on grapes, it does not collect similar statistics for wine. As wine is an alcohol, wine data are collected by the Bureau of Alcohol, Tobacco and Firearms (ATF) within the U.S. Department of Treasury. Monthly domestic consumption of wine is calculated using data from ATF's "Monthly Statistical Reports—Wine," where domestic consumption equals taxable withdrawals (production) minus exports minus the change in stocks, and the monthly proportion of red wines consumed is calculated using consumption shares obtained from *The U.S. Wine Market* annual editions. For prices, a domestic red-wine producer-price index series was obtained from the Bureau of Labor Statistics (BLS), and it includes monthly price indices from December 1983 through December 2000. A price estimate for red wines in the year 2000 was obtained from Gomberg-Fredrikson based on AC Nielsen scanner data, which allows us to convert the BLS monthly price indices into actual prices. A summary of descriptive statistics is presented in Table 1.

AIDS and Time-Series Data

Deaton and Muelbauer developed a flexible-functional-form demand system from a PIGLOG expenditure or cost function. It is almost ideal, in their opinion, because it provides an arbitrary first-order approximation of any demand system, it provides perfect aggregation over consumers without maintaining homothetic preferences, its functional form is consistent with known household-budget data, it satisfies the axioms of choice exactly, it allows statistical testing of homogeneity and symmetry, and its linear approximation is simple to estimate (Deaton and Muelbauer). One of its serious limitations is that no restrictions on its parameters can insure negativity or concavity of the cost function.

In the Almost Ideal Demand System (AIDS), the log of a price index, derived analytically from the AIDS' cost function, deflates the log of (nominal) income. In practice, Deaton and Muelbauer recommend replacing the analytically derived log of the price index with an approximation, the Stone index, that is,

$$(1) \qquad \ln P^* = \sum_{i=1}^{n} w_i \ln p_i$$

where w_i $(= p_i q_i / x)$ is the budget share of good i (=1, …., n), $x = \sum_i p_i q_i$ is (nominal) expenditure or income, p_i is the price of good i, q_i is the quantity of good i, and n is the number of goods.

This approximation is not without its problems. Unless preferences are homothetic (unitary elasticities), the Stone index, even with constant prices, can vary because expenditure shares vary with income levels. Other deficiencies of this approximation exist including that simultaneity is introduced by the dependent variables entering into the Stone index (Eales and Unnevehr) and that parameter estimates are not invariant to the units of measure (Moshini).

Fortunately, in the time-series case, an alternative exists, the first-difference AIDS,[1] that is,

$$(2) \qquad dw_{it} = \beta_i d \ln\left(\frac{x_t}{P_t^{**}}\right) + \sum_j \gamma_{ij} d \ln p_{jt}$$

where d represents the first difference of a variable, dln P^{**} is the first difference of the analytically derived log-price index, t (=1,…,T) represents the time period, and β_i and γ_{ij} are parameters assumed to be constant. In principle, equation (2) is estimable, but, in practice, Deaton and Muelbauer suggest replacing dln P^{**} with either the first difference of the Stone index or the Divisia price index,

$$(3) \qquad d \ln P_t = \sum_{i=1}^{n} \overline{w}_{it} d \ln p_{it} \ ,$$

where $\overline{w}_{it} = (w_{it} + w_{i,t-1})/2$ is the average budget share between time t and t-1.[2] Unlike the Stone index, the Divisia price index in (3) does not vary with constant prices even if income changes and preferences are non-homothetic. Further, parameter estimates based on the first-difference version of AIDS utilizing the Divisia price index are invariant to

[1] Deaton and Muelbauer coined this name for the first differencing of the AIDS. The model itself without first differencing was named by them as the “levels version.” We adhere to this terminology throughout the paper.

