On the Use of Panel Data Methods to Estimate Rational Addiction Models*

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Abstract

This paper reviews some of the econometrics problems faced when estimating a rational addiction model of Becker and Murphy (1988) using panel data. This is illustrated with three empirical applications. The first application looks at cigarette consumption using macro panel data on American states over time (see Baltagi and Griffin (2001)). The second application looks at liquor consumption using macro panel data on American states over time (see Baltagi and Griffin (2002)). The third application uses micro panel data on Russian alcohol consumption (see Baltagi and Geishecker (2006)).

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There is no doubt that econometrics is subject to important limitations, which stem largely from the incompleteness of the economic theory and the non-experimental nature of economic data. But these limitations should not distract us from recognizing the fundamental role that econometrics has come to play in the development of economics as a scientific discipline. It may not be possible conclusively to reject economic theories by means of econometric methods, but it does not mean that nothing useful can be learned from attempts at testing particular formulations of a given theory against (possible) rival alternatives.

Pesaran (1990, pp. 25-26)

1 INTRODUCTION

Stigler and Becker (1977) and Becker and Murphy (1988) proposed an elegant theory of rational addiction which has been used to explain addiction to cigarettes, alcohol, caffeine, illicit drugs, opium, cocaine, etc. The theoretical model assumes forward-looking, utility maximizing consumers who may become addicted to the consumption of a good. Consumers are rational in that they anticipate the expected future consequences of their current actions. They may recognize the addictive nature of their choices but decide to make them because the gains from the activity exceed the costs of possible future addiction. In particular, the more a person drinks alcohol or smokes cigarettes, the higher the current utility derived. However, the individual recognizes that he or she is building up a stock of this addictive good that is harmful. The individual rationally trades off these factors to determine the appropriate level of drinking or smoking. This is in contrast to prior ‘myopic’ theories of habit formation where current utility depended only on past and current consumption of the addictive good, see Pollak (1970, 1976). The myopic consumer ignores the possibility that his or her current choices would affect future tastes or health.

This theory is not without its critics; for example, Winston (1980) argues that addicts in this model are happy, which is inconsistent with observed regret among addicts. Akerlof (1991) argues that addicts in this model choose to become addicts and there is no scope for curbing their addictions with education programs, which is incompatible with any role for information and public policy. However, Orphanides and Zervos (1995) provide a rational theory of addiction with learning and regret that resolves some of these criticisms. The basic idea is to allow for uncertainty rather than perfect foresight and a process of learning through experimentation. Their theory explains how individuals can be voluntarily drawn

into a harmful addiction and later regret it. Gruber and Köszegi (2001) question the ‘time consistent preferences’ assumption required by the Becker and Murphy (1988) theory. Dropping this time consistent preferences assumption still yields forward-looking behavior but strikingly different normative policy implications.

The Becker and Murphy (1988) theory has been applied to the consumption of cigarettes, (see Chaloupka (1991), Becker, Grossman and Murphy (1994), Labeaga (1993, 1999), Baltagi and Griffin (2001), Gruber and Köszegi (2001) and Jones and Labeaga (2003) to mention a few); to the consumption of alcohol, (see Grossman, Chaloupka and Sirtalan (1998) and Baltagi and Griffin (2002) and Baltagi and Geishecker (2006)); to the consumption of caffeine, (see Olekalns and Bardsley (1996)); cocaine, (see Grossman and Chaloupka (1998)); illicit drugs, (see Saffer and Chaloupka (1999)) and opium, (see Liu, et al. (1999)). A key feature of this theory is that consumption of an addictive good will depend on future, as well as, past consumption. Finding future consumption statistically significant is a rejection of the myopic model of consumption behavior.

2 THE RATIONAL ADDICTION MODEL

Following Becker, Grossman and Murphy (1994), denoted by BGM, the consumer’s problem is to maximize the sum of lifetime utility discounted at rate r :

$$\sum_{t=1}^{\infty} \beta^{t-1} U(C_t, C_{t-1}, Y_t, e_t), \quad (1)$$

where $\beta = 1/(1 + r)$, C_t is the quantity of cigarettes consumed in period t , Y_t is the consumption of a composite commodity in period t , and e_t reflects the impact of unmeasured life-cycle variables on utility. BGM take the composite commodity Y as the numeraire and the rate of interest is assumed to be equal to the rate of time preference. This maximization is subject to the following constraints:

$$C_o = C^o \quad \text{and} \quad \sum_{t=1}^{\infty} \beta^{t-1} (Y_t + P_t C_t) = A^o, \quad (2)$$

where P_t is the price of cigarettes at period t , C^o is the initial condition indicating the level of cigarette consumption at period zero, and A^o is the present value of wealth. The associated first-order conditions are the usual marginal utility of wealth w is equal to marginal utility of other consumption in each period U_y :

$$U_y(C_t, C_{t-1}, Y_t, e_t) = w.$$

Also, marginal utility of current cigarette consumption U_1 plus the discounted marginal effect on next period's utility of today's consumption, U_2 , equals the current price multiplied by the marginal utility of wealth,

$$U_1(C_t, C_{t-1}, Y_t, e_t) + \beta U_2(C_{t+1}, C_t, Y_{t+1}, e_{t+1}) = wP_t.$$

