The consumption of ordinary wines in France : the effect of administered prices

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Abstract

Since 1970 nearly all-Mediterranean countries of the EC had undertaken measures to regulate their domestic market for ordinary wines, in the context of constant fall in domestic demand for that product. This paper provides an empirical modelling framework for understanding the effect on the domestic demand for ordinary wines of the orientation price policy in the case of France. The result of a static demand co-integrating equation is compared with those obtained from alternative distributed lag and dynamic versions. It is shown that a non-static specification can result from a consideration of the effects in the short-run on demand responses of the price orientation mechanism. A restricted version of the static equation is tested for constancy using the Hansen's L_C test of parameter instability. We reject parameter stability, which confirms the trend followed by the series of ordinary wines over the sample period and supports the use of alternative specification for the demand equation. Across specifications, demand for ordinary wines is well explained by own-price changes, the price of juice and expenditures, especially in the short run. We also find evidence of habit persistence and a significant difference between unobserved and actual farm to retail spreads. A structural parameter associated to the degree of influence of the price support policy is estimated. Its statistical significance leads to the main conclusion that in the short run the presence of price regulation reduces demand responses after a change in the retail prices of ordinary wines.

JEL classification: C22, D12, Q11

Keywords : common agricultural policy, wine market, structural change, co-integration

1 Introduction

Domestic demand for ordinary wines in France and the major wine-producing countries of the European Community has been decreasing importantly since the mid-fifties. It was 160 litres per capita in 1965 and 80 litres in 1994. Poor quality production in this sector, changes in health habits and standard shifters such as low substitute prices and increased real income turned consumer purchases more to soft drinks and high-quality wines. In this context, price support policies were supposed to secure a sufficient level of income to producers. The welfare effects of the wine policy have been given a theoretical framework (see e.g., Jayet, 2001). However, as far as we know the influence of this policy on demand responses has not been investigated. The main purpose of this paper is to provide an empirical framework for understanding the impact on demand responses for ordinary wines of the orientation price policy. [1] In addition to this main objective, the present paper provides further estimates of demand responses for ordinary wines after a change in own-price, income, and other relevant variables in the demand function.

The issue addressed here is essentially concerned with the accuracy of known measures of own-price demand elasticities for regulated commodity. As far as we know, most studies consider rich consumer baskets that include wine with several beverages. Own-price and income elasticities for wine are generally found significant and close to unity. For instance Johnson et al. (1992), in a single equation co-integration framework for Canadian consumption of wine, beer and spirits, derive short run and long run own-price responses for wine and find respective values of –0.70 and –1.17. Other papers estimate systems of demand equations, employing AIDS models. Recently, Larivière et al. (2000) use a system wide approach to estimate demand equations for beer, wine and spirits and find total own-price elasticities essentially unitary across specifications. It is also confirmed in the pure dynamic multivariate time-series approach developed by McAvinchey (2000) for the U.K where the author finds significant own-price elasticity for wine very close to –1. It will be shown that these results are only roughly consistent with the ones found in the present analysis. Indeed, in the case of ordinary wines, whose price has been regulated for approximately 30 years, the values found for own-price demand responses are erroneous unless the constraints that the orientation price policy impose on retail price fluctuations are taken into account.

Regarding demand analysis on French data, we found an international comparison of price and income elasticities for wines across six nations for the period 1954-1971 in Labys (1976). For France the author estimated a static Marshallian demand equation and found a non-significant own-price elasticity of –0.06. An obvious reason that might explain this low value is that long run and short-run own-price responses were not independent of the model specification. It is now routine in applied researches on demand analysis with time series models to insert lagged consumption and lagged prices before estimating the demand equation. Unlike lagged consumption, a few economic justifications seem to support lagged prices in demand equations for alcoholic beverages whereas they generally improve statistical performances (see e.g., Johnson et al., 1992; McAvinchey, 2000). Lagged prices may serve to model an extrapolative process of expectation in demands for storage and speculation. However, we observe that consumers purchase ordinary wines for immediate consumption, frequently and futures markets do not exist for ordinary wines. Besides, we observe that the quality of ordinary wines does not generally increase with age or maturity so that ordinary wines are not held for investment purposes (for a discussion on these aspects, see Pompelli and Heien, 1991; Nerlove, 1995, p. 1704). Accordingly, in empirical studies on demand responses for wine, final consumers should not be supposed to extrapolate or forecast future prices thus taking only current price as exogenous. Consequently, consumer demand for ordinary wine at each period should be based only on the price of wine during that period. An immediate question arises out of these considerations: when lagged prices are significant in a demand equation, how may they be justified? A possible answer is that wine consumers do observe past prices for other reasons than that given above.

We propose an economic justification of the presence of significant lagged prices in a demand equation for ordinary wines. We relate those lags to the price intervention mechanism. As far as we know, no studies seem to have accounted for the CAP influence on ordinary wine purchases although wine is a major commodity of most Mediterranean countries of the EC (today in France, consumption for ordinary wines still represents about fifty percent of total wine consumption). We show that instead of imposing lagged prices *ad hoc*, a set of structural equations that account for the intervention price policy in this sector supports the presence of those lags. [2] We opt for a single equation model. This simple framework is sufficient to account for the constraints that the wine policy imposed on demand responses in the short-run.