[2] It is the usual practice in time-series demand studies to replace w_{it} with $\overline{w}_{it}$ (Theil).

units of measure, and there is no simultaneity problem as in the levels version. A careful reading of Deaton and Muelbauer suggests that the first-difference version fit their data better than the levels approximation in terms of homogeneity and symmetry restrictions.[3]

Empirical Estimation

The system we estimate is conditional on U.S. expenditure on red wines, both domestically produced and imported. To make the estimation manageable, we follow the multi-stage budgeting approach as described by Barten (1977) and used, for example, by Seale, Sparks, and Buxton. Specifically, we maintain that U.S. consumers allocate total expenditure among groups of goods, red wine being one of those groups.[4] Preferences among these groups are blockwise dependent (Theil) or weakly separable (Barten 1977). Having allocated expenditure for the group, red wines, U.S. consumers further allocate red-wine expenditure among U.S. domestically produced and imported red wines. By including U.S. red wines with imports, we follow the argument by Winters that domestic and imported goods of the same type may not be additively separable or preference independent.

Deaton and Muelbaurer fit the first-difference version of AIDS to British time-series data and introduced intercept terms to allow for time trends in the levels version. Similarly, we allow for monthly-time trends in the levels version by introducing 12 monthly-dummy variables in the difference version. Because the model is fit to monthly instead of annual data and to deseasonalize the data, we twelfth difference the data such that $d\ln z_t = \ln z_t - \ln z_{t-12}$ where z represents w, q, p, or x (Kmenta, pp. 325-326).

[3] Eales and Unnevehr and Kastens and Brester fit the first-difference version with the Divisia price index as suggested by Deaton and Muelbauer.
[4] One of the other groups could be white wines and, based on blockwise dependence, the marginal utility of consuming red wines would be affected by the consumption of white wines (Theil).

Specifically, we estimate the following twelfth-difference system, combining equations (2) and (3),

$$(4) \qquad dw_{it}^{*} = \beta_i d\ln\left(\frac{x_{gt}}{P_{gt}}\right) + \sum_{j\in g}\gamma_{ij} d\ln p_{jt} + \sum_t d_{it} + \varepsilon_{it}\,,$$

where g represent the group, in this case red wines, d represents a (twelfth) change in a variable, x_g $(=\sum_{i\in g} p_i q_i)$ is the expenditure on red wines from countries $i \in g$ (= Italy, France, Spain, Australia, Chile, U.S., or ROW), p_i and q_i are the price and quantity of red wines from countries i, $w_i^{*} = w_i / W_g$ is the (conditional) budget share of red wines from imported or domestic sources where w_i $(= p_i q_i / x)$ and $W_g = \sum_{i\in g} p_i q_i / x$, $d\ln P_t = \sum_{i\in g} \overline{w}_{it}^{*} d\ln p_{it}$ is the (conditional) Divisia price index where $\overline{w}_{it}^{*} = \left(w_{it}^{*} + w_{i,t-1}^{*}\right)/2$, d_{it} is a dummy variable for the i^{th} good in the t^{th} month, ε_i is the residual term for the i^{th} good with zero mean, and ε $(=\varepsilon_i, \ldots, \varepsilon_n)$ has covariance equal to Ω. If ν_t is the original error term for time t in the levels versions, then $\varepsilon_{it} = \nu_{it} - \nu_{i,t-12}$. If ν_t is autoregressive with ν_{t-12} (i.e., $\nu_{it=} \tau\, \nu_{i,t-12} + \zeta_{it}$), then ε_{it} is autoregressive (Kmenta, p. 321-322).

The adding-up conditions,

$$(5) \qquad \sum_i \beta_i = 0\,,\ \sum_i\sum_j d_{ij} = 0\,, \text{ and } \sum_i \gamma_{ij} = 0\,,$$

are met automatically because $\sum_i dw_i^{*} = 0$, and, accordingly, the full n x n covariance matrix is singular. Under this condition, Barten (1969) shows that the system parameters can be estimated by dropping one equation and that these estimates are invariant to the specific equation dropped; we drop the ROW equation for estimation purposes.

Homogeneity can be imposed on the system by constraining $\sum_{j \in g} \gamma_{ij} = 0$ and symmetry by constraining $\gamma_{ij} = \gamma_{ji} \quad \forall \quad i,j \in g$. Economic theory suggests that these conditions should be maintained, but it has become common practice to test for these restrictions (e.g., Deaton and Muelbauer; Laitinen; Meisner).