Assuming the utility function is quadratic and solving the first-order conditions for C_t , BGM obtain the following first-difference equation:

$$C_t = \theta C_{t-1} + \beta \theta C_{t+1} + \theta_1 P_t + \theta_2 e_t + \theta_3 e_{t+1}, \quad (3)$$

where

$$\begin{aligned} \theta_1 &= (u_{yy}\lambda / [(u_{11}u_{yy} - u_{1y}^2) + \beta(u_{22}u_{yy} - u_{2y}^2)]) < 0, \\ \theta_2 &= -(u_{yy}u_{1e} - u_{1y}u_{ey}) / [(u_{11}u_{yy} - u_{1y}^2) + \beta(u_{22}u_{yy} - u_{2y}^2)], \\ \theta_3 &= -\beta(u_{yy}u_{2e} - u_{2y}u_{2e}) / [(u_{11}u_{yy} - u_{1y}^2) + \beta(u_{22}u_{yy} - u_{2y}^2)]. \end{aligned}$$

The lower case letters denote the coefficients of the quadratic utility function. Equation (3) gives current cigarette consumption as a function of past and future cigarette consumption, P_t and the unobservable shift variables e_t and e_{t+1} . The latter variables reflect the impact of unmeasured life cycle variables. BGM recognize that the unobservable errors e_t that affect the utility in each period are likely to be serially correlated. Even if these variables are uncorrelated, e_t directly affects utility in each period and affects consumption at all dates through the optimizing equation (3). Therefore, BGM treat C_{t-1} and C_{t+1} as endogenous and use lagged and future prices as instruments.

3 EMPIRICAL RESULTS

3.1 The U.S. Cigarette Data

BGM used cigarette data from 1955-1985 for the 50 American states, including the District of Columbia. Their empirical equation for cigarettes includes other exogenous variables, such

as income, short and long distance smuggling indexes and taxes. BGM estimate their model using fixed effects 2SLS, which is essentially 2SLS with state dummies included in the model. Their preferred instrument set, includes future prices of cigarettes, and yields a positive and statistically significant effect of lagged consumption, i.e., θ . This confirms that cigarette smoking is addictive. The estimated coefficient of C_{t+1} , i.e., $\beta\theta$, is positive and significant indicating that future consumption of cigarettes affects current levels of consumption. This finding clearly rejects the hypothesis of myopic addiction behavior. The implied short-run and long-run price elasticities computed at the sample mean of BGM's data are -0.41 and -0.73, respectively. Using the notation in equation (3), the second difference equation has real roots provided $4\beta\theta^2 < 1$. Appendix A of BGM characterizes this condition as a 'stability condition'. Given that the solution to a rational addiction model is generally assumed to be a saddle point, its roots could not pass a stability test, see Ferguson (2000). A saddle point solution, however, does require that the roots be real, and in that context, this condition is useful and has to be checked. For the U.S. state cigarette data considered by BGM, the roots are real and the long-run price elasticity is larger than the short-run price elasticity in absolute value as predicted by the theory. Some of the troubling results obtained by BGM that are in conflict with the rational addiction theory were the negative interest rates ranging from -81% to -88% (see their Table 5). This occurred when BGM excluded future price, P_{t+1} , from the set of instruments. BGM find that a Hausman (1978) test rejects the hypothesis that future price values like P_{t+1} , are legitimate instruments, but nevertheless argue for its use. BGM justify their choice of P_{t+1} as an instrument even though it fails the Hausman test, on several grounds. First, BGM reason that using only past prices yield poor instruments, which in turn can produce a large bias in the estimated coefficients of the endogenous variables. This is the weak IV story (see Nelson and Startz (1990)). Second, BGM argue that using P_{t+1} as an instrument yields more robust estimates against slight misspecification of the demand equation than instrumental variables estimators that omit P_{t+1} . Finally, they argue that consumers may have relevant information to forecast future cigarette prices. For example, tax hikes are announced in advance and it is plausible that consumers anticipate the price increase well in advance. Even with using P_{t+1} as an instrument, the positive implied interest rate ranges between 56% and 223%.

