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DO FLUCTUATIONS IN WINE STOCKS AFFECT WINE PRICES?

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Abstract: Globalization and the expansion of world wine trade have caused a wine boom that together with agricultural subsidies have made fluctuations in wine inventories a more critical issue. In the case of domestic and international wine markets, little is known about intertemporal inventory adjustments and how they relate to prices. We investigate possible dynamic relations between these variables in a time series context, so as to better understand how wine producers and traders can face growing price and financial volatility. Countries for whom meaningful data series could be constructed include: Argentina, Australia, France, Germany, Italy, Spain and the United States. The study begins by examining the empirical evidence on inventories in these markets and their relation to prices. Stationarity tests are first performed to assess likely trends in the wine inventory and price variables. Cointegration analysis follows to analyze the stationary relationships between these variables. To explain the dynamics of this relationship, vector autoregressions have been estimated and impulse functions are computed to measure possible delays between variable reactions.

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1. INTRODUCTION

Globalization and the expansion of world wine trade have caused a wine boom that together with agricultural subsidies have made fluctuations in wine inventories a more critical issue. In the case of domestic and international wine markets, little is known about intertemporal inventory adjustments and how they relate to prices. Data on wine inventories and prices have been very difficult to obtain; and this problem has been aggravated by the fact that wines are such hetereogenous commodities. Yet we do know that wine inventories increase during times of abundant grape harvests and decline during years of poor grape harvests. And these conditions ultimately do affect wine prices. There is thus a need to investigate possible dynamic relations between these variables in a time series context, so as to better understand how wine producers and traders can face growing price and financial volatility. We perform this analysis for a group of countries where meaningful data series could be constructed: Argentina, Australia, France, Germany, Italy, Spain and the United States.

The study begins by examining the empirical evidence on inventories in these markets and their relation to prices. Stationarity tests are first performed to assess likely trends in the wine inventory and price variables. Cointegration analysis follows to analyze the stationary relationships between these variables. To explain the dynamics of the interrelationship between these variables, vector autoregressions have been estimated and impulse functions are computed to measure possible delays between variable reactions.

The remainder of the paper consists of the following parts: (2) Background, (3) Inventory and Price Behavior, (4) Testing for Trends, (5) Cointegration between Stocks and Prices, (6) Vector Autoregression and Impulse Results, and (7) Conclusions.

2. BACKGROUND

The observation that the role of commodity stocks is little understood is surprising. In the case of wine markets, this topic has hardly been researched at all. Most often commodity inventory and price behavior have been considered as an intertemporal adjustment process reflecting demand and supply disequilibrium in a closed market or economy. Inventory adjustments as such can influence domestic and international price fluctuations and their understanding is important for agricultural commodity producers and consumers. Since demand (particularly for agricultural food and beverages) and supply (particularly for agricultural perennial crops) tend to be relatively price-inelastic in the short run, inventory movements, which are more price elastic, provide the vehicle whereby markets achieve equilibrium. Nonetheless, stock behavior should not be explained just in terms of residual adjustments or unintended accumulations. Stock holding can also take place for precautionary, transactions, and speculative purposes. Other intervening influences also exist. Stock changes not only reflect agricultural surpluses and deficits due to climatic changes, but exogenous business cycle effects as well, e.g. for the case of wine see Labys (2001). Also of some concern are uncertain flows of information, income and financial fluctuations, market persistence to shocks, and sudden trade flow disruptions. Such factors together with more complex behavioral motives exist in the inventory-theoretic literature, as briefly summarized below.

The most basic aspect of commodity stockholding behavior is that it represents intertemporal arbitrage, initially in a closed system. Earlier studies of this behavior appear in Labys (1973); later reviews include Antonini (1988), Blinder and Maccini (1991), Chikan (1984), Labys (1989), Williams and Wright (1991), and Wright and Williams (1982). Among these studies, the popular supply of storage model is based on the premise that each firm will adjust its inventory level until the marginal revenue of holding stocks equals it marginal costs. As developed by Working (1949), Brennan (1958, 1959) and Weymar (1969a and b), the motives for inventory holding include convenience yield, stockout yield, and coverage yield. Intertwined in this process is the role of inventories as the allocating agent, particularly to avoid stockout and to facilitate the scheduling of production and sales.

Research in this area expanded in the 1980's with consideration of the importance of disequilibrium adjustments and rational expectations in the work of Kawai (1983) and Otani (1983). Attention to problems of speculative demand and convenience yield appear in Newberry and Stiglitz (1982), Gilbert (1991), and Larson (1994). Of special interest in analyzing convenience yield is the test of speculative carry over. Also of significance are theories that consider inventories in the context of portfolio theory of asset-holding suggested earlier by Yver (1971). While Yver applied that theory to cattle inventories, Orden (1982) later developed an asset theory generally relevant for agricultural commodities. The production cost-smoothing model, whereby inventories are used to shift output to periods in which production costs are low also can be used to avoid stockouts and to reduce scheduling costs. Eichenbaum (1984, 1989) and Eckstein and Eichenbaum (1985) analyze a target level of inventories and the linear-quadratic cost of deviating from that level. Further tests of this model were made by Labys and Lord (1992) using error-correction model (ECM) analysis for several of the major traded agricultural commodities.

Agricultural producers have long sought for an inventory-holding model that would optimize the level of inventories to be held, given other influencing factors. Gilbert (1991), Knapp (1982), Pindyck (1994) and others have suggested an inventory theory based on intertemporal optimization. The few studies available that address inventory holdings in the commodity specific case of wine employ inventory variables either as measure of disequilibrium market adjustment present in the price equation (as is the present study) and/or as variables that define the market closing of a system. Further insights can be obtained from Amspacher (1988) and Gijsbers and Labys (1988). Wohlgenant (1978,1982), in particular, developed a dynamic model of wine processor behavior in which price, production and input demand functions are derived from an optimal control model that takes into account inventory growth from ageing and the linkage between product inventories and input purchases.

3. INVENTORY AND PRICE BEHAVIOR

Wine researchers have neglected the problems of wine market disequilibrium for some time, though market instability problems have long been recognized by policy makers. This major international wine problem has been exacerbated because of growing wine trade and excess of world wine production relative to declining consumption (i.e. Labys and Cohen, 2006). The greatest impact of the disequilibrium between production and consumption appears to be in the European Union (EU), because it has proven difficult politically to modify its Common Agricultural Policy. Rather large surpluses have created particular problems, such as the need to distill considerable volumes of wine every year or to export wines below cost. In other parts of the world, the production and consumption imbalances appear to have declined to be closer to what could be considered stable market equilibrium. While vineyard and wine management programs have resulted in severe reduction in the surfaces planted, restructuring and replanting activities have resulted in increases in productivity. The general trend is in the direction of expanding the production of wine of better quality, although table wine production still remains relatively large.