Testing Restrictions

The model of equation (4) is estimated in three ways: without homogeneity or symmetry (unrestricted), with homogeneity imposed, and with homogeneity and symmetry jointly imposed. We also test each of the above estimations for autocorrelation of the form $\varepsilon_{it} = \rho\varepsilon_{i,t-1} + \xi_{it}$ where ξ_{it} ~N(0,Σ_ξ). To preserve adding up, we constrain ρ to be the same across all equations (Berndt and Savin).

With autocorrelation, iterative seemingly unrelated regressions (SUR) is not maximum likelihood because the log of the Jacobian term of the likelihood function is not equal to zero. Accordingly, we fit the model to the monthly data under the different restrictions using the maximum-likelihood scoring method with and without autocorrelation of degree one (AR(1)). See Seale, Walker, and Kim for a thorough treatment and discussion of the scoring method when allowing for autoregressive errors of degree one.

Concentrated log-likelihood values from the different estimations are presented in Table 2. The values in the parentheses represent the number of restrictions going from the previous model to the next more restricted one. For example, imposing homogeneity on the unrestricted model necessitates 6 restrictions and imposing symmetry on the homogeneity-constrained model necessitates 15 further restrictions. Column (2) of table 2 presents the concentrated log-likelihood values when the autoregressive term, ρ, is constrained to equal zero, and column (3) of the table presents the concentrated log-likelihood values when ρ is not constrained to equal zero.

We use likelihood ratio (LR) tests to statistically determine whether or not the restrictions above should be imposed on the model. These LR tests are easily calculated based on the information reported in table 2. The test statistic is χ^2_q where q is the number of restrictions, and it is equal to the negative of $2(L_r\text{-}L_u)$ where L_r and L_u are the log-likelihood values for the restricted and unrestricted models, respectively (Kmenta, p. 491).[5] For example, a test of AR(1) versus no AR(1) is χ^2_1 that, at the 95% confidence

Table 2. Concentrated Log-likelihood Values from Twelfth-Difference AIDS under Different Restrictions for Monthly Data, U.S. Red-Wine Demand.

Restrictions (1)	No AR(1)[a] (2)	AR(1) (3)
unrestricted	2701.04	2705.00
homogeneity	2694.13 (6)[b]	2700.16 (6)
homogeneity and symmetry, unrestricted monthly dummies	2684.83 (15)	2690.91 (15)
homogeneity and symmetry, restricted monthly dummies	2613.36 (60)	2630.37 (60)

[a]AR(1) respresents autocorrelation of degree one.
[b]Numbers in parentheses represent the number of restrictions from previous model.

level, has a critical value of 3.84 (Johnston, p. 427). When making row comparisons between the values in column (2) and (3), the constraint that ρ = 0 is rejected in all cases.

[5] The test is invariant as to whether one uses the log value of the concentrated-likelihood or the unconcentrated-likelihood functions.

Accordingly, correcting for AR(1) is necessary to ensure that the asymptotic standard errors are consistent.[6]

Concentrating on the model corrected for AR(1), homogeneity is not rejected at the 95% confidence level by the LR test when comparing the concentrated log-likelihood values of the homogeneity-constrained model and the unrestricted one. To test for symmetry, we compare the concentrated log-likelihood values of the symmetry-and-homogeneity constrained model with that of the homogeneity-constrained one. Again, we do not reject symmetry at the 95% confidence level. We also test whether it is appropriate to constrain the monthly dummy parameters to be the same in each equation and reject these restrictions at the 95% confidence level.

Parameter estimates

Based on the LR test results above, we do not reject homogeneity, symmetry, or AR(1) although we do reject the hypothesis that the monthly-dummy parameters are the same in each equation. Accordingly, all reported parameter and elasticity estimates are based on the AR(1) corrected model with homogeneity and symmetry imposed and allowing the monthly-dummy parameters to differ among equations. Estimated expenditure parameters, price parameters, the AR(1) parameter, and their associated asymptotic standard errors (within parentheses) are reported in Table 3. We estimate the model with maximum likelihood using the scoring method dropping the ROW equation due to the singularity of the full-covariance matrix. The parameter estimates of the ROW equation and their associated asymptotic standard errors are calculated based on the

[6] Without correcting for AR(1), the asymptotic standard errors are inconsistent, and any statistical tests based upon them are invalid.

adding-up restrictions. Parameter estimates are used to calculate elasticity measures as discussed below and presented in tables 4 and 5.