Baltagi and Griffin (2001) consider a variant of (3) as follows:

$$C_{it} = \delta_0 + \delta_1 C_{i,t-1} + \delta_2 C_{i,t+1} + \delta_3 P_{it} + \delta_4 Y_{it} + \delta_5 Pn_{it} + u_{it}, \quad (4)$$

where the subscript i denotes the i -th state ($i = 1, \dots, 46$) and the subscript t denotes the t -th

year ($t = 1, \dots, 30$), namely, 1963-1992. C_{it} is real per capita sales of cigarettes (measured in packs of cigarettes per head) by persons of smoking age (16 years and older). P_{it} is the average retail price of a pack of cigarettes measured in real terms. Y_{it} is real per capita disposable income. Pn_{it} denotes the minimum real price of cigarettes in any neighboring state.¹ The last variable allows for possible “bootlegging effects,” which Baltagi and Levin (1986) found important in explaining why some very low tax states enjoy much higher cigarette sales than neighboring states with higher taxes. For other ways of dealing with short and long distance smuggling of cigarettes, see BGM.

We follow the usual convention of assuming that the disturbance term in equation (4) is specified as a two-way error component model:

$$u_{it} = \mu_i + \lambda_t + \nu_{it} \quad i = 1, 2, \dots, 46; \quad t = 1, 2, \dots, 30, \quad (5)$$

where μ_i denotes a state-specific effect, λ_t denotes a year-specific effect, and ν_{it} is a remainder disturbance. One of the advantages of a panel is its ability to control for all time-invariant variables or state-invariant variables, whose omission could bias the estimates in a typical cross-section or time series study. Both effects can be assumed either fixed or random. We assume that the time-period effects (the λ_t 's) are fixed parameters to be estimated as coefficients of time dummies (D_t) for each year in the sample. This can be justified given the numerous policy interventions, as well as, health warnings and Surgeon General's reports which previous studies accounted for using time dummy variables, (see Baltagi and Levin (1986)). Similarly, the μ_i 's are state-specific effects which can represent any state-specific characteristic, including, but not limited to the following: (i) Indian reservations selling tax-exempt cigarettes, (ii) tax-exempt military bases in certain states, (iii) a high percentage of Mormons (a religion which forbids smoking), (iv) a highly touristic state, such as Nevada, with a per capita consumption of twice the national average. BGM included both a set of dummies for states and time in all specifications considered. This model is known as the fixed effects model.

Baltagi and Griffin (2001) suggested alternative dynamic panel data estimators suggested by Anderson and Hsiao (1982), Arellano and Bond (1991) and Keane and Runkle (1992). BGM relied solely upon the fixed effects 2SLS estimator using two sets of instrumental variables (IV) estimators. The first IV estimator uses lagged prices $[P_{t-1}, P_{t-2}]$ among the set

¹For data sources, as well as details on the construction of the neighboring price, see Baltagi and Levin (1986). Per capita sales of cigarettes and the average retail price per pack were obtained from The Tobacco Institute (1993).

of instruments. The second IV estimator adds future prices $[P_{t+1}]$ to the set of instruments. However, it is important to emphasize that the fixed effects 2SLS estimator is inconsistent for fixed T , since the regressors include lagged and future values of the dependent variable. Note that the bias in the fixed effects estimator can be small when T is large and the processes are weakly dependent. In response to the problems with FE2SLS, GMM estimators were developed that exploit the numerous orthogonality conditions between P_t and u_t . In fact, if P_t is uncorrelated with u_t for all $t = 1, \dots, T$, then in principle, all prices $[P_1, P_2, \dots, P_T]$ should be used as instruments for each time period. In this case, one first differences (4) to get rid of state specific unobservable effects. The resulting equation is:

$$\Delta C_{it} = \delta_1 \Delta C_{i,t-1} + \delta_2 \Delta C_{i,t+1} + \delta_3 \Delta P_{it} + \delta_4 \Delta Y_{it} + \delta_5 \Delta Pn_{it} + \Delta u_{it}. \quad (6)$$

Note that if P_t is uncorrelated with u_s for all $t, s = 1, 2, \dots, T$, one gets the following moment conditions:

$$E[W_i' u_i] = 0 \quad \text{for} \quad i = 1, \dots, N, \quad (7)$$

where W_i is a block diagonal matrix with $[P_{i1}, P_{i2}, \dots, P_{iT}]$ in each block. Also, i denotes the i -th state and u_i is the disturbance vector for that state with $u_i' = (u_{i1}, u_{i2}, \dots, u_{iT})$. Additional orthogonality restrictions are obtained assuming strict or weak exogeneity of Y_{it} and Pn_{it} . These are also included in W_i' . In this case, we form the matrix $W' = [W_1', \dots, W_N']$ by stacking it for N states. The Arellano and Bond (1991) GMM estimator is obtained by performing GLS on (6) after premultiplying it by W' . Using standard assumptions on the initial observation, one can obtain additional moment conditions that could lead to more efficient GMM estimators, see Ahn and Schmidt (1995), Arellano and Bover (1995) and Blundell and Bond (1998).