To be able to study the nature of stock and price movements and their interactions in wine markets, we have undertaken to construct national wine inventory and price series beginning 1975 for France and the United States, and 1986 for Argentina. Australia, Germany, Italy, and Spain. All series end in 2003. Our series result from individual year by year compilation and estimation, based on inventory and price data reported to trade and governmental organizations. Sources of the data include historical documents from the Office of International Wines and Vines (OIV) in Paris, the Food and Agricultural Organization in Rome (FAO), along with the world wine data published by Wittwer and Rothfield (2006), Anderson and Norman (2003) and Berger, Anderson and Stringer (1998). The price data have the same origins, but the longer French series come from INSEE historical data and the longer US series stem from US Department of Labor data archives. Even with this care, the inventory data represent only bulk wine holdings by private and commercial holders, and in the case of France wine producer and wholesaler stocks. The US stocks for the most part are those held in California. No data are available by quality designation. Similarly the price data are in the form of aggregate indexes and do not include individual wine types. Definitions and sources of the variables appear in Appendix A.

Figure 1 demonstrates changes in the trends and fluctations in inventory holdings for our sample group of countries. Beginning 1975, French wine stocks have fluctuated about an average of 394 million litres (Ml) annually, with a sharp dip occurring between 1987 and 1991. US wine stocks have fluctuated less, around an average of 182 Ml. Beginning 1986, Argentinian stocks have declined continuously around an average of 246 Ml. Australian wine stocks have risen about a smaller average of 78 Ml. Germany, Italy and Spain stocks also show fluctuations over these years, around 171 Ml, 281 Ml, and 26 Ml, respectively.

Figure 1. Wine Inventory Fluctuations: 1975-2003

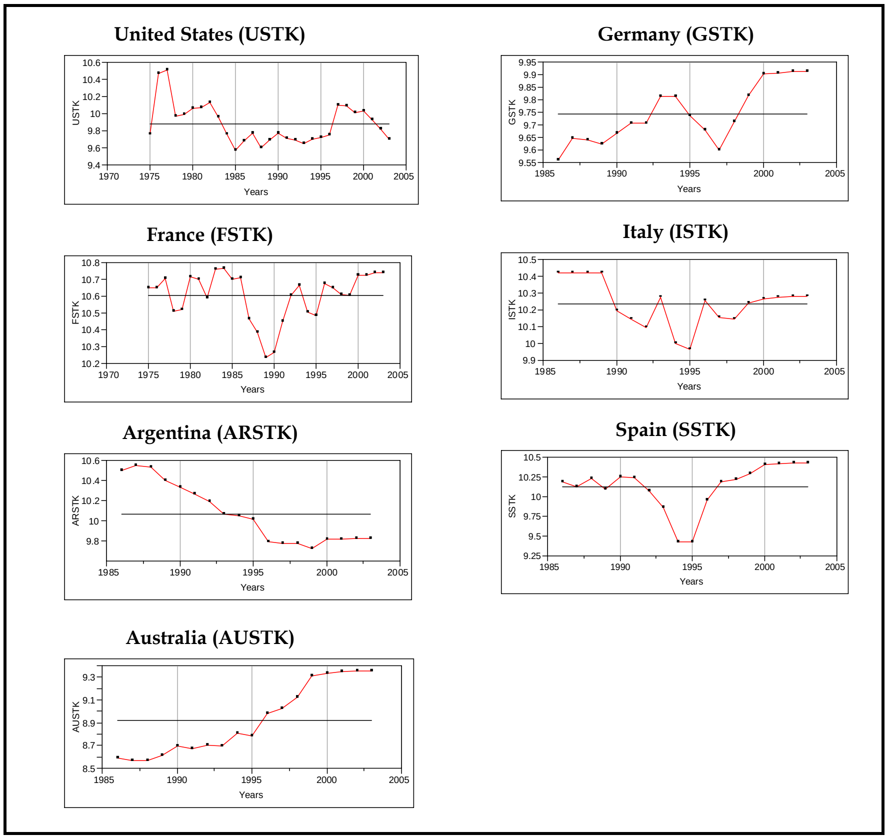
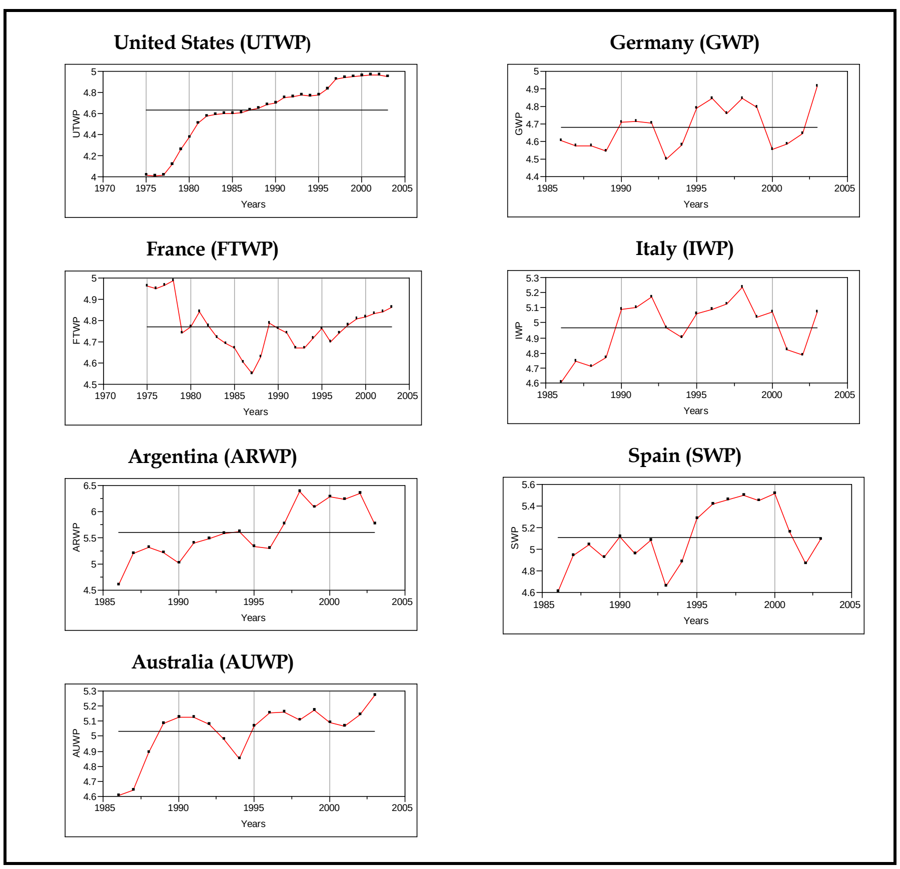


Table wine prices reflect changes in production, consumption and their differences or stock changes. Stocks themselves embody the history of past market imbalances. When table wine demand is growing relative to production and stocks, prices will rise accordingly. When fluctuations in weather conditions cause wine production to decrease or increase, prices will also adjust. Other factors also influence table wine prices such as price supports (the Common Agricultural Policy), rates of inflation or even interest rates. Quality wine prices, in contrast, are determined in other markets and reflect other price-making influences. Because the disparities between wine production and the actual grape varieties vary appreciably from country to country, it is difficult to determine how wine prices move in the face of wine market disequilibria.