In the next section we start estimating a static co-integrating equation of domestic demand for ordinary wines and discuss the results. As a complement to these results, and given the constant changes in demand during the last three decades we will also test for parameter instability employing the L_C test proposed by Hansen (1992). An alternative dynamic/distributed lag approach of the demand equation will be estimated in Section 3. Short-run responses will be derived from an associated error correction model (ECM) that is combined with the habit formation specification implied by lagged consumption. All regressions use an instrumental variable estimator to reduce a potential endogeneity biases. We present and discuss empirical results for the distributed lag specifications in Section 4 then conclude in Section 5.

2 Co-integration approach to domestic demand for ordinary wine

2.1 Description of the variables

The data span the period 1960-1995. [3] We retained this period since it turns out that the CAP ceased to apply its "Reference" price as of July 1995 that is the year of the Uruguay Round Accords. As from 1995 we anticipate that further structural changes have occurred. Including the most recent years may complicate the present analysis. The dependent variable selected for the study is annual per capita consumption of ordinary wines (q). Table 1 gives some basic statistics for this variable,

[Insert Table 1.]

Other variables are a retail price index for ordinary wines (p^r), a price index of soft drinks (p^j), an aggregate price index of "quality wines produced in specific regions" (p^h), and annual expenditures for all beverages, e. [4] Regarding demographic characteristics we create a variable that captures the essence of the shift in the composition of the population of wine consumers, without excluding a per-capita basis conversion for both consumption and expenditures. It is a ratio between two age groups, the population aged between 14 and 59 divided by that aged above 59. It is expected that this ratio accounts for a potential effect on wine consumption of the decreasing in mortality at older ages. It will be called d. We also considered originally the average annual temperature (w). See appendix 1 for a detailed exposition of the demographic and other variables. All models include a constant.

2.2 Static co-integrating equation

We make use of a static co-integrating regression only to capture the long-run relationships between our variables and compare it with the results of the distributed lags model set up in the next section. The first and necessary stage before estimating a co-integrating equation is to test the non-stationarity of the series. We used the Augmented

Dickey-Fuller tests combined with the Akaike and Schwartz information criterion to detect optimal size of lags. ADF test results are given in Table 2. We can not reject the null hypothesis of one unit root for all variables. Different specifications have been tested depending on the significance of the deterministic components. We regressed the first differences of the logarithm of the series against the lagged log-level and up to three lags for the first differences were used. However, two and three lags did not give consistent results for all variables. Lagged first differences are significant for all variables except for consumption for which a trend and a constant were significantly different from zero in the regression. In overall results the residuals present very little auto-correlation. Similar regressions were run in which the second difference in the series is regressed against lagged first differences. In this case, the null of a unit root was rejected for all variables although this result was less evident for the consumption variable when removing the deterministic components.

[Insert table 2.]

We tested for co-integration between the logarithms of the following I(1) variables: taxed consumption, retail price of ordinary wines, the price of soft drinks, that of high quality wines and total expenditures. The static demand equation tested for co-integration is

(1) $$q_t = c + \beta_r p_t^r + \xi^{c\prime} X_t^c + \xi^{e\prime} X_t^e + \varepsilon_t ,$$

where the ε_ts are assumed to be zero-mean identically distributed errors. The normalised vector of coefficients ξ^c is associated to X^c, the vector containing logarithms of the I(1) variables of the model other than p_t^r. The coefficients associated to the vector of exogenous regressors X^e are gathered together into ξ^e. In order to find a relevant estimator of β_r that measures the response of per capita consumption to retail price for ordinary wines, an important consideration must be given to shifts of supply and demand that have affected this coefficient. Supply during the first half of the sample remained at a high level and ordinary wine surpluses grew rapidly. At the converse, the second half reveals more important shifts in supply and recently a decrease at a lower rate of domestic purchases for ordinary wines. As a consequence of both factors, average prices of ordinary wines have maintained a slightly upward sloping trend. Shifts in demand were not independent of those in supply. These latter resulted from both the set-aside policy and especially vines replanting replacing grape varieties of poor quality with higher-quality grape varieties. Given these simultaneous shifts in both demand and suppply, a simple OLS regression could give an underestimated value of the own-price elasticity of demand. The large supply shocks (higher than that of demand) of the last fifteen years should help identify our equation. However, during the first part of the sample, demand shocks seem to have affected price more than supply did so. Accordingly an estimator employing instrumental variable is used.[5] Bearing acreages stand for the major instrumental variable that has shifted the supply curve and thus influenced retail prices. Instruments employed are thus the current and previous annual values for bearing acreage, the first lag of variables contained in X^c, the vector X^e plus the demographic variable lagged once. Lagged consumption was never inserted in the list of instruments. The test of the null of non co-integration remains based on the residual series $\hat{\varepsilon}_t$ derived from the estimated above equation. Given that constant and exogenous variables were included into the co-integrating equation, the procedure is a simple test of the existence of a unit root in the series $\hat{\varepsilon}_t$. Regression results for the coefficients of the demand equation are given in table 3a. The residual regression is performed on the equation $\Delta\hat{\varepsilon} = \rho^* \hat{\varepsilon}_{t-1} + a\Delta\hat{\varepsilon}_{t-1} + u_t$. The DF and ADF(1) statistics are –3.81 and –4.12, respectively. The co-integrating regression Durbin-

Watson (CRDW) test statistics equals 1.32. The 5% critical values for these three statistics are respectively -1.95, -4.51 and 1.36. Critical values for the CRDW is linearly extrapolated from the table reproduced in Banerjee et al. (1993, p. 209). Critical values for the ADF statistics are also contained in that reference.