All expenditure parameters, β_i, are significant at the 95% confidence level and are reported in column (9) of table 3. The β_i for imported red wines are all negative while that of the U.S. is positive. As we shall see below, this means that imported red wines are conditionally expenditure inelastic while U.S. domestic red wines are conditionally expenditure elastic; none are conditionally unitary elastic.

Conditional own-price parameters are reported along the diagonal of columns (2)-(8) of table 3. Four (Italy, France, Chile, and U.S.) are statistically different from zero at the 95% confidence level, two (Spain and Australia) are statistically different from zero at the 90% confidence level, and that of ROW is statistically zero. All are positive except that of Chilean red wines.

Conditional cross-price parameters are reported as the non-diagonal elements of columns (2)-(8) of table 3. Of the 21 conditional cross-price parameters, seven are statistically different from zero at the 95% confidence level, and five others are statistically different from zero at the 90% confidence level. Twelve are negative while the others are positive. It is also of interest that all price parameters of U.S. domestic red wines are significantly different from zero at the 95% or 90% confidence levels. The autocorrelation parameter, ρ, is equal to .17 with an asymptotic standard error of .04 indicating that the AR(1) corrected model is the appropriate specification.

Table 3. Conditional Parameter Estimates from Twelfth-Difference AIDS for U.S. Red Wine Demand, Monthly Data, 1990:4-1999:8, AR(1)[a] Corrected.

	parameters							
	γ_{ij}							β_i
Countries (1)	Italy (2)	France (3)	Spain (4)	Australia (5)	Chile (6)	U.S.[b] (7)	ROW[c] (8)	Expenditure (9)
Italy	.088	-.044	.006	.000	.004	-.060	.005	-.054
	(.017)[d]	(.012)	(.003)	(.005)	(.006)	(.025)	(.003)	(.012)
France		.107	-.004	-.002	-.005	-.049	-.003	-.056
		(.017)	(.002)	(.003)	(.004)	(.027)	(.003)	(.017)
Spain			.003	.001	.000	-.004	-.002	-.007
			(.002)	(.002)	(.003)	(.004)	(.003)	(.002)
Australia				.009	.004	-.013	.001	-.024
				(.005)	(.005)	(.007)	(.003)	(.003)
Chile					-.021	.019	-.002	-.021
					(.008)	(.009)	(.003)	(.004)
U.S.						.112	-.004	.172
						(.055)	(.002)	(.030)
ROW							.005	-.010
							(.007)	(.003)

[a]AR(1) is autocorrelation of degree one. The estimated estimate of this parameter, ρ, is .17 with an asymptotic standard error of .04.
[b]U.S. represents United States.
[c]ROW represents rest-of-world.
[d] Asymptotic standard errors are in parentheses.

Conditional Expenditure Elasticities and Marginal Shares

Conditional expenditure elasticities are calculated as $\eta_i = 1 + \beta_i / w_i^*$ (Chalfant; Alston, Foster and Green). If β_i is significantly different from zero and positive (negative), the conditional expenditure elasticity is elastic (inelastic); if β_i equals zero, the conditional expenditure is unitary. These conditional elasticities estimate the percent change in quantity demanded for red wines when total U.S. expenditure on red wines increases by 1%.

Conditional expenditure elasticities, reported in column (2) of table 4, are calculated base on the average (conditional) expenditure shares for the red wines in 1999. All are positive but only that of the U.S. is greater than one (1.3); those of the imported red wines are .4 (Chile and ROW), .5 (Italy and Australia), and .6 (France and Spain). These results indicate that if total U.S. expenditure on red wines increase 1%, the quantity demanded for U.S. red wines will increase more than 1% while that of imported country-specific red wines will increase less than 1%.

Marginal shares equal $\beta_i + w_i^*$ and are calculated based on the average (conditional) expenditure shares for the red wines in 1999 and are reported in column (3) of table 4. Marginal shares indicate how an additional dollar spent on red wines would be allocated. U.S. domestic wines would benefit the most with 80 cents of the additional dollar being spent on them. Eight cents and six cents of the additional dollar would be allocated to French and Italian wines, respectively, while two cents of the additional dollar would be spent on Australian red wines and only a penny each on Spanish, Chilean, and ROW red wines. Thus, the U.S. red wine industry would be by far the biggest gainer if an additional dollar were spent on red wines.