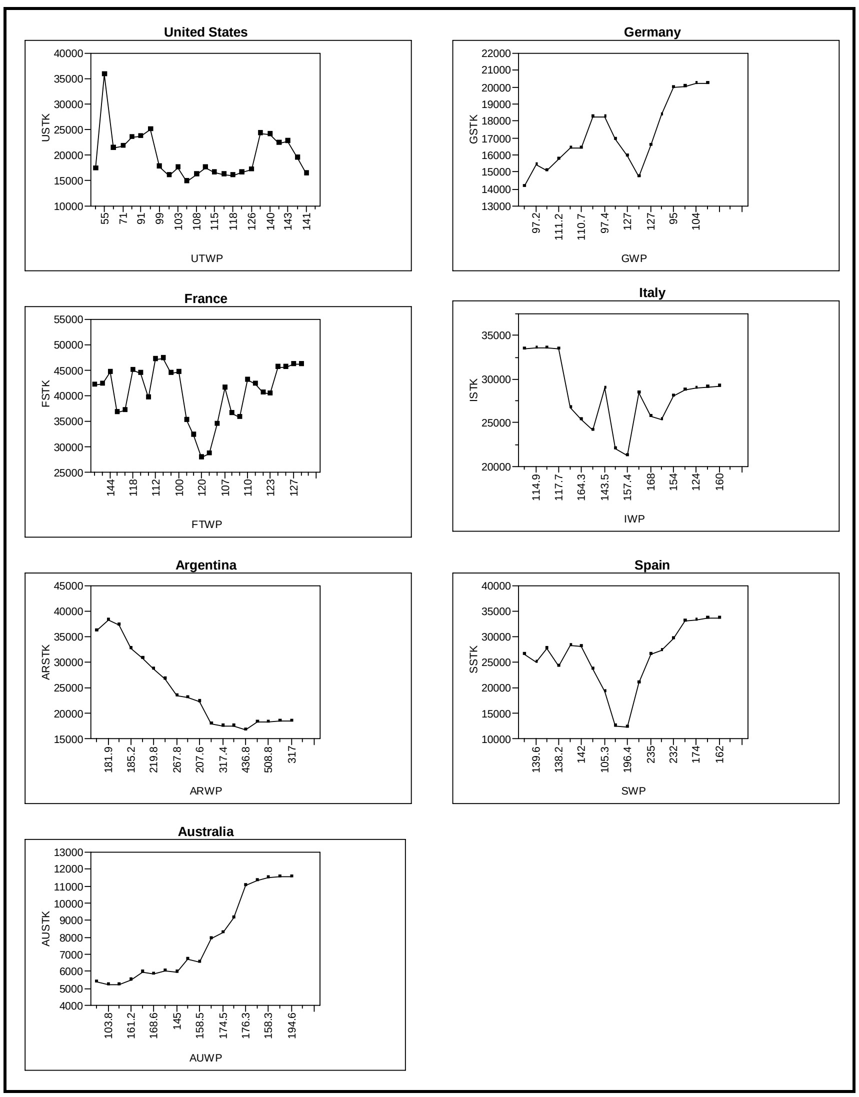
Figure 2 features the results of compiling the wine price indexes. Since 1975, French wine prices have fluctuated considerably around a slightly falling trend, reflecting the impact of unstable wine production relative to a fairly constant, but steadily declining, demand. Relative wine surpluses in the United States have been reduced, leading to increases in domestic table wine prices. Since 1986, Australian wine prices also have continued to rise. The prices for Germany, Italy and Spain are shown to fluctuate over the same period. The lower-quantity vintages in 1988-1989 for many major European countries, the strengthening of market interventions to remove surpluses from the market, and increases in EU allocations to structural programs to reduce wine acreage have caused wine prices to recover in Europe. Some of the price increases shown around 1995 are due to the relatively smaller harvests and vintages.

Figure 2. Wine Price Fluctuations: 1975-2003



In terms of interactions between wine inventories and prices, Figure 3 provides the graphs of these series against each other over the relevant time periods. The individual country plots are not easy to interpret because of the mentioned intervening factors that confound the normal relationship between these variables. There is also the confusion of defining exactly what the inventory data mean. Stocks can be held by (1) producers in the form of carryover, (2) consumers such as restaurants awaiting resale or wine connoisseurs for storage, and (3) traders or speculators who are anticipating resale to others.

Figure 3. Wine Inventory versus Price Fluctuations: 1970-2003



Carryover stocks can be considered a supply of stocks in which more stocks are held as prices rise, in anticipation of further price increases. The alternative approach considers stockholding in the form of a demand; more stocks are purchased when prices are low and fewer stocks are held when prices are higher. Where more detailed data exist, then one can employ the more complex theories of stockholding mentioned earlier. Returning to Figure 3, only Argentina, Italy and the United Sates appear to have a stock-price relationship roughly of a demand nature. The curves of Germany, France, Australia and Spain appear to reflect carryover.

4. TESTING FOR TRENDS

Examination of the autocorrelation and partial autocorrelation coefficients for the inventory and prices series has suggested that the series are dissimilar regarding persistency and mostly have unstable trends. With the exception of Australia, Germany, Italy and Spain series, the graphs of the autocorrelation functions decay very slowly, which might suggest possible unit roots for the series. Trend tests thus follow.

Unit Root Tests

The results of performing the ADF test in levels under the no-constant-no-trend specifications suggest that the null hypothesis of non-stationarity cannot be rejected at the 5 percent significance level (Table 1). Therefore, we conclude that both wine prices and wine inventories are non-stationary in their levels at the 5 percent significance level. The series were differenced and the ADF tests run again. The results for the first differences suggest that the null hypothesis of a unit root can be rejected at the 5 percent significance level for all series. Thus, the ADF tests indicate that wine prices and wine inventory series in their first differences are stationary at the 5 percent significance level for all countries.

Table 1: Augmented Dickey-Fuller Tests

No Intercept and No Trend				
Country	Wine Prices		Wine Inventory	
	Levels	1^{ST} Differences	Levels	1^{ST} Differences
Germany	0.41	-2.92*	0.45	-3.10*
Argentina	0.03	-2.57*	-1.50	-2.01*
Spain	-0.25	-2.81*	-0.09	-2.88*
Italy	0.18	-2.70*	0.28	-4.47*
Australia	0.68	-2.87*	1.93	-1.65*
France	0.54	-5.38*	-0.11	-4.81*
United States	0.78	-2.37*	-1.73	-7.47*

* Significant at 5% critical value

***Note*:** Critical values = -1.95 for 29 observations and -1.96 for 16 observations. With the exception of the USA and France, all countries have 16 observations.