It is not clear that more instruments of this nature provide better estimators, especially when N is relatively small ($N=46$) and T is relatively large ($T=30$) by micro panel standards. In fact, Ziliak (1997) showed that GMM may perform better with suboptimal instruments and argued against exploiting all moments in panel data GMM estimation. With $T=30$, there are too many orthogonality conditions to use. Hence, Baltagi and Griffin (2001) consider a subset of these orthogonality restrictions, testing their validity using the Sargan over-identification test.

An alternative method of dealing with predetermined instruments was proposed by Keane and Runkle (1992). Using the forward filtering idea from the time series literature to get

rid of the serial correlation in the disturbances allows one to use the original predetermined set of instruments. The general variance-covariance matrix across time $\sum$ can be estimated using the first-differenced 2SLS residuals obtained from (6), i.e.,

$$\widehat{\sum}_{FD} = \sum_{i=1}^N \widehat{u}_{i,FD2SLS} \widehat{u}_{i,FD2SLS}' / N, \quad (8)$$

where $\widehat{u}_{i,FD2SLS}$ is the $(T-1) \times 1$ vector of first-differenced 2SLS residuals for state i . Then one premultiplies (6) by Cholesky's decomposition of the inverse of this variance-covariance matrix $\widehat{\sum}_{FD}^{-1}$ replicated for each cross-section, i.e., by $Q = I_N \otimes P$, where $P = \text{Cholesky of } \widehat{\sum}_{FD}^{-1}$. Finally, 2SLS is performed on the forward filtered model using the original set of instruments. This method of estimating dynamic panel data models has been shown to perform well in Monte Carlo experiments by Ziliak (1997). This estimator, like the GMM estimator, is consistent and has the advantage of being able to account for any arbitrary serial correlation as implied by equation (3).

Baltagi and Griffin (2001) compare FE2SLS with GMM and FD2SLS-KR for two alternative instrument sets. Following BGM, they computed the fixed effects two-stage least squares (FE2SLS) estimator based on $[P_t, P_n, Y_t]$ and their lagged values. FE2SLS results give more reasonable estimates than BGM's fixed effects 2SLS estimator given in their Table 5, model (i). In BGM, the estimated coefficient of C_{t-1} is negative, the estimated coefficient of P_t is positive and both are insignificant. This is why BGM recommended future price as an IV. Baltagi and Griffin (2001), on the other hand, find a positive and significant coefficient estimate of C_{t-1} confirming that smoking is addictive. Also, a positive and significant coefficient estimate of C_{t+1} , rejecting the myopic model and supporting the forward looking consumer without including future price to the set of instruments. Using the Becker, Grossman and Murphy (1994) formulas used to compute the short-run and long-run price elasticities, one obtains the short-run price elasticity from:

$$dC_t/dP_t = 2\delta_3/[1 - 2\delta_2 + (1 - 4\delta_1\delta_2)^{\frac{1}{2}}],$$

and the long-run price elasticity from:

$$dC_\infty/dP = \delta_3/(1 - \delta_1 - \delta_2).$$

The corresponding elasticities are evaluated at the means of the data. BGM's implied elasticities from their FE2SLS estimator are -0.66 and -0.56, respectively, which are in conflict

with the implications of the theory. These are compared with -0.69 and -1.38, the implied price elasticities from Baltagi and Griffin's (2001) FE2SLS. BGM's implied interest rate estimate (from their FE2SLS estimator) is -225%, whereas the implied interest rate from the FE2SLS estimator in Baltagi and Griffin (2001) is 41%. Despite the fact that the latter estimates are plausible, the standard error is large 1.16, implying an unreliable estimate of this discount rate that includes zero and negative values in its confidence interval.

Baltagi and Griffin (2001) then add P_{t+1} to the set of instruments as in BGM. The estimated coefficients of C_{t-1} are uniformly positive and statistically significant ranging between 0.24 and 0.50. Again, this confirms the addictive effects of smoking. The estimated coefficients of C_{t+1} are also positive and statistically significant ranging between 0.16 and 0.38. The short-run price elasticity at the sample mean ranges between -0.42 and -0.69 whereas the long-run price elasticity ranges between -0.56 and -2.04. All of the estimators yield positive estimates of the interest rate ranging from 32% to 84%. However, these estimates have large standard errors.