Table 4. Conditional Expenditure Elasticities, Marginal Shares, and Own-Price Elasticities for Imported and Domestic Red Wines in the United States in 1999.

Source country	Expenditure elasticities (η_i)	Marginal shares	Own-price elasticities	
			Slutsky	Cournot
(1)	(2)	(3)	(4)	(5)
Italy	.52	.06	-.10	-.27
France	.60	.08	-.09	-.31
Spain	.60	.01	-.78	-.80
Australia	.50	.02	-.77	-.84
Chile	.40	.01	-1.56	-1.61
United States	1.27	.80	-.19	-1.63
Rest of world	.44	.01	-.69	-.72

Conditional Own-Price Elasticities

Several price elasticity formulas have appeared in the literature for the AIDS (Alston, Foster and Green), and we use those suggested by Chalfant. The conditional Slutsky own-price elasticity of demand is

(6) $$S_{ii} = -1 + \frac{\gamma_{ii}}{w_i^*} + w_i^*$$

while that of the conditional Cournot own-price elasticity of demand is

(7) $$C_{ii} = -1 + \frac{\gamma_{ii}}{w_i^*} + \beta_i \ .$$

Conditional Slutsky and Cournot own-price elasticities are calculated based on the average (conditional) expenditure shares for the red wines in 1999 and are reported in columns (4) and (5) of table 4, respectively. Slutsky price elasticities are compensated while Cournot price elasticities are uncompensated. According, Cournot own-price elasticities are larger in absolute value than Slutsky ones because Cournot elasticities measure pure substitution effects (as do Slutsky elasticities) plus income effects of price changes.

All conditional own-price elasticities are negative. Slutsky and Cournot own-price elasticities of demand for Chilean red wine are conditionally elastic while the conditional Slutsky (Cournot) own-price elasticity of demand for U.S. red wines is highly inelastic (elastic); the conditional Slutsky (Cournot) own-price elasticity of demand for U.S. red wines is -.2 (-1.6) and that for Chilean wines –1.6 (-1.6). The difference in the U.S. conditional Slutsky and Cournot own-price elasticities reflects the expenditure sensitivity of demand for U.S. domestic red wines.

Italian and French wines are much less own-price elastic. This indicates that U.S. consumers' demand for Italian and French wines are not that sensitive to own-price changes. Conditional Slutsky (Cournot) own-price elasticities are only -.1 (-.3) for French and Italian red wines. The conditional own-price elasticities of the other wines are more elastic than French and Italian red wines but less elastic than Chilean red wines. Australian red wines have a conditional Slutsky (Cournot) own-price elasticity of -.77 (-.84), Spanish red wines of -.78 (-.80) , and that of the ROW red wine is -.69 (-.72).

These results suggest significantly different effects from price changes on U.S. red wine demand. If the price of French and Italian red wines increase by 1%, uncompensated quantity demanded for these wines will decrease only about .3%. U.S. and Chilean red wines are much more affected by uncompensated own-price changes. A

1% increase in U.S. and Chilean red wine own-prices would decrease the uncompensated quantity demanded for these wines by about 1.6%.

Conditional cross-price elasticities

Conditional cross-price elasticities measure the effect on the quantity demanded of a good when the price of a substitute or complimentary good changes. Again, we choose the formulas for calculation as suggested by Chalfont. Specifically, the Slutsky cross-price elasticity is

$$(8) \qquad S_{ij} = \frac{\gamma_{ij}}{w_i^*} + w_j^* \qquad\qquad i \neq j$$

while that of the conditional Cournot cross-price elasticity of demand is

$$(9) \qquad C_{ij} = \frac{\gamma_{ij}}{w_i^*} - \beta_i \frac{w_j^*}{w_i^*}, \quad i \neq j\,.$$

These elasticities are calculated based on the average (conditional) expenditure shares for the red wines in 1999 and are reported in table 5. The conditional Slutsky cross-price elasticities are reported in the top matrix of the table, and the conditional Cournot cross-price elasticities in the bottom matrix. A positive Slutsky cross-price elasticity indicates that an increase (decrease) in the i^{th} good's price will cause the quantity demanded of the j^{th} good to increase (decrease), that is, the goods are substitutes (complements). Of the 42 conditional Slutsky cross-price elasticities, 33 are positive, and nine are negative.