When a constant and trend are included (Table 2), the ADF test results for the levels series suggest that in all series, except one (Australia), the null hypothesis of a unit root cannot be rejected at the 5 percent level, implying that with the exception of wine prices in Australia, all series are non-stationary in their levels. In their first differences, the null hypothesis of a unit root can be rejected for France, Italy and US wine inventories, and only France for wine prices. For all the other series, namely wine prices for US, Argentina, Australia, Germany, Italy and Spain; and wine inventories for Germany, Argentina, Spain and Australia, the null hypothesis of a unit root cannot be rejected at the 5 percent significance level. Overall, the unit root tests suggest that under the no-constant-no-trend specification, all series are I(1) but when a constant and trend are introduced, only wine price series for France and wine inventory series for Italy, France and the US are I(1).

Table 2: Augmented Dickey-Fuller Tests

With Intercept and Trend				
Country	Wine Prices		Wine Inventory	
	Levels	1^{ST} Differences	Levels	1^{ST} Differences
Germany	-3.38	-2.64	3.08	-3.25
Argentina	-2.09	-2.30	-1.80	-2.23
Spain	-1.82	-2.65	-2.03	-2.86
Italy	-2.19	-2.74	-1.82	-4.40*
Australia	-4.51*	-2.56	-1.99	-2.23
France	-1.56	-6.19*	-2.38	-4.78*
United States	-2.62	-3.54	-3.63	-8.04*

* Significant at 5% critical value
***Note*:** Critical values = -3.57 for 29 observations and -3.74 for 16 observations. With the exception of the USA and France, all countries have 16 observations.

Since the results of conventional tests can change with sample size, we conducted periodogram based unit root tests which are not influenced by sample size to check the stability of our results.

Periodogram based Unit Root Test

The Akdi and Dickey's (1998) periodogram based unit root test method has certain advantages over conventional tests. Firstly, conventional tests require the estimation of too many AR parameters to account for the dynamics of the series. Secondly, test results change with the sample size in conventional tests, while the periodiogram based method requires no parameter estimation except for variance. Thirdly, the critical values of the test statistics are free of sample size constraints. Thus, these might have considerable advantages, especially for the small sample wine data.

As described in Akdi, Berument and Cilasun (2007), one may use the trigonometric transformation of the series for this test. Given a time series $\{Y_1, Y_2, ..., Y_n\}$, the periodogram ordinate (without any model specification) is,

$$I_n(w_k) = \frac{n}{2}\left(a_k^2 + b_k^2\right) \tag{1}$$

where a_k, b_k are the Fourier coefficients and defined as

$$a_k = \frac{2}{n}\sum_{t=1}^{n}(Y_t - \overline{Y})\cos(w_k t) \text{ and } b_k = \frac{2}{n}\sum_{t=1}^{n}(Y_t - \overline{Y})\sin(w_k t)\,. \tag{2}$$

It should be noted that when $w_k = 2\pi k/n$, the following equality appears $\sum_{t=1}^{n}\cos(w_k t) = \sum_{t=1}^{n}\sin(w_k t) = 0$ and this causes the Fourier coefficients to be invariant to the mean, and, therefore, the periodogram ordinate is invariant to the mean (Akdi, Berument and Cilasun, 2007). Moreover, periodogram based unitroot/cointegration tests have the advantage of being seasonality robust, and model free from the selection of the lag lengths (see Akdi, 1995 and Akdi and Dickey, 1998).

In order to reject the null hypothesis of a unit root, one needs to observe small values of the periodiogram ordinates. Therefore, the values of the test statistics, $T(w_k)$ can be used to test for a unit root where

$$T(w_k) = \frac{2(1 - \cos(w_k))}{\hat{\sigma}^2} I_n(w_k) \tag{3}$$

The test statistics are distributed as a mixture of chi-squares exactly for AR(1) series under the assumption of stationarity (Akdi, Berument and Cilasun, 2007). In this case, the normalized periodogram will be distributed as chi-square with two degrees of freedom asymptotically. In conventional applications,, the power of the test is not exact. However, the power can be calculated analytically for the periodogram method to test for a unit root (see, Akdi, 1995). For higher order series, the same distribution is obtained asymptotically; that is,

$$T(w_k) = \frac{2(1 - \cos(w_k))}{\hat{\sigma}^2} I_n(w_k) \xrightarrow{D} Z_1^2 + 3Z_2^2 \tag{4}$$

where Z_1 and Z_2 are independent standard normal random variables and σ^2 is the variance of the error term. Here, the notation "$\xrightarrow{D}$" stands for convergence in distribution. The critical values of this distribution are provided by Akdi and Dickey (1998). Table 3 presents the results and the critical values. As shown in the table, the periodogram results support the conclusion derived from the conventional tests that all series are I(1) and hence, we can conduct cointegration tests under the no-constant-no-trend specification.

Table 3 Periodogram based unit root test

	$I_n(w_1)$	$\hat{\sigma}^2$	$T_n(w_1)$	$I_n(w_1)$	$\hat{\sigma}^2$	$T_n(w_1)$	Critical Values
	Wine Prices			Wine Inventory			
Levels							
Germany	554.95	0.0192	49605	16647.43	0.0038	13359228	0.178
Argentina	2393.95	0.0081	3063.6	746583.3	0.0044	6487305.3	0.178
Spain	21779.64	0.0508	105621	432053.8	0.0371	24125950	0.178
Italy	4722.25	0.0263	707118	150380.8	0.0168	1460875.5	0.178
Australia	1636.31	0.0127	433656	6250.414	0.0053	4508908.3	0.178
France	2989.43	0.0049	148776	243767.1	0.0126	62384683	0.178
US	11105.26	0.0013	571908	316242.3	0.0312	38028975	0.178
First Difference							
Germany	0.00434	0.02006	0.000062	0.0035	0.0043	0.000225	0.178
Argentina	0.01991	0.11683	0.000343	0.0304	0.0060	0.006389	0.178
Spain	0.01906	0.06249	0.000349	0.1045	0.0454	0.001044	0.178
Italy	0.03476	0.02903	0.002112	0.0251	0.0178	0.000422	0.178
Australia	0.01910	0.00639	0.005258	0.0278	0.0139	0.006379	0.178
France	0.01057	0.00526	0.000244	0.0096	0.0149	0.000046	0.178
US	0.00712	0.00161	0.000691	0.0378	0.0473	0.000012	0.178

5. COINTEGRATION BETWEEN STOCKS AND PRICES

Next we conduct cointegration analysis using three alternative techniques: the Engle-Granger (1987) two-step test, the maximum likelihood method developed by Johansen (1988) and Johansen and Juselius (1990) and the periodiogram method proposed by Akdi (1995). The Johansen method is preferred when there are more than two time series variables involved because it can determine the number of cointegrating vectors. Furthermore, less error is involved in the Johansen technique because only one step is involved rather than the two steps required in the Engle-Granger technique. As noted earlier, the periodiogram based method also has certain advantages over conventional tests, especially for small samples.