To summarize, the Becker, Grossman and Murphy (1994) paper provides empirical support for the rational addiction model that relies on a fixed effects 2SLS estimator with future prices included in the instrument set. Even then, the estimates of the interest rate were positive but seemed implausibly large, ranging from 56% to 223%. Exclusion of P_{t+1} from the set of instruments yielded negative interest rates and perverse signs on lagged consumption and own price. Furthermore, Hausman's test for the consistency of instrumental variables estimators using P_{t+1} as instruments was decisively rejected. Baltagi and Griffin (2001) argue that before the RA model can be widely accepted, plausible and statistically significant estimates of the implied discount rate are needed. Based on BGM and the Baltagi and Griffin (2001) results, aggregate panel data do not seem likely to provide sharp estimates of the discount rate. They conjectured that a more promising approach may be to estimate this RA model with micro-panel data as in Chaloupka (1991) but hopefully with samples much longer than $T=3$. In fact, Chaloupka (1991) used micro data on cigarette consumption from the National Health and Nutrition Examination Survey to estimate a rational addiction model. The data set involved approximately 28,000 individuals between the years 1976-1980. Let us now turn to other empirical applications of the rational addiction model.

Olekalns and Bardsley (1996) test rational addiction to caffeine using annual (time series) data on coffee consumption for the U.S. over the period 1967-1992. Imposing an annual discount rate of 0.10, they find support for the rational addiction theory with significant coefficients on lagged and future consumption. However, their long-run price elasticity has

the wrong sign. Imposing prior, so called, reasonable estimates of the discount rate, is not a legitimate, nor scientifically acceptable, answer to the problem of negative estimates of the discount rate. In this day and age of abundant panel micro and macro panel data, one has to estimate this discount rate from the model.

Grossman, Chaloupka and Sirtalan (1998) used surveys of high school seniors as part of the monitoring of the future research program to test the rational addiction hypothesis for liquor consumption. Consumption is measured as the number of drinks of alcohol consumed in the past year. The price variable is that of a six-pack of beer. Grossman, Chaloupka and Sirtalan (1998) find support for the rational addiction theory rather than the myopic theory of addiction. They report negative and significant price effects, positive and significant future consumption effects, and a long-run price elasticity that is approximately 60% larger than the short-run price elasticity. However, Grossman, Chaloupka and Sirtalan (1998, p.46) report that their estimates are not fully consistent with rational addiction because their estimates of the discount factor were negative and implausibly high, yielding interest rates in the range of -20% to -60%. They conclude that these results, along with the detailed analysis of Becker, Grossman and Murphy (1994), suggest that the data on alcohol consumption or cigarette smoking are not rich enough to pin down the discount factor with precision, even if the rational addiction model is accepted. Their long-run price elasticity estimates range from -0.26 to -1.26 , with an average of -0.65 , while their short-run price elasticity ranges from -0.18 to -0.86 , with an average of -0.41 .

Bentzen, Eriksson and Smith (1999) used time series data over the period 1960-1994 to test the rational addiction model for alcohol consumption across the four Nordic countries: Denmark, Finland, Norway and Sweden. Liu, et al. (1999) use annual time series data from Taiwan over the period 1914-1942 to estimate a rational addiction model for opium use. They confirm that opium is addictive, but find no support for the rational addiction model. The short-run and long-run price elasticities for opium are estimated at -0.48 and -1.38 , respectively. Auld and Grootendorst (2004) critique the empirical application of rational addiction models using aggregate data showing that non-addictive commodities, such as milk, eggs and oranges, can be found to be rationally addictive. Brettville-Jensen (1999) examines the rational addiction theory assumptions of interpersonal variation and stability in time preferences based on the behaviour of three groups of people: drug users, non-drug users, and former drug users. This study finds that active injectors of heroin and amphetamine have a higher discount rate than a group of non-users. However, this study also finds that the discount rate among active and former users differs significantly, which is

not supportive of the stability assumption. This questions whether a high time-preference rate leads to addiction or whether the onset of addiction itself alters people's inter-temporal equilibrium.

3.2 The U.S. Liquor Data

Baltagi and Griffin (2002) used annual per capita distilled spirits consumption for 42 American states over 36 years (1959-1994). Their results find some of the implications of the rational addiction hypothesis satisfied for liquor consumption in the U.S. However, these results are sensitive to the assumption of homogeneity across states and suffer from unreasonable estimates of the discount rate. For their empirical implementation, C_{it} is per capita consumption of liquor (measured in gallons of distilled spirits per head) by persons drinking age (16 years and older). P_{it} is the average retail price of 750 ml of Seagram 7 deflated by the CPI. Y_{it} is real per capita disposable income. Per capita sales of liquor is obtained from the *Distilled Spirits Institute* and the average retail price per 750 ml of Seagram 7 is obtained from various issues of *The Liquor Handbook* and updated using the price of alcoholic beverages from the inter city cost of living index published quarterly by the American Chamber of Commerce Researchers Associates. Per capita disposable income data on a state basis are published in various issues of the *Survey of Current Business*. Population data are obtained from various issues of the *Current Population Reports*. Price deflators are obtained from the *Bureau of Labor Statistics*. P_{nit} denotes the minimum real price of liquor in any neighboring state. The last variable allows for possible "border purchasing effects," found important in explaining why some very low tax states enjoy much higher per capita liquor sales than neighboring states with higher taxes.