[Insert table 3a]

Whereas the Dickey-Fuller statistics lead us to reject the null of no co-integration, the augmented Dickey-Fuller test (ADF) does not so. This latter with CRDW must be interpreted with caution for the following reasons. First, the number of observations is small. Second, the coefficient of the first lagged difference of $\hat{\varepsilon}_t$ in the residual regression is not always significant whereas higher order residual autocorrelation might be detected which would make the use of the CRDW statistics insufficient. Finally, the parameters of the demand equation might have shifted during the period we consider. The first issue has been justified earlier in this section. Regarding the second problem, a Breusch-Godfrey test for serial correlation was performed, regressing the estimated residual on all explanatory variables and residual lagged twice and thrice. The statistics are respectively 6.57 and 6.38 with 0.03 and 0.09 significance levels. Only the former exceeds the $\chi^2(2)$ critical value which is 5.99 at 5 % hence we reject autocorrelation of order more than two and accept autocorrelation at the second order. Consequently we can not sufficiently rely on the CRDW statistics. If we remove the constant and the weather variable, the ADF statistics are now in favour of co-integration ($t_{\rho^*=0;.05} = -5.53 < -4.39$) and the coefficient of first lagged difference in the residual regression becomes statistically significant ($t_{a=0} = 3.34 > 1.96$). Regarding the third possible source explaining the lack of significance of the residual based tests for co-integration in the present analysis a test of parameter instability has been conducted.

Before to run this test it seems important to summarize briefly the past trends in consumption over the sample period. Anti-alcoholism campaigns during the eighties influenced purchases by people over 50 who reduced or definitely stopped drinking wine (Aigrain et al. 1996). Whereas it seems obvious that strong shifts in tastes occurred for consumers who left the market, it is less evident for the others. Empirical suveys demonstrated that young people tend to drink less than their parents did at the same age and regular consumers aged 45 or over purchase less wine. Besides, the second half of the sample is also related to an increasing interest by consumers in wines identified by a geographical indicator (local wines or vin de pays from specific areas, and wines made from particular grapes). They have improved the general quality of the sector of ordinary wines.[6] Moreover, it is observed that drinking habits for local wines are slightly different. Household purchases of local wines are less occasional than those for wines labelled vin de table. Conversely to the first sub-period, during about the last 10 years covered by the sample, people are less sensitive to anti-alcoholism campaigns which are regressing (Aigrain et al., 1996, 1998).

Therefore, we suppose that the likelihood of parameter variation is relatively constant throughout the sample. In this case, Hansen (1992) suggests to use its L_C test statistics in the context of fully modified estimation. This statistics requires no decisions for trimming the sample. Critical values are computed for up to four I(1) regressors in addition to the deterministic terms. [7] The test requires a constant to be inserted (the underlying idea and demonstration are given in Hansen, 1992, pp. 327-328). Conversely to the residual-based tests above, the null hypothesis is parameter stability of the co-integrating equation. We test first for parameter stability in a co-integrating relation with the following regressors: retail prices of ordinary wines, soft drinks, quality wines and expenditures. The statistics was quite high, 19.70 and the constant not significant. The non significance of the parameter for the constant suggests that the relationship may be unstable with one coefficient, the intercept (in our case it is the constant) following a random walk. Another reason might be the presence of co-integrating relations between subsets of the regressors, which would decrease considerably the size of the test. If we remove the price of quality wines then we find L_C = 1.87 with a critical value of 1.31 at 1 per cent thus we still reject parameter stability but the rejection is less evident in this case. Finally, a further restricted specification has been tested where also the price of soft drinks is removed. It is then a test of parameter stability in a co-integrating relation relating consumption its own-price and expenditures. The statistics becomes .96 that is below the critical value, at 1 per cent. In this latter case, we hardly reject the null of a stable co-integrating equation relating taxed consumption to its own-price, income and a constant. Table 3b presents the last two results.

[Insert table 3b]

Though these results tend to reject the possibility of constant parameters in the standard co-integration model for the estimated restricted versions of the demand equation, they might also simply confirm the inability of a static model to capture the long-run relationships between the co-integrated variables. Indeed, the superiority of a specification involving lags of the explanatory variable to estimating a co-integrating equation when the sample size is small for example, is increasingly supported by numerous authors (Masih and Masih, 2000). The following section follows this branch of the literature. It turns out that a distributed lags demand equation is a relevant approach given the main purpose of this paper. Indeed, the structure of administered prices imply lagged prices in the demand equation.