Beginning with the Engle-Granger cointegration test, if a series Y_t is non-stationary and there is a β vector (or matrix) such that $W_t = \beta' Y_t$ becomes stationary, then Y_t is considered cointegrated and the vector β is called the cointegrating vector. Previously in Table 1, it was shown that under the no-constant-no-trend specification, both wine price series (WP) and wine inventory series (WI) are I(1). Thus, these non-stationary series can be written as a linear combination of stationary and non-stationary series as

$$\begin{aligned} WP_t &= a_{11}\phi_t + a_{12}\varpi_t \\ WI_t &= a_{21}\phi_t + a_{22}\varpi_t \end{aligned} \tag{5}$$

where ϕ_t and ϖ_t represent the unit root and stationary component of these series, respectively. Since each component of the bivariate series includes the nonstationary component ϕ_t, both components of Y_t are non-stationary. However, if the coefficients $(a_{ij}, i, j = 1,2)$ are known, then

$$WP_t - \frac{a_{21}}{a_{11}} WI_t = \left(a_{22} - \frac{a_{21}a_{12}}{a_{11}} \right) \varpi_t = c\,\varpi_t \tag{6}$$

is stationary and the system is cointegrated with the cointegrating vector $\beta = \left(-\frac{a_{21}}{a_{11}}, 1 \right)'$. Since we do not know the coefficients, we normally need to estimate all the coefficients in equation (5). But now, it is sufficient only to estimate the ratio $\frac{a_{21}}{a_{11}}$ using OLS. The differenced series in (6) look like the residuals from the regression of WP on WI, and hence if the residual series is stationary, then the bivariate series is cointegrated. Moreover, the OLS estimator of the parameter WP obtained from that regression is a consistent estimator for the ratio $\frac{a_{21}}{a_{11}}$ (Engle and Granger, 1987). The results for the co-integration equations when wine prices (WP) are regressed on wine inventories (WI) and verse versa are reported in Tables 4a and 4b, respectively.

Table 4a. Co-Integration Regression: Prices regressed on Inventories.

Country Name	Coefficients	t-Ratio	R-Square	D-W Test
USA	-77.570*	-2.289	0.162	0.999
France	65.075	0.835	0.025	0.931
Germany	4.379	0.119	0.001	0.295
Italy	-103.65*	-3.893	0.487	1.403
Spain	22.677	0.684	0.028	0.352
Argentina	-35.355*	-4.147	0.518	0.768
Australia	59.901*	2.963	0.354	0.306

* Denotes significance at 5% level or higher

Table 4b. Co-Integration Regression: Inventories regressed on prices.

Country Name	Coefficients	t-Ratio	R-Square	D-W Test
USA	-0.002*	-2.289	0.163	0.199
France	0.000	0.835	0.025	0.477
Germany	0.000	0.119	0.001	1.075
Italy	-0.005*	-3.893	0.487	1.418
Spain	0.001	0.684	0.028	0.556
Argentina	-0.015*	-4.147	0.518	1.218
Australia	0.006*	2.963	0.354	0.672

* Denotes significance at 5% level or higher

To check for co-integration, the errors from the co-integration equations are recovered to perform non-stationarity tests using equation 7 since co-integration requires stationary residuals:

$$\Delta\varepsilon_t = \varpi\varepsilon_{t-1} + \sum_{i=1}^{p}\psi_i\Delta\varepsilon_{t-i} + \eta_t \tag{7}$$

where ε_t is the error from the co-integration equation, η_t is a stationary random error; here the null-hypothesis of non-stationarity is rejected when ϖ is significantly negative. The summation runs to ‘p’ where p is 2. Tables 5a and 5b report the ADF test statistics and the critical values. As shown in Table 5a, non-stationary of the residuals can be rejected at the 5-percent significance level for US, France, Italy and Argentina series, but the hypothesis cannot be rejected for Germany, Spain and Australia series. In Table 5b, non-stationary of the residuals can be rejected at the 5-percent level for all series except Spain and the US series.

Table 5a. ADF Tests for Errors when prices are regressed on Inventories

Country Name	ADF Test (Levels Series)	5 % Critical value
USA	-2.082*	-1.955
France	-2.117*	-1.953
Germany	-1.767	-1.961
Italy	-2.244*	-1.961
Spain	-1.793	-1.961
Argentina	-2.941*	-1.961
Australia	-1.436	-1.961

* Denotes significance at 5% level or higher

Table 5b. ADF Tests for Errors when Inventories are regressed on Prices

Country Name	ADF Test (Levels Series)	5 % Critical value
USA	-0.994	-1.955
France	-2.571*	-1.953
Germany	-2.909*	-1.964
Italy	-9.573*	-1.961
Spain	-1.764	-1.964
Argentina	-2.065*	-1.964
Australia	-6.003*	-1.961

* Denotes significance at 5% level or higher

Johansen's Cointegration Test

Next, we perform the Johansen cointegration test using the above procedure. The test results aappear in Table 6. Although the residual based test rejected the null hypothesis of no cointegration for US, France, Italy and Argentina series, the Johansen's cointegration test results suggest that the null hypothesis of no cointegration can be rejected only for the US and Argentina series. The observed inconsistencies in the cointegration results might be due to problems associated with sample size. Therefore, we examine cointegration using the periodogram test which is not influenced by sample size.

Table 6: Johansen's Cointegration Test Results

No intercept no trend Specifications				
Series: USTK vs UTWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.850186	59.10055	25.32	30.45	None **
0.312529	9.743133	12.25	16.26	At most 1
Series: FSTK vs FTWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.434781	22.13221	25.32	30.45	None
0.244742	7.298108	12.25	16.26	At most 1
Series: GSTK vs GWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.708330	25.22351	25.32	30.45	None
0.362012	6.741530	12.25	16.26	At most 1
Series: ISTK vs IWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.610421	21.29830	25.32	30.45	None
0.379480	7.157956	12.25	16.26	At most 1
Series: SSTK vs SWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.618633	20.14434	25.32	30.45	None
0.315428	5.684425	12.25	16.26	At most 1
Series: ARSTK vs ARWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.861254	34.15469	25.32	30.45	None **
0.260565	4.528031	12.25	16.26	At most 1
Series: AUSTK vs AUWP				
Lags interval: 1 to 2				
Eigenvalue	Likelihood Ratio	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.610966	19.03397	25.32	30.45	None
0.277359	4.872631	12.25	16.26	At most 1

* Denotes rejection of the hypothesis at the 5 percent significance level.