Following BGM, both state and time effects are assumed to be fixed. This can be justified given the numerous policy interventions, as well as health warnings and Surgeon General's reports. Government interventions include: (i) the Alcohol Traffic Safety Act of 1983, which provided financial incentives for states to enact and enforce stringent drunk-driving laws. See Kenkel (1993), who reports that between 1981 and 1986, 729 state laws pertaining to drunk driving were enacted. (ii) The Federal Uniform Drinking Age Act of 1984 passed by Congress to pressure all states into raising the minimum legal drinking age to 21. (iii) The introduction of warning labels on all alcoholic beverages warning pregnant women about the dangers of drinking and the public about the dangers of drinking and driving. Similarly, state dummy variables are included to control for state-specific characteristics, including, but not limited to the following: (i) States like Montana, New Mexico and Arizona with Indian reservations

that are allowed to sell tax-exempt liquor, (ii) States like Florida, Texas, Washington and Georgia with tax-exempt military bases, (iii) A state like Utah with a large percentage of Mormons (a religion which forbids drinking). Utah's adult per capita consumption in 1994 was 1.20 gallons. This was less than the national average of 1.82 gallons per adult. (iv) A highly touristic state, such as Nevada with a per capita consumption of liquor in 1994 of 4.68 gallons, which is more than twice the national average of 1.82 gallons.

The Distilled Spirits Council (DISCUS) reports that adult per capita liquor consumption in the U.S. dropped from a high of 3.05 gallons per head in 1974 to 1.97 gallons per head in 1992. In 1992, excise taxes on distilled spirits generated \$7.7 billion in tax revenues in federal, state and local taxes. The Federal excise tax on distilled spirits remained the same from 1951 till 1985 at \$10.50 per proof gallon, when it was increased by 19% to reach \$12.50 per proof gallon. This was increased again in 1991 by 8% to \$13.50 per proof gallon. In fact, Grossman, Chaloupka, Safer and Laixuthai (1994) argue that the real prices on alcoholic beverages have declined by 32% between 1975 and 1990 due to the stability of these taxes. States have also taxed liquor quite differently, ranging in 1983 from \$1.50 per gallon in a low tax state like Maryland to a high of \$6.50 per gallon in Florida. Variations in state tax rates are primarily responsible for substantial interstate variation in liquor prices, especially among neighboring states. Another factor is state-owned versus privately-owned liquor stores. In 1984, apparent per capita consumption of alcohol for persons 14 years and older in New Hampshire was 4.91 gallons, nearly twice the national median of 2.63 gallons per capita. This does not imply that New Hampshire residents are heavy drinkers. Carlson (1985, p.31) reports that "about 55% of New Hampshire's 155 million in annual liquor sales is to out of state tipplers."

Allowing the coefficients of equation (4) to vary from state to state, Baltagi and Griffin (2002) estimated a second difference equation for each state, using time-series data. The effective sample period is 1960-1993. Only 22 states have estimates that yield real roots for the second difference equation. For these state by state heterogeneous time-series estimates, all of the coefficients of C_{t-1} and C_{t+1} are significant and positive rejecting the myopic model in favor of the rational addiction model. Pesaran and Smith (1995) suggest averaging these heterogeneous estimates to obtain a pooled estimator. The implied short-run and long-run liquor price elasticity estimates at the mean of the data are -0.52 and -1.39 , respectively, and the implied interest rate for the mean estimates is -6% . The mean neighboring price effect is positive and significant while the mean real per capita income effect is small and insignificant. Baltagi and Griffin (2002) applied 2SLS using two period lagged and forward

variables which included price, neighboring price and income. The results yielded 20 states with real roots for the second difference equation. For these state by state heterogeneous time series regressions, most of the coefficients of lagged and future consumption are significant and positive, rejecting the myopic model in favor of the rational addiction model. The implied interest rate at the mean of the data is -15.7% and the implied short and long-run liquor price elasticity estimates at the mean are -0.51 and -1.38 , respectively. They also performed the Breusch and Godfrey LM test for first order serial correlation across 42 states. This yields a rejection of the null of no serial correlation in 41 out of 42 cases at the 5% level. Recently, Auld and Grootendorst (2004) argued that time series data are generally insufficient to differentiate true rational addiction from serial correlation.