3 Distributed-lags dynamic model: CAP influence on final consumption

3.1 Structure of retail prices

The declining level of domestic consumption for ordinary wines during the late sixties gave rise to unreasonable measures largely in favor of producers (Jayet, 2001). Market surpluses were distilled at a guaranteed price that was on average 30% of the orientation price whose structure is described in this section. Though price support together with the effect of the set aside policy had a considerable effect on the supply side of the market, we shall demonstrate in this section that the former measure might have affected demand responses as well. More precisely, both the magnitude and delay of the response of consumption to a change in its own-price must depend on the orientation price's structure. That should be the case for any regulated commodity. Assuming that the level of support prices influence market prices and given that the actual structure of the former is basically a function of past farm prices, then past retail prices also depend on support prices. This suggests that consumers of ordinary wines should observe past prices since current prices do not convey all the information about price.

The orientation of prices in the sector of ordinary wines was fully effective until 1995, the year of Uruguay Round reforms. A reference price is fixed at time *t* as an average of the farm prices observed at time *t*-1 and *t*-2 then this arithmetic mean is updated weekly with the prices from different market places until the end of the campaign (European commission, 1998, p. 72). We simply retain the levels of farm prices observed at time *t*-1 and *t*-2. We denote p_t^a the administered price and p_t^f the price cleared at farm. Computation of p_t^a is similar to an extrapolative process with respect to past values of p_t^f

$$p_t^a = \varphi p_{t-1}^f + \varphi' p_{t-2}^f, \quad 0 \le \varphi, \varphi' \le 1.$$

Over the sample, φ and $\varphi^{'}$ estimates are about 0.286 and 0.671 respectively with only the latter being significant at 5 per cent. A Fisher statistics support the linear restriction that the above coefficients sum to one.[8] This equation is re-parameterised as follows

(2) $$p_t^a = \phi p_{t-1}^f + (1-\phi)p_{t-2}^f, \quad 0 \le \phi \le 1$$

In the following equation we suppose that the unobserved farm price p_t^{*f} or the price that would be considered as normal by all market participants (essentially producers and wholesalers) depends upon a linear combination between observed farm price p_t^f and orientation price. Accordingly we write p_t^{*f} as follows

(3) $$p_t^{*f} = \lambda p_t^f + (1-\lambda)p_t^a, \quad 0 < \lambda \le 1.$$

The parameter in this equation reflects the trade off between selling more production at a low market price and receiving a guaranteed price. This parameter should enter wholesale prices hence retail prices. Equivalently, (3) can be rewritten as $p_t^f - p_t^a = (1/\lambda)\left(p_t^{*f} - p_t^a\right)$ where given $0 < \lambda \le 1$, the difference between actual farm price and orientation price is greater or equal to the difference between desired and orientation price. To put in other way, under regulation actual farm prices should exceed their unobservable counterpart or $p_t^f > p_t^{*f}$. From (3), two symmetrical situations might occur. In the former ($\lambda \downarrow 0$) the market is totally regulated. Conversely, if $\lambda = 1$ the market tends to be perfectly competitive and actual farm prices should reflect the price judged normal by market participants. We add the following mark-up formula that resume the transmission mechanism between farm and retail prices

(4) $$p_t^{*r} = p_t^{*f} + m^*, \quad m^* > 0,$$

(5) $$p_t^r = p_t^f + m, \quad m > 0.$$

These mark-up stand for constant equilibrium farm-to-retail price spreads on the raw materials contained in the retail product and marketing costs (see Chavas and Helmberger, 1996, p. 133). Assume that m^* equals the unobservable marginal cost of marketing in the long-run. Even though indirect taxes on alcohol are included in retail prices, we may assume that farm and retail prices should not drift too far apart in the long-run. We recognise that assuming m and m^* constant is a strong assumption which implies that the amount of production on the market has no effect on the marketing spread or that retail and farm demands should move closer at least in the long run (Ferris, 1997, p. 18). Under perfect competitive market conditions, in the long run, m should reflect marginal cost of marketing (Chavas and Helmberger, 1996, p. 140) and we expect the difference $m^* - m$ to tend to zero. Equations (2) and (3) together with (4) and (5) imply that under regulation ($\lambda < 1$), p_t^f and hence p_t^r will fluctuate in response to changes in p_t^a, i.e. in p_{t-1}^f, p_{t-2}^f. Given Eq. (4)-(5), consumers should consider the past values of retail prices when purchasing ordinary wines. Current prices alone provides only partial information suggesting that the amounts purchased would be different in the case of perfect information ($\lambda = 1$).