Periodiogram Cointegration Tests

Now we apply the periodiogram method proposed by Akdi (1995) to determine if there is a cointegrating relationship between the series. When the real part of the cross periodiogram ordinate of WI and WP series (say y_k) is regressed on the periodogram of WI (or WP) series (say x_k), the coefficient of x_k is also a consistent estimator for the ratio $\frac{a_{21}}{a_{11}}$ (Akdi, 1995). That is, when we consider the model,

$$y_k = \alpha + \beta \ x_k + \eta_k \ , \ k = 1,2,3,\ldots,[n/2], \tag{8}$$

the OLS estimator of β is a consistent estimator for the ratio $\frac{a_{21}}{a_{11}}$. The calculated values of $\hat{\beta}$ from equation 8 are reported in Table 7.

Table 7. OLS Regression Results for the Periodogram based Tests

Country	Constant	$\hat{\beta}$	P-value	R-Square
U.S.A	1.337	75.3	0.000	0.96
France	-2.386	159.2	0.004	0.52
Argentina	9.096	40.3	0.000	0.89
Australia	7.288	43.3	0.000	0.86
Germany	0.998	83.3	0.005	0.71
Italy	1.758	40.3	0.004	0.71
Spain	4.383	105.6	0.000	0.84

If the series $Z_t = Y_{2,t} - \hat{\beta} Y_{1,t}$ is stationary, then these two series are cointegrated. If Z_t is stationary, we will conclude that the WI and WP series are cointegrated. In order to perform this test, we regress ΔZ_t on Z_{t-1} and calculate the value of the $t-$statistics. The critical values are –3.43564 at the 5% level and –3.12867 at the 10 % level. The estimated results of this regression ($Z_t = Y_{2,t} - \hat{\beta} Y_{1,t}$) are reported in Table 8.

Table 8. Estimated Values of the Periodogram based Tests

Country	t-Statistics	R-Square	5% Critical Value
U.S.A	1.78541	0.11	-3.43564
France	2.19396	0.16	-3.43564
Argentina	-0.51608	0.02	-3.43564
Australia	-0.48016	0.02	-3.43564
Germany	-3.03394	0.38	-3.43564
Italy	-1.33291	0.11	-3.43564
Spain	-1.29607	0.10	-3.43564

Based on the periodogram results in Table 8, we fail to reject the null hypothesis of no cointegration for all series at the 5% significance level. Thus, while the evidence from the conventional tests was mixed, the periodiogram based analysis suggests that wine inventories and wine prices are indeed not cointegrated. There might be various reasons for the discrepancy of the test results. One possible reason for this is that the conventional tests require estimation of too many parameters to address the dynamics of the series with AR parameters. However, the periodogram based method is seasonally robust and requires no parameter estimation except for the variance (any consistent estimator of the variance can be used in the test statistics). These may account for the differences in the test results (see, Akdi, 1995, for details).

6. VECTOR AUTOREGRESSION AND IMPULSE RESULTS

The interrelationship between wine inventories and prices can be more directly examined using causality and vector autoregression analysis (VAR), e.g see Cromwell et al (1994). By incorporating time lags between these variables, these approaches are particularly relevant because changes in inventories typically my not cause changes in prices immediately but rather over several periods and vice-versa. Producers and consumers must first realize the stock changes in order to form price expectations. Wine inventories in bottles are not perishable in the short to medium run; some storage can even increase wine quality; and thus wine consumers may not be pressed to purchase quickly. In the case of high quality wines such as Bordeaux or Burgundy, the life expectancy may well last beyond 20 years or even 50 years.

The results of performing the Granger (1969) causality test between the inventory-price pairs revealed no strong forms of causality. We thus move directly to the VAR approach, which provides a useful means of analyzing the broad correlation in the variables of a system. This approach sidesteps the need for structural modeling by modeling every endogenous variable in the system as a function of the lagged values of all of the endogenous variables in the system. While the approach does not confirm causality, it at least evaluates the intertemporal influences between the variables. Estimated VAR's are used to calculate the percentages of each endogenous variable that are explained by innovations in each of the other endogenous as well as the explanatory variables and provides information about the relative importance of each random innovation to the variable in the VAR.

The mathematical form of a VAR is

$$Y_t = A_1 Y_{t-1} + \dots + A_p Y_{t-p} + \beta X_t + \varepsilon_t \qquad (9)$$

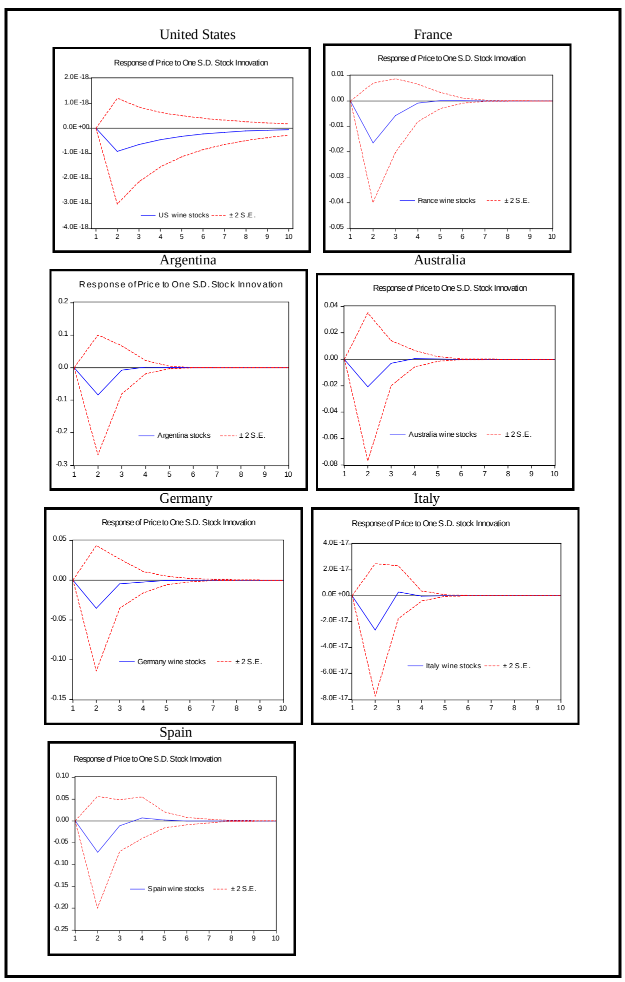
where Y_t is a k vector of endogenous variables, X_t is a d vector of exogenous variables, $A_1, \dots, A_p$ and β are matrices of coefficients to be estimated, and ε_t is a vector of innovations that may vary contemporaneously. The VAR model is used to highlight the impact of changes in wine stocks on wine prices in two ways: decomposition of the variance into forecast errors and secondly the analysis of impulse shocks

Our present interest is in discovering the lags and the signs of these lags, as they measure the impacts of wine stock changes on prices. Wohlgenant (1982) found that lags in inventory and shipment variables play an important role in explaining wine prices. This is best accomplished through impulse response functions that simulate the impacts of a shock of a given variable (leaving all variables endogenous) and then compute the predicted dynamic responses of each of the included variables. By treating the residuals of each variable/equation as unexplained innovations, the impacts of innovations are traced through the system by shocking the error terms. To employ the impulse functions, the VAR equations must first be estimated and the impulse response computed. The lack of strong cointegration between the endogenous wine variables permits us to procede in this direction. Because some nonstationarity was found in the time series of these variables, it is best to ensure stationaity by using some transform, in this case percentage changes. This transformation also conforms to the tenets of price theory. It is really changes in stocks that induce changes in prices in the short term (rather than the relationship in levels).