Australian and U.S. red wines are the only ones that are substitutes for all wines. A compensated price change in U.S. red wines has a small positive effect on the quantity

Table 5. Conditional Cross-Price Elasticities of United States Demand for Red Wine by Source in 1999.

Source Country (1)	Italy (2)	France (3)	Spain (4)	Australia (5)	Chile (6)	U.S. [a] (7)	ROW[b] (8)
				Slutsky			
Italy	---	-.25	.07	.05	.07	.10	.06
France	-.20	---	-.01	.03	.00	.27	.00
Spain	.45	-.11	---	.10	.05	.04	-.13
Australia	.12	.10	.04	---	.12	.36	.04
Chile	.22	.00	.03	.17	---	1.17	-.03
U.S.	.02	.06	.01	.03	.06	---	.01
ROW	.41	-.03	-.12	.10	-.06	.40	---
				Cournot			
Italy	---	-.46	.04	-.02	.02	-.86	.04
France	-.38	---	-.04	-.04	-.06	-.73	-.03
Spain	.27	-.33	---	.03	.00	-.60	-.16
Australia	-.05	-.11	.01	---	.07	-.59	.01
Chile	.06	-.19	.00	.11	---	.28	-.06
U.S.	-.24	-.26	-.03	-.08	-.01	--	-.03
ROW	.24	-.23	-.15	.03	-.11	-.51	---

[a]U.S. represents United States.
[b]ROW represents rest-of-world.

demanded for all other red wines; a 1% increase in the price of U.S. domestic red wines will increase the (compensated) quantity demanded for the other wines by less than .1%. However, the price changes of the imported red wines have more positive effects on the quantity demanded of U.S. domestic red wines. This is particularly true for a price increase of Chilean red wines. A 1% increase in the price of Chilean red wines will increase the (compensated) quantity demanded for U.S. red wines by approximately 1.2%. A 1% increase in the price of Australian and ROW red wines will increase the (compensated) quantity demanded for U.S. red wines by approximately .4% while a 1% increase in the prices of French, Spanish, and Italian red wines will increase the (compensated) quantity demanded of U.S. red wines by approximately .3%, .04%, and

.1%, respectively. Another interesting finding is that French and Italian red wines are complementary as are French and Spanish red wines.

One of the more interesting findings concerning the Cournot cross-price elasticities and reported in the bottom matrix of table 5 is that when the income effect of a price change is taken into account in addition to the pure substitution effect, an increase in the U.S. price of domestic red wines decreases the demand for all other red wines. Additionally, except in the case of Chilean red wines, increases in the prices of imported red wines decrease the quantity demanded for U.S. domestic red wines. For example, a 1% increase in the price of Italian (French) red wines will decrease the (uncompensated) quantity demanded of U.S. domestic red wines by approximately .9% (.7%), respectively. A 1% increase in the price of Spanish or Australian red wines will decrease the (uncompensated) quantity demanded for U.S. domestic red wines by approximately .6%, while a 1% increase in the price of ROW red wines will decrease the quantity demanded for U.S. domestic red wines by about .5%. In the case of Chilean red wines, a 1% increase in price will *increase* the (uncompensated) quantity demanded of U.S. domestic red wines about .3%, indicating that U.S. consumers would prefer to switch to domestically produced red wines rather than pay higher prices for Chilean red wines.

Conclusions

Key finding of this research include the following. Firstly, although the volume of imports has dramatically increased in the past decade, U.S. consumption of domestically produced red wines far exceeds that of imports. Secondly, expenditure elasticities of U.S. imported red wines are all inelastic. In contrast to this, the expenditure elasticity of demand for U.S. domestic red wines is elastic. This result implies that if U.S. total expenditure on red wines were to increase, U.S. demand for domestic red wines would increase more than the demand for red-wine imports. This would certainly benefit the U.S. red wine industry. However, increases in total U.S. expenditures for red wines stem from non-price demand determinants such as rising incomes, changing demographics, or changes in preferences and may be beyond the influence of the U.S. wine industry. One possible avenue of influence may be through specific label or generic advertising.