In order to obtain the reduced form equation for p_t^{*r}, insert (2) into (3) and use (4)-(5) to replace p_t^{*f} with $p_t^{*r} - m^*$ and p_t^f with $p_t^r - m$ which gives

(6) $$p_t^{*r} = \Delta m^* + \lambda p_t^r + (1-\lambda)\phi p_{t-1}^r + (1-\lambda)(1-\phi)p_{t-2}^r,$$

where $\Delta m^* \equiv m^* - m$. According to the value of λ, p_t^{*r} is readily interpretable. If $\lambda = 1$, then (3) becomes $p_t^{*f} = p_t^f$ and (4) is equivalent to $p_t^{*r} = p_t^f + m^*$. Given (5) then $p_t^{*r} = p_t^r + \Delta m^*$. In the long run Δm^* should tend to zero so that p_t^r becomes close to p_t^{*r}. Symmetrically, under regulation ($\lambda \downarrow 0$) then $p_t^{*f} \to p_t^a$ and $p_t^{*r} \to \phi p_{t-1}^r + (1-\phi) p_{t-2}^r$, suggesting that changes in the unobserved level of retail prices will highly depend upon the coefficient associated to the structure of intervention prices. Suppose now that ϕ equals zero then $p_t^{*r} = (1-\lambda) p_t^r + \lambda p_{t-2}^r$. Interestingly, in the long run, fluctuations in the unobserved level of retail prices only depend upon the coefficient measuring the degree to which support prices influence agents on market places. However, generally consumers are not aware of the level and structure of administered prices, which does not mean they ignore the information conveyed by current and past retail prices since all are observable.

3.2 Dynamics of quantities

Now we have defined a distributed lags specification of retail price for ordinary wines, a demand equation with lagged consumption also would be relevant to estimating long-run and short-run responses for three main reasons. Unexpected fluctuations in p_t^r and other variables might induce delays of adjustment of consumption to its desired value. Second, lagged consumption in the demand equation allows us to test for habit persistence (Baltagi and Griffin, 1995). The corresponding partial adjustment formulation (Nerlove, 1958) has the advantage that in some cases the underlying adjustment coefficient will be that of the error-correction form (see for instance Reziti and Ozanne, 1999). Unlike Johnson et al. (1992) who test and reject a partial adjustment mechanism for the demand equation for several beverages in favour of the two-step Engle and Granger approach, we combine both models.

Accordingly, q_t^* is defined as the desired (or long-run equilibrium) consumption and current consumption changes in proportion to the difference between its desired and current values or

(7) $$\Delta q_t = \gamma\left(q_t^* - q_{t-1}\right),$$

with $0 < \gamma \leq 1$. Combine (6) and (7) and assume the following model for the demand equation

(8) $$q_t^* = \beta_r p_t^{*r} + \xi^{c'} X_t^c + \xi^{e'} X_t^e + \varepsilon_t,$$

which is (1) with unobservable values for consumption and its own-priced. Using (6) and (8), the reduced dynamic co-integrating equation to be estimated is

(9) $$q_t = \gamma\beta_r \Delta m^* + \gamma\beta_r \lambda p_t^r + \gamma\beta_r (1-\lambda)\phi p_{t-1}^r + \gamma\beta_r (1-\lambda)(1-\phi) p_{t-2}^r + (1-\gamma) q_{t-1} + \gamma\xi^{c'} X_t^c + \gamma\xi^{e'} X_t^e + \gamma\varepsilon_t$$

We note that this equation does not depend on the administered price but only on the parameter relating this price to farm prices. Similarly, farm prices do not appear into the equation given the simple mechanism between farm and retail prices given in Eq. (4)-(5). An ECM is obtained after some linear transformations of (9)

(10) $$\Delta q_t = \gamma\hat{\beta}_r \Delta m^* + \gamma\hat{\beta}_r \lambda \Delta p_t^r - \gamma\left(q_{t-1} - \hat{\beta}_r p_{t-1}^r - \hat{\xi}^{c'} X_{t-1}^c\right) - \gamma\hat{\beta}_r (1-\lambda)(1-\phi)\Delta p_{t-1}^r + \gamma\xi^{c'} \Delta X_t^c + \gamma\tilde{\xi}^{e'} \tilde{X}_t^e$$

$\tilde{\xi}^e$ is associated with the transformed vector of exogenous variables $\tilde{X}^e$ which contains lagged and first differences of the exogenous variables *d* and *w*. This results from the linear transformation of (8). We restrict the elements in $\tilde{\xi}^e$ associated with the lagged exogenous variables to zero. Hence, only changes are considered. This equation has two main implications. First, the effect of a change in retail price on current consumption is inferior or equal to β_r not only because of γ but also λ which lies between zero and one. Thus λ provides also a measure of the degree to which the orientation price reduces observed consumer responses after a change in retail prices. Second, adjustment to disequilibrium error in the previous period is exactly γ, the adjustment coefficient of the PAM model. [9]

4 Empirical Results for the non static demand equation

Empirical results for the static regression have been discussed and were given Table 3a. Those related to the non static co-integrating regression are in Table 4. Table 5 shows ECM estimates in these cases.

[Insert Tables 4,5]

Residuals of the co-integrating regressions were tested for a unit root using the DF and ADF tests. First lagged changes never appeared significant in this extended framework, thus we focus on the former statistics. First order autocorrelation in the residuals is never rejected and higher serial correlation is found only in the regression cases 2 and 5. In the unconstrained regression (case 2) a Breusch-Godfrey test was performed on an auxiliary regression including all explanatory variables and residual lagged once then twice. In the latter case, the computed value is 8.03 and exceeds the $\chi^2(2)$ critical value that is 5.99 at 5 %. However, at the 1 % critical value, residual autocorrelation is rejected in all dynamic co-integrating regression. In case 2, the DF *t*-statistics (-3.79) leads us to accept the alternative of co-integration while the CRDW statistic (1.31) is very close to the critical value (1.36).