The method of estimating the VAR’s is normally of an unrestricted nature. All endogenous variables are thus of the same lag length in the estimation process. These forms of regressions were first performed for both variables. An attempt also was made to divide the later data span into 1986-1994 and 1995-2003 periods and to re-estmate the equations to investigae whether the inventory-price relation varied between the earlier and the later parts of the total time span. No clear evidence of a difference was found here. However, we discovered that the regression results could be improved by estimating restricted VAR’s in which only first order price lags and country specific inventory lags were used. The method employed was to estimate the cross-correlograms between the variable pairs and to discover which variable lags (i.e.1, 2, 3 or higher) had the most statistical significance. The results of estimating these restricted VAR’s are summarized in Appendix B. (the equation t-statistics and R-squared values are normally low because of the use of first-differences). Only first or second inventory lags are significant for all countries, except for Spain where a third-order lag proved meaninful.

In order to use the estimated VAR to analyze the interaction among wine prices and stocks in the structural models, impulse-response functions are computed by recovering structural innovations from the estimated residuals (linear combinations of uncorrelated structural shocks) coming from the VAR. The computed impulse functions (which show the difference between the expected value of the variable at time $t + i$ after a hypothetical shock at time t, and the expected value of the same variable at time $t + i$ given the observed history of the system) for each equation are given in Figure 4.

Figure 4. Impulse Response Functions



Summarizing these results, it is observed, for instance, that the equation results and impulses are explained best for France and Australia. By looking at the impulse response functions, there is some evidence in favor of French inventories affecting wine prices. Positive changes in French inventories are shown to decrease prices for up to 4 periods. Lag dependency is strongest for Australia at 2 periods but declines by the third period. The explanation for the United Sates is the next best. Here the lag dependency is the longest, extending to 7 periods. The results for Argentina, Italy and Spain show weaker dependence. In all three cases, the stongest inventory lag is at 2 periods and then finishes at 3 lags. For Germany the inventory influence is the weakest, most important by 2 lags but quitting at 3.

7. CONCLUSIONS

This study has employed carefully constructed national wine inventory and price series to determine what kind of relationships, if any, might exist between these variables. The underlying data, however, have been generated by national statistical agencies that have not compiled the kinds of data most useful for economic analysis. Microeconomic theory provides a clear background as to possible theories of inventory behavior and how inventory and prices variables might be interlinked. Only elementary aspects of this relationship have been examined in previous studies of wine inventory and price behavior.

The present empirical results employing univariate and multivariate time series analysis suggest that only a weak relationship can be confirmed. Trend stationary, cointegration and causality tests have provided some insights into the behavior of these variables. The results of the VAR analysis are correspondingly limited. However the shapes of the impulse functions do confirm the proper negative relationship between positive stock changes and falling prices, and negative stock changes and rising prices. These results are strongest for the wine markets of Argentina, Australia, France, Italy and Spain, and weakest for Germany.

It is hoped that this study has whetted the appetite for further research in this area. Most obviously more adequate inventory and price data are necessary. The best results would probably be obtained from data accumulated for specific wines and quality levels. This would permit not only improved price explanation but also the evaluation of some of the more complex inventory behavioral theories mentioned earlier.

Appendix A

WINE DATA SOURCES

Prices

Argentina: OIV Price Index, *Bulletin de L'OIV*, International Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield,, The Global Wine Statistical Compendium, AWBC and GWRDC, University of Monash.

Australia: OIV Price Index, *Bulletin de L'OIV*, Intenational Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield, *The Global Wine Statistical Compendium*, AWBC and GWRDC, University of Monash

France: Wines for current consumption, *vins de consommation courante*, Institut National de Statistiques et des Etudes Economiques, Paris.

Germany: OIV Price Index, *Bulletin de L'OIV*, Intenational Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield, The Global Wine Statistical Compendium, AWBC and GWRDC, University of Monash

Italy: Average price from the markets for table wine. ISMEA, Italy. OIV Price Index, *Bulletin de L'OIV*, International Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield,, The Global Wine Statistical Compendium, AWBC and GWRDC, University of Monash.

Spain: Average of daily quotations, *Semana Vitivinicola*, Spain. OIV Price Index, *Bulletin de L'OIV*, International Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield, The Global Wine Statistical Compendium, AWBC and GWRDC, University of Monash.

United States: Grape table wines, Producer price index WPU02610431, US Department of Labor, Washington DC.

Inventories

All Countries: Reported private and commercial stocks. *Bulletin de L'OIV*, International Office for Wines and Vines, Paris. G. Wittwer and J. Rothfield, The Global Wine Statistical Compendium, AWBC and GWRDC, University of Monash, 2006.

Appendix B

VECTOR AUTOREGRESSION RESULTS

Germany: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DGWP	DGSTK
DGWP(-1)	-0.061869	-0.088249
	(0.35410)	(0.16115)
	[-0.17472]	[-0.54761]
DGSTK(-1)	-0.579763	0.192011
	(0.63496)	(0.28897)
	[-0.91307]	[0.66446]
C	0.034944	0.013663
	(0.03930)	(0.01789)
	[0.88916]	[0.76392]
DGSTK(-2)	-0.082583	-0.093852
	(0.47880)	(0.21791)
	[-0.17248]	[-0.43070]
R-squared	0.070643	0.099725
Adj. R-squared	-0.161697	-0.125344
Sum sq. resids	0.249360	0.051647
S.E. equation	0.144153	0.065604
F-statistic	0.304049	0.443086
Log likelihood	10.58856	23.18423
Akaike AIC	-0.823570	-2.398029
Schwarz SC	-0.630423	-2.204881
Mean dependent	0.020993	0.016617
S.D. dependent	0.133745	0.061843
Determinant resid covariance (dof adj.)		7.79E-05
Determinant resid covariance		4.38E-05
Log likelihood		34.87741
Akaike information criterion		-3.359677
Schwarz criterion		-2.973382