Allowing the coefficients of equation (4) to vary from year to year, they obtained a heterogeneous second difference equation for each time period. Baltagi and Griffin (2002) estimated these heterogeneous equations using cross-section data for 42 states. Only 21 years have estimates that yielded real roots for the second difference equation. All of the coefficients of lagged and future consumption for these year by year heterogeneous cross-section estimates, are significant and positive. They support the rational addiction model rather than the myopic model. The Pesaran and Smith (1995) average pooled estimator yield short and long run liquor price elasticity estimates at the mean of the data at -0.28 and -1.45 , respectively, and the implied interest rate for the mean estimates is -1.4% . They applied 2SLS using two period lagged and forward exogenous variables, which include price, neighboring price and income. The results yield 24 years with real roots for the second difference equation. For these year by year heterogeneous cross-section estimates, all of the coefficients of lagged and future consumption are positive and the majority are significant. This evidence is in favor of the rational addiction model and against the myopic model. The implied interest rate for the mean estimates is 1.8% and the implied short and long-run liquor price elasticity estimates at the mean are -0.31 and -1.62 , respectively.

Assuming the coefficients of equation (4) do not vary over time or states one can pool the data and run 2SLS with fixed effects for states and years. Following BGM, the fixed effects 2SLS estimates using P , Pn_t , and Y_t and their lagged and future values as instruments yields coefficient estimates of C_{t-1} and C_{t+1} that are significant and positive, rejecting the myopic model. The implied short and long-run liquor price elasticity estimates at the mean of the data are -0.10 and -1.24 , respectively. The implied interest rate estimate is 150% . The range of short-run and long-run liquor price elasticity estimates were not that sensitive to the homogeneity versus heterogeneity assumption, but the estimates of the interest rate were

very sensitive to this assumption.

The empirical evidence against the rational addiction theory are: (i) The finding of complex roots for the second difference equation for 20 states out of 42 for the heterogeneous state by state model and 12 years out of 34 for the heterogeneous year by year model. (ii) The unreasonable estimates of the discount rate. The implied interest rates were negative for the heterogeneous state by state model, small and positive for the 2SLS heterogeneous year by year model and large and positive for the homogeneous model. However, it should be emphasized that none of these interest rates were significant. The empirical evidence in support of the rational addiction theory were: (i) The significance of future, as well as, lagged consumption for both heterogeneous and homogeneous rational addiction models. Also, (ii) the persistent and robust evidence of smaller short-run, rather than long-run, liquor price elasticities for both homogeneous and heterogeneous models. The policy implications from the latter results are that there is scope for raising revenues from increasing liquor taxes in the short-run. However, these gains in tax revenue are short lived since the long-run elasticity is larger than one. This also means that taxes are an effective weapon as a social control device to reduce consumption of alcohol in the long-run.

3.3 The Russian Longitudinal Monitoring Survey Data

Alcohol consumption in Russia is legendary and has been reported to be the third leading cause of death in the former Soviet Union after heart disease and cancer, see the *Economist* (1989). In 1985, President Mikhail Gorbachev initiated an anti-drinking campaign that reduced the production of vodka and cognac, set the minimum legal drinking age at 21, prohibited the sale of beverages in public places, restricted the hours of sale and the number of sales outlets, increased the price, prohibited advertising, prosecuted home distillers, developed anti-alcohol programs, and introduced a policy of intolerance to drinking in the workplace. A more recent campaign to raise the tax rate on alcohol by 40% in 2000 provoked long lines outside distilleries and prompted regional governments to refuse to implement the new taxes, fearing civil disobedience. Are Russian alcohol consumers rational addicts? Baltagi and Geishecker (2006) used eight rounds of a nationally representative Russian survey, spanning the period 1994-2003, to estimate a rational addiction model for alcohol consumption. This is done in a panel data setting, as well as on a wave by wave basis. They do not find support for the RA model in Russia for women. For men, the results are supportive of several implications of the RA model, but overall, they fail to endorse the

model empirically on grounds of implausible negative estimates of the discount rate. C_{it} is consumption of alcohol (measured in grams of alcohol consumed per day). P_{it} is the real price of alcohol described below. Y_{it} is real household income and in equation (3) additional variables Z_{it} were included covering a vector of demographic characteristics for the i -th individual at time t .

The Baltagi and Geishecker (2006) study is based on phase II of the Russian Longitudinal Monitoring Survey (RLMS). This is a nationally representative survey designed to measure the effects of Russian reforms.² It used rounds 5 to 12 of the RLMS spanning the period 1994-2003. This is a rich data set with detailed information on alcohol consumption, demographics, education, income, health, occupation, and region of residence. For model estimation, Baltagi and Geishecker (2006) restrict the sample to respondents who were at least 18 years old, drank alcohol, and completed at least three successive interviews.