Parameter estimates produced with 2SLS for the static co-integrating equation stated in the previous section (case 1) are identically obtained in the dynamic framework assuming that $\gamma = \lambda = 1$ implying $q_t = \beta_r \Delta m^* + \beta_r p_t^r + \xi_t^{c\prime} X_t^c + \xi_t^{e\prime} X_t^e + \varepsilon_t$. Interestingly the constant can be interpreted since it contains the parameter Δm^* (*c* of Eq. (1) is replaced with $\beta_r \Delta m^*$ in Eq. (9) where $\gamma = \lambda = 1$). An estimate of Δm^* and its standard error can be given. They are –4.84 and 0.70 respectively thus this constant significantly different from zero whereas our model assumed the converse. It should be remembered that $\Delta m^* \equiv m^* - m$, hence the negative value of Δm^* would reflect the existence of large farm to retail spreads in the short-run are. Among the possible factors we might have high input prices, elevated mark-up and increasing government taxes. Large costs other than simple marks-up may exist at different levels of marketing. Between 1960-1961 and 1994-1995, the real price for wines of 9-12 degrees at farm decreased from FF2.03 per hectolitre to FF1.43, while at retail it decreased from FF5.29 to FF3.55. Interestingly the ratio of retail to farm prices remained about constant over the sample period, about 2.5.

Except for own-price responses and the demographic variable coefficient, estimates derived from the different dynamic co-integrating equations are close to those found in the static regression (case 1). Besides all coefficients have the anticipated sign. Long-run own-price responses derived from the dynamic co-integrating equations range from –1.13 to –1.41. These values are below that found in the static case that amounts to –1.10. These values in absolute terms are relatively high supporting that the 2SLS coefficient estimator approach reduced simultaneous bias. Such bias generally produces lower estimates for own-price changes in absolute value. Only in the cases 1 and 3 consumption responds significantly to the demographic variable. The significance of the coefficient reveals that the recent changes in the demographic structure (the decreasing in mortality at older ages) had an effect on consumption of ordinary wines. Larivière et al. (2000) finds the same result in their Almost Ideal Demand System applied to US data. Our result in the case of France supports their conclusion with taxed consumption as the dependent variable. Consumers respond more to a change in the price of juice than in that of high quality wines, reflecting that the former is a closer substitute in the daily consumption of beverages. This result holds across specifications. The coefficient on expenditure is the most significant one with a long-run positive value which is about 0.60. Interestingly, the weather variable has a significant positive effect at 5 per cent but only in the static regression case. We may expect this significance to increase in the short run, which will be discussed from the ECM estimates.

One parameter of interest is λ which is estimated to be 0.58 and ranges between 0.56 and 0.74. Presumably, wholesalers and producers have a different and imperfect information regarding the structure of p_t^a. Thus any value for λ (the degree to which actual farm prices and hence farm price fluctuations will depend on administered prices) must be interpreted with caution. The range of estimated values for this parameter suggests that less than half of the fluctuations in the price of ordinary wines are determined by the administered price policy. Estimated value of ϕ, the orientation price policy coefficient is about 0.90, which leads us to compare the model where $\phi = 0$ (case 3) with case 4 where $\phi = 1$. The most significant regression is that imposing $\phi = 1$. Comparing this regression to case 3, the F-statistic is largely increased and the hypothesis of no higher serial correlation than of order one is merely accepted. The PAM formulation (case 5) is not satisfactory in so far as it deletes lagged prices and grounds on the idea of habit persistence. Indeed, habit persistence does not seem to be a good assumption given the past structural changes in demand.

Table 5 reports short-run responses obtained with the ECM approach for cases 2-5. The demographic variable was deleted because it was not significant and affected the results. Therefore only the coefficient associated with a change in the weather variable was retained in the set of exogenous factors. As expected, weather changes appear more significant for both cases 2 and 3 with or without the constant term. The estimated short run own-price response has the anticipated negative sign (-0.68) which is about half the long-run values that ranged from –1.41 to –1.13. Even though some adjustment seem to have occurred ($\gamma > 0$), the error correction is not really significant with *t*- statistics ranging from -1.40 to -2.08 suggesting that consumption may be weakly exogenous (Engle and Granger, 1987). However, deleting the error-correction term gives worse results. The low significance of this coefficient might reflect changes in tastes, a result that was supported by the Hansen (1992)'s L_C test of parameter instability in section 2. In such context consumers are less likely to correct any disequilibrium from the previous period and it may be difficult for them to do so since prices do not convey the right information about market structural changes. For example, prices tended to remain high whereas consumption was decreasing.

5 Conclusion

The declining level of domestic consumption for ordinary wines during the late sixties gave rise to important measures applied to production. Whereas administered prices are more often considered in a supply-side analyses, this paper proposes a framework to estimate domestic demand responses for ordinary wines accounting for the constraints that the orientation price policy imposed on retail price fluctuations. It is motivated by the following idea that retail prices for a regulated commodity can not convey to the final consumer information about market conditions unless she is perfectly informed of the complex mechanism that relates retail prices to administered prices.