Thirdly, the conditional own-price Cournot elasticities of U.S. and Chilean red wines are elastic but inelastic for all other country wines. Thus, if U.S. producers increase the price of their red wines, U.S. consumers are price sensitive (when income effects are considered) and will decrease their consumption of domestic red wines by a greater percent than the price increase. Because of this, total revenue to U.S. red-wine producers will fall if price rises, and this would be harmful to the U.S. red-wine industry.

Fourthly, the conditional own-price elasticities (Cournot and Slutsky) for Italian and French red wines are strongly inelastic relative to Spanish and Australian red wines. This combined with the fact that Italian and French red wines dominate U.S. red-wine imports, comprising 78% of the total red-wine imports, gives red-wine producers from these countries a competitive advantage. Thus, Italian and French red-wine producers could increase their prices with little impact on the quantity demanded for their wines and thereby increase their revenue.

Fifthly, regarding conditional Slutsky (compensated) cross-price elasticities, U.S. and Australian wines are substitutes for red wines from other countries. Additionally, a 1% increase in the price of U.S. red wines has little (compensated) effect on the demand for red wines. However, a 1% increase in the price of imported red wines does increase the demand for domestically produced red wine, with Chilean wines having the most dramatic impact.

Finally, when comparing own- and cross-price elasticities, results indicate that an increase in the price of U.S. red wines will reduce compensated demand by only .19 (Slutsky), but, when income effects are considered (Cournot), a 1% increase in the price of U.S. red wines leads to a 1.6% reduction in its quantity demanded. This decline is six times greater than the decline for Italian and French red wines and over 20 times greater than that of any other country-specific wines.

The bottom-line implications are that U.S. red-wine producers are able to increase their total revenues by decreasing price because their own-price elasticity of demand (Cournot) is elastic. In contrast, Italian and French red-wine producers could increase their total revenue by increasing prices because of the highly inelastic nature of demand by U.S. consumers for their red wines.

References

Alston, J. M., K. A. Foster and R. D. Green. "Estimating Elasticities with the Linear Approximate Almost Ideal Demand System: Some Monte Carlo Results." *Review of Economics and Statistics* 76(1994):351-356.

Barten, A. P. "Maximum Likelihood Estimation of a Complete System of Demand Equations." *European Economic Review* 1(1969):7-73.

________. "The Systems of Consumer Demand Functions Approach: A Review." *Econometrica* 45(1977):23-51.

Berndt, E. R. and N. E. Savin. "Estimation and Hypothesis Testing in Singular Equation Systems with Autoregressive Disturbances." *Econometrica* 43(1975):937-957.

Chalfant, J. "A Globally Flexible, Almost Ideal Demand System." *Journal of Business and Economic Statistics* 5(1987):233-241.

Deaton, A., and J. Muellbauer. "An Almost Ideal Demand System." *The American Economic Review* 70:3(1980):312-326.

Eales, J.S., and L.J. Unnevehr. "Demand for Beef and Chicken Products: Separability and Structural Change." *American Journal of Agricultural Economics* 70(1988):521-532.

Food and Agriculture Organization (FAO) of the United Nations. "FAOSTAT Statistical Database," Rome, Italy, 1998 at http://apps.fao.org/lim500/nph-wrap.pl?Trade CropsLivestockProducts&Domain=SUA&servlet=1.

Gomberg-Fredrikson—A consulting firm specializing in wine industry economics. Personal correspondence with Eileen Fredrikson. Woodside, California.

Kastens, T. L., and G. W. Brester. "Model Selection and Forecasting Ability of Theory-Constrained Food Demand Systems." *American Journal of Agricultural Economics* 78(1996): 301-312.

Kmenta, J. *Elements of Econometrics, 2nd edition*. New York City, NY: Macmillan Publishing Co., 1986.

Johnston, J. *Econometric Methods*. New York City, NY: McGraw-Hill Book Co., 1972.

Laitinen, K. "Why is Demand Homogeneity So Often Rejected?" *Economics Letters* 1(1978):187-191.

M. Shanken Communications. *The U.S. Wine Market.* New York City, NY: M. Shanken Communications, various annual editions.