Italy: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DIWP	DISTK
DIWP(-1)	-0.103283	-9.71E-17
	(0.33450)	(1.3E-16)
	[-0.30877]	[-0.74631]
DISTK(-1)	-0.409931	2.03E-16
	(0.38959)	(1.5E-16)
	[-1.05220]	[1.34171]
C	0.012991	3.48E-18
	(0.04201)	(1.6E-17)
	[0.30920]	[0.21314]
DISTK	-0.537577	1.000000
	(0.36356)	(1.4E-16)
	[-1.47864]	[7.1e+15]
DISTK(-2)	-0.115273	1.36E-16
	(0.36694)	(1.4E-16)
	[-0.31414]	[0.95195]
R-squared	0.210925	1.000000
Adj. R-squared	-0.076012	1.000000
Sum sq. resids	0.302567	4.58E-32
S.E. equation	0.165850	6.45E-17
F-statistic	0.735092	1.62E+31
Log likelihood	9.041316	
Akaike AIC	-0.505165	
Schwarz SC	-0.263731	
Mean dependent	0.020694	-0.008708
S.D. dependent	0.159884	0.134249
Determinant resid covariance (dof adj.)		1.14E-34
Determinant resid covariance		5.40E-35
Log likelihood		585.8201
Akaike information criterion		-71.97751
Schwarz criterion		-71.49464

Spain: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DSWP	DSSTK
DSWP(-1)	-0.153130	0.439610
	(0.30194)	(0.25293)
	[-0.50715]	[1.73804]
DSSTK(-1)	-0.383418	0.301977
	(0.33448)	(0.28019)
	[-1.14631]	[1.07775]
C	0.019372	0.012158
	(0.06174)	(0.05172)
	[0.31376]	[0.23507]
DSSTK(-3)	-0.477549	-0.062099
	(0.32098)	(0.26888)
	[-1.48778]	[-0.23095]
DSSTK(-2)	-0.002601	-0.117863
	(0.35667)	(0.29878)
	[-0.00729]	[-0.39448]
R-squared	0.278636	0.402614
Adj. R-squared	-0.009909	0.163659
Sum sq. resids	0.559834	0.392854
S.E. equation	0.236608	0.198205
F-statistic	0.965658	1.684896
Log likelihood	3.377164	6.033685
Akaike AIC	0.216378	-0.137825
Schwarz SC	0.452395	0.098192
Mean dependent	0.003160	0.013110
S.D. dependent	0.235444	0.216732
Determinant resid covariance (dof adj.)		0.001961
Determinant resid covariance		0.000872
Log likelihood		10.26973
Akaike information criterion		-0.035964
Schwarz criterion		0.436070

United States: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DUTWP	DUSTK
DUTWP(-1)	0.704785	-3.35E-16
	(0.15431)	(1.5E-16)
	[4.56728]	[-2.18636]
DUSTK(-1)	-0.027712	-1.97E-17
	(0.03174)	(3.2E-17)
	[-0.87301]	[-0.62408]
C	0.009630	2.07E-17
	(0.00873)	(8.7E-18)
	[1.10358]	[2.38768]
DUSTK	-0.016292	1.000000
	(0.04417)	(4.4E-17)
	[-0.36882]	[2.3e+16]
R-squared	0.500410	1.000000
Adj. R-squared	0.435246	1.000000
Sum sq. resids	0.025700	2.54E-32
S.E. equation	0.033427	3.32E-17
F-statistic	7.679247	1.86E+32
Log likelihood	55.60964	
Akaike AIC	-3.822937	
Schwarz SC	-3.630961	
Mean dependent	0.034878	-0.028445
S.D. dependent	0.044481	0.153987
Determinant resid covariance (dof adj.)		1.22E-36
Determinant resid covariance		8.88E-37
Log likelihood		1044.035
Akaike information criterion		-76.74332
Schwarz criterion		-76.35937

Australia: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DAUWP	DAUSTK
DAUWP(-1)	0.210177	0.176227
	(0.25283)	(0.16354)
	[0.83130]	[1.07756]
DAUSTK(-1)	-0.301299	-0.065912
	(0.40022)	(0.25888)
	[-0.75283]	[-0.25460]
C	0.078423	0.032394
	(0.04034)	(0.02609)
	[1.94427]	[1.24159]
DAUSTK(-2)	-0.694038	0.309379
	(0.39344)	(0.25449)
	[-1.76404]	[1.21567]
R-squared	0.288268	0.172361
Adj. R-squared	0.110335	-0.034549
Sum sq. resids	0.138980	0.058151
S.E. equation	0.107618	0.069612
F-statistic	1.620091	0.833025
Log likelihood	15.26511	22.23543
Akaike AIC	-1.408139	-2.279428
Schwarz SC	-1.214992	-2.086281
Mean dependent	0.039280	0.049392
S.D. dependent	0.114096	0.068440
Determinant resid covariance (dof adj.)		5.55E-05
Determinant resid covariance		3.12E-05
Log likelihood		37.59268
Akaike information criterion		-3.699085
Schwarz criterion		-3.312791

France: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DFCWP	DFSTK
DFCWP(-1)	0.099775	0.346445
	(0.25174)	(0.68953)
	[0.39633]	[0.50243]
DFSTK(-1)	-0.141769	0.251110
	(0.09795)	(0.26829)
	[-1.44737]	[0.93598]
C	0.026975	-0.008802
	(0.01305)	(0.03574)
	[2.06731]	[-0.24628]
DFSTK(-2)	-0.141692	-0.133356
	(0.10599)	(0.29032)
	[-1.33679]	[-0.45934]
R-squared	0.319468	0.114097
Adj. R-squared	0.173639	-0.075740
Sum sq. resids	0.026753	0.200707
S.E. equation	0.043714	0.119734
F-statistic	2.190711	0.601027
Log likelihood	33.06250	14.92562
Akaike AIC	-3.229167	-1.213958
Schwarz SC	-3.031307	-1.016098
Mean dependent	0.030638	0.002213
S.D. dependent	0.048088	0.115442
Determinant Residual Covariance		2.59E-05
Log Likelihood		43.97841
Akaike Information Criteria		-3.997601
Schwarz Criteria		-3.601880

Argentina: Vector Autoregression Estimates
Standard errors in ( ) & t-statistics in [ ]

	DARWP	DARSTK
DARWP(-1)	0.015152	0.022086
	(0.30269)	(0.07880)
	[0.05006]	[0.28030]
DARSTK(-1)	-1.073901	0.071431
	(1.16381)	(0.30297)
	[-0.92274]	[0.23577]
C	-0.060259	-0.045857
	(0.10456)	(0.02722)
	[-0.57634]	[-1.68479]
DARSTK(-2)	-1.228213	-0.023028
	(1.03307)	(0.26893)
	[-1.18890]	[-0.08563]
R-squared	0.198556	0.018653
Adj. R-squared	-0.001805	-0.226684
Sum sq. resids	1.110202	0.075235
S.E. equation	0.304166	0.079181
F-statistic	0.990993	0.076030
Log likelihood	-1.358643	20.17478
Akaike AIC	0.669830	-2.021848
Schwarz SC	0.862978	-1.828701
Mean dependent	0.034715	-0.045552
S.D. dependent	0.303892	0.071491
Determinant Residual Covariance		0.000565
Log Likelihood		14.42798
Akaike Information Criteria		-0.803498
Schwarz Criteria		-0.417203

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