Long-run structural parameters derived from a static co-integrating demand equation were compared with those obtained from alternative non-static specifications. We followed a distributed lags and dynamic co-integration/error-correction approach in order to estimate separately long run and short run responses. Demand is found to be own-price elastic with a value generally below minus one whereas consumption responses after a change in income are close to .5. An ECM specification reveals that in the short run, only current and lagged own-price, expenditures and to a certain extent the price of juice explain ordinary wine consumption satisfactorily.

It is shown that only the distributed lags demand equations allow us to account for the influence from the orientation price policy. Parameters related to the orientation price policy are significant and the following conclusion is borne out that the structure of administered prices play a significant role in the study of the demand responses for the market of ordinary wines. Though less than half of retail price fluctuations seem to be determined by the price policy, nevertheless this latter would have the effect to reduce consumer responses after a change in the retail prices of ordinary wines. Consequently, consumers would need more than one period to fully respond to this change. We found evidence of habit formation, reflecting that responses of consumption are further delayed after a shock on price but also on the other variables of the model.

We plan to apply the present modelling framework to panel data in order to account for consumers' perception of the wine policy. Our results might be useful for the future Common Market Organisation for ordinary wines and for cross-countries comparisons. We suggest that more attention should be focused on the influence of farm policies on demand responses for a regulated commodity.

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NOTES

[1]. A recent survey of the literature on the demand for alcoholic beverages is Larivière et al. (2000). It reflects significantly the small number of papers in international publications dedicated to the study of demand responses for oridinary wines in the case of France. One exception is Labys (1976) who provided a pioneer study, estimating a single demand equation for different countries. Jayet (2001) addresses market welfare issues.

[2]. Incentive aspects of the Common Market Organisation for ordinary wines are analysed in Jayet (2001).

[3]. Data were supplied by the National Institute of Agronomic Research (INRA), the General Customs Service (DGI) and may be found in part in the report of the Inter-professional National Organism for Wines (ONIVIN).

[4]. The French designation for non ordinary wines is vins de qualité produits dans une région déterminée (VQPRD), which includes vins d'appélation d'origine contrôlée (AOC) and vins de qualité supérieure (VDQS).

[5]. Until the mid-eighties demand shifted strongly and supply remained high because of the existence of attractive floor prices. For this period, OLS would give a consistent estimate of the own-price elasticity of supply. After this period, vine-removing policy reduced supply through destruction of existing plants while demand was decreasing more slowly thus OLS should give a consistent estimate of the own-price demand elasticity. Consequently, an OLS regression applied to the whole sample might produce an average of both the own-price demand and supply elasticities. A 2SLS should reduce the endogeneity bias associated to the former period.

[6]. Local wines were introduced in 1974. Though local wines are ordinary wines, they are not perfectly homogenous with wines corresponding to the appellation vin de table, nevertheless the Commission of the European Community placed both products together in a single category.

[7]. We have developed a program to run the test under RATS econometric software. The program and the data on which the test is applied are available on request from the author.

[8]. A regression of orientation price at time *t* on farm prices at times *t*-1 and *t*-2 plus a constant is done on a sample covering the period 1970-1971 to 1993/1994. We test the null hypothesis that the lagged market price coefficients sum to one. A Fisher test is conducted. The calculated *f*-statistic equals 0.098 which is less than the $F(1, 21)=4.32$ statistics at 95%. Hence, we do not reject the null of a sum of coefficient of unity. Notice that this result holds only if a constant is added. Its estimated value is about –4.07. This estimate of the constant certainly stands for the part of the structure of the administered price we have arbitrarily omitted.

[9]. See Reziti and Ozanne (1999) for a discussion of different forms such as PAM and ECM that can be derived from a general first-order dynamic model.

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APPENDIX

q : taxed consumption of red, rosé and white non-sparkling wines produced in France and other countries (except for sweet and fortified wines such as *Porto*) sold in bottles (and small bulk containers). In Europe ordinary wine_ always denotes a still wine containing at least 8.5 and below 14-15 percent alcohol, with no flavour additives, and meeting minimum standards regarding grape varieties and acidity (European Commission, 1998: 69).

p^r : the real retail price for red wines of 11 degrees in Paris and its suburbs. Note that red and rosé ordinary wines represent the majority of the market for bottled wine. Deflator is the CPI.

p^j : real price index of soft drinks. Deflator is the CPI.

p^q : real price index of VQPRD wines. Deflator is the CPI.

e : expenditures for overall beverages. Have been deflated by the general CPI and divided by population aged over 14.

d : a ratio between the population aged between 14 and 59 and that over 59. During the 20th century, death rates at older ages have fallen in France. The pace of decline has been more rapid in recent decades than it was in the earlier decades. Two age groups (see appendix) characterized with different positive average rate of growth over the last fourty years are considered in this study. The young-aged group has an inferior rate than the old group. Given the reduction in the consumption of ordinary wines and the existing interactions between drinking habits, health and longivity, using a ratio between the two groups may be a way to account for a potential effect of the decreasing in mortality at older ages on wine consumption. A positive elasticity value would demonstrate that sold quantities of ordinary wines may not benefit from that structural change

w : a ratio between the maximal temperature and the average temperature of the corresponding year. Variables are measured in the Montsouris Centre of Meteo France in Paris.

p^a : orientation price for the RI and RII wine categories. Since 1975-1976, the categories RI and RII have had equal orientation prices. This price was deflated by the CPI. Ordinary wines of RI type are red wines from grapes other than *Portugieser* type. They contain between 10 and 15 per cent alcohol.

p^f : farm prices given in the different quotation places for ordinary wines situated in Bastia and in the South of France: Beziers, Montpellier, Narbonne, Nîmes, Perpignan and Brignoles. Local wines are not quoted in these places. Red and Rosé wines are quoted together while white wines are quoted apart. The farm price is simply an average of both quotations. This price is deflated by the CPI.