Meisner, J. F. "The Sad Fate of the Asymptotic Slutsky Symmetry Testing for Large Systems." *Economics Letters* 2(1979):187-191.

Moschini, G. "Units of Measurement and the Stone Index in Demand System Estimation." *American Journal of Agricultural Economics* 77(1995):63-68.

Seale, J, L., Jr., W. E. Walker, and I. M. Kim. "The Demand for Energy: Cross-Country Evidence Using the Florida Model." *Energy Economics* 13(1991):33-40.

Seale, J.L., Jr., A. L. Sparks, and B. M. Buxton. "A Rotterdam Application to International Trade in Fresh Apples: A Differential Approach." *Journal of Agricultural and Resource Economics.* 17:1(1992):138-149.

Stone, R. "Linear Expenditure Systems and Demand Analysis: An Application to the Pattern of British Demand." *Economic Journal* 64(1954):511-527.

Theil, H. *The System-Wide Approach to Microeconomics*. Chicago, IL: University of Chicago Press, 1980.

U.S. Department of Agriculture, Economic Research Service. *Foreign Agricultural Trade of the United States (USDA-FATUS).* 1999 at http://www.econ.ag.gov/db/FATUS/index.asp?Type=MS.

U.S. Department of Agriculture, Foreign Agricultural Service (USDA-FAS). "FAS Statistics--World Horticultural Trade and U.S. Export Opportunities," 1998 at http://www.fas.usda.gov/htp2/circular/1998/ 98-08/circular.htm.

U.S. Department of Commerce-Bureau of the Census. "U.S. Imports of Merchandise." 1999 (CD-ROM) Washington, DC. C3.278/2: 990/4-999/8.

U.S. Department of Labor-Bureau of Labor Statistics (BLS). "Producer Price Index—Red grape wine with 14 percent or less alcohol content; series id: PCU2084#814" at http://146.142.4.24/cgi-bin/srgate

U.S. Department of Treasury-Bureau of Alcohol, Tobacco and Firearms (ATF). "Monthly Statistical Report—Wine." Washington, DC. January 1990 – December 1999.

Wine Institute. "Wine Institute Statistics--Economic and Industry Background," 1999 at http://www.wineinstitute.org/communications/statistics/index.htm.

Winters, L. A. "Separability and the Specification of Foreign Trade Functions." *Journal of International Economics* 17(1984):239-263.

Table 1. Descriptive Statistics on U.S. Consumption of Red Wine by Source, April 1990 - August 1999.

	Italy	**France**	**Spain**	**Australia**	**Chile**	*Rest of Wor*
Price (USD per Liter)						
Mean	$3.43	$6.49	$4.38	$4.88	$2.48	$2.77
median	3.35	6.31	4.26	4.72	2.46	2.73
minimum	2.76	4.74	2.58	3.98	2.06	1.93
maximum	4.88	10.89	6.39	7.26	3.10	4.60
coefficient of variation	12.10	16.64	18.22	12.88	9.01	18.34
Import Values (000 USD)						
Mean	$14,150	$20,021	$1,813	$3,254	$3,603	$2,103
median	12,652	17,629	1,561	2,088	2,767	1,903
minimum	5,312	6,441	338	281	442	326
maximum	32,891	49,705	5,000	12,994	8,752	4,707
coefficient of variation	45.04	44.31	59.24	92.55	69.39	50.32
Import Quantity (Hectoliters)						
Mean	40,288	30,719	3,939	6,268	14,214	7,510
median	38,921	28,237	3,593	4,920	11,676	6,661
minimum	16,723	12,438	1,036	455	1,932	1,382
maximum	80,182	73,624	10,312	21,705	35,556	15,657
coefficient of variation	36.60	40.03	46.34	82.37	65.17	44.54
Expenditure Shares						
Mean	0.1185	0.1684	0.0143	0.0231	0.0265	0.0172
median	0.1100	0.1600	0.0100	0.0200	0.0300	0.0200
minimum	0.0700	0.1000	0.0100	0.0000	0.0100	0.0100
maximum	0.2200	0.2800	0.0400	0.0700	0.0700	0.0300
coefficient of variation	27.16	25.42	39.62	64.40	46.30	29.39