Table 1. Descriptive Statistics of the demand variables in level, 1960-1992.

Mean	93.31[a]
Standard Deviation	26.28[a]
Min/Max values	42.00/128.61[a]
Corresponding harvesting year	1994-1995/1959-1960

[a] In litre per capita a year.

Table 2. Stationarity tests

	p^r	q	p^j	p^h	E
ADF Statistics	-0.87	-1.20	-1.13 [a]	0.72	-3.54
5% critical values	-2.63	-4.27	-2.63	-2.63	-4.27
Number of lags	1	0	1	2	2
t-of trend coefficient	None	-2.42	None	None	-3.11
t- of constant	None	1.19	None	None	-3.31
D&W	2.06	2.16	1.89	2.04	2.08

[a]. The values reported in the first row are the pseudo *t*-statistics from the Augmented Dickey-Fuller tests.

Table 3a. Static co-integrating demand equation, case 1.

p^r	p^j	p^q	e	d	w	c
-1.10	0.48	0.24	0.60	0.73	0.17	5.34
(-11.24)	(3.74)	(2.26)	(53.60)	(2.98)	(2.12)	(7.07)

t-statistics in brackets. Coefficient of determination: R_a^2 = 0.995

Table 3b. Fully Modified Estimation of a
restricted version of the static co-integrating demand equation.

p^r	p^j	e	c	L_c
-1.03	0.91	0.55	4.90	1.87
(-3.30)	(3.79)	(12.48)	(3.92)	
-6.65		1.07	36.01	.95
(-11.06)		(14.19)	(12.62)	

t-statistics in brackets. Fully modified estimation of the co-integrating regression between consumption and four I(1) regressors. The corrected standard errors are computed following Hansen (1992).

Table 4. Dynamic co-integrating demand equations. Estimated long-run coefficients [a]

p^r	p^j	P^q	e	D	w	γ	ϕ	λ	Δm^*
Unrestricted dynamic equation (regulation and adjustment costs), case 2									
-1.41	0.59	0.29	0.64	0.37	0.19	0.71	0.89	0.56	-4.75
(-3.26)	(3.17)	(2.12)	(14.14)	(1.05)	(1.64)	(1.88)	(3.11)	(3.01)	(-3.95)
Restricted dynamic equation (regulation with $\phi=0$ and adjustment costs), case 3.									
-1.30	0.45	0.18	0.63	0.68	0.15	0.70		0.74	-5.17
(-5.12)	(1.88)	(1.05)	(13.64)	(2.01)	(1.22)	(1.68)		(4.92)	(-4.60)
Restricted dynamic equation (regulation with $\phi=1$ and adjustment costs), case 4.									
-1.37	0.61	0.31	0.63	0.37	0.19	0.73		0.58	-4.43
(-7.73)	(3.74)	(2.51)	(21.64)	(1.20)	(1.92)	(2.66)		(4.11)	(-6.50)
Partial adjustment model ($\lambda=1$), case 5.									
-1.13	0.36	0.17	0.62	0.80	0.20	0.56			-5.45
(-6.44)	(1.30)	(0.89)	(16.96)	(1.81)	(1.80)	(1.77)			(-3.83)

[a] t-statistics in brackets.

Table 5. Error-correction models. Estimated short-run coefficients

Model	Δp^r	Δp^j	Δp^q	Δe	z_{t-1}	Δp^r_{t-1}	Δw	c^a
Constant								
Cases 2, 3	-0.68	0.42	0.36	0.39	-0.21	-0.36	0.15	-0.01
	(-4.12)	(2.09)	(2.56)	(2.69)	(-1.40)	(-2.96)	(2.36)	(-1.36)
Cases 4, 5	-0.45	0.19	0.18	0.23	-0.21		0.09	-0.02
	(-2.74)	(0.94)	(1.30)	(1.57)	(-1.42)		(1.46)	(-2.07)
No Constant								
Cases 2, 3	-0.78	0.64	0.41	0.57	-0.30	-0.39	0.17	
	(-5.27)	(4.89)	(3.08)	(8.95)	(-2.08)	(-3.20)	(2.78)	
Cases 4, 5	-0.60	0.52	0.26	0.51	-0.30		0.10	
	(-3.78)	(3.56)	(1.75)	(7.18)	(-1.96)		(1.55)	

[a] t-statistics in brackets. $c \equiv \gamma\beta_r \Delta m^*$ of the ECM.