Alcohol consumption involves various types of alcoholic beverages. Each respondent was asked to state how many grams of beer, wine, fortified wine, home-made liquor, vodka and other hard liquor, and other alcohol they usually drank per day over the last 30 days. From this we constructed two different measures of alcohol consumption: the first is a simple additive measure, and the second is a weighted average adjusted for pure alcohol content.³ Self-reported measures of alcohol consumption have their critics; see Midanik (1989) on the validity of such measures. To the extent that there is no stigma attached to drinking in Russia, respondents can be more truthful in their response to this question; see Tekin (2002).

Prices for alcohol came from the community files of the RLMS. Maximum and minimum prices for certain food items and alcohol are sampled at the community level. All prices are transformed into real values using the monthly consumer price index obtained from Goskomstat (Statistics Russia).

Baltagi and Geishecker (2006) used minimum prices for vodka, beer, fortified wine, and table wine to construct a weighted alcohol price measure. Naturally, prices for home made liquor and other types of unspecified alcohol are not known and cannot be incorporated in the alcohol price measure. This is unfortunate, as it is well documented that consumers substitute home made liquor for branded alcohol as alcohol prices rise.

²This survey is coordinated by the Carolina Population Center at the University of North Carolina (<http://www.cpc.unc.edu/projects/rlms>).

³We used 5% alcohol content for beer, 10% for wine, 19% for fortified wine, 45% for home made liquor, 40% for vodka, and 20% for other alcohol. Similar weights were used by Mullahy and Sindelar (1991) and Tekin (2002).

In Russia the official minimum age for purchasing and drinking alcohol is 18 years. However, respondents as young as 14 years reported drinking alcohol. Fifty-three percent of all respondents in our sample drink alcohol. Among men, the frequency of respondents who reported drinking alcohol is 66%, which is significantly higher than the 44% share among women. The profile of a male drinker shows that, on average, male drinkers are older (41) than non-drinkers (39). They are more likely to be married (67% as compared to 54%). They are also less likely to have children (53% as compared to 57%) and less likely to be foreigners (16% as compared to 23%). Controlling for three levels of education, drinkers are less likely to have a primary education (15%) than non-drinkers (25%). Drinkers are more likely to hold higher ranking occupations (like managers, officials, technicians) than non-drinkers (25% as compared to 21%). Men who drink on average have higher real household income than men who do not drink. Unemployment is significantly less prevalent among drinkers (15%) than among non-drinkers (19%). Male drinkers are more likely to have a higher body mass index (24.9) than non-drinkers (24.2), i.e., they are slightly more likely to be overweight. In addition, male drinkers are significantly more likely to smoke than non-drinkers (67% as compared to 43%).

For women, a slightly different picture emerges. On average, women who drink are significantly younger (40) than women who do not drink (48). They are more likely to be married (55% as compared to 45%). They are more likely to have children (54% as compared to 48%) and less likely to be foreigners (14% as compared to 20%). Among three levels of education, the same pattern emerges for women as for men. The frequency of primary education among drinkers is significantly lower (12%) than that among non-drinkers (32%). Also, women who drink are more likely to have higher occupational placement (51% as compared to 46%). Women who drink have significantly higher real household income than woman who do not drink. Like men, unemployment is less widespread among women who drink as compared to women who do not drink (12% as compared to 14%). Unlike men, women who drink are slightly less likely to be overweight than women who do not drink (with body mass index 26.2 compared to 26.9). Like men, women who drink are also significantly more likely to smoke than women who do not drink (20% as compared to 6%).

Turning to the quantity of alcohol that the individual drinks, male drinkers, on average, drink more than twice as much alcohol as female drinkers (887 grams of alcohol per day compared to 413). This remains the case even after we adjust for pure alcohol content (168 grams of alcohol content compared to 67). Since women differ substantially from men in the frequency and amount of alcohol consumed, Baltagi and Geishecker (2006) studied men and

women separately.

Pooled OLS with robust cluster standard errors, for the entire sample, as well as for men and women separately were reported. OLS ignores the endogeneity of lagged and lead consumption, and controls for unobserved heterogeneity only through the inclusion of demographic characteristics for each individual. These include gender, age, marital status, level of education, region of residence, whether the individual has children, whether a foreigner, whether in a top occupation, and whether the individual drinks without eating. All regressions also include time dummies. For the full sample, lead and lagged consumption were significant, rejecting the myopic model in favor of future looking consumers. Price was significant, but income was not. Some of the regional dummies were significant.⁴ The implied interest rate is negative. The short and long-run price elasticities evaluated at the mean are -0.15 and -0.18 and are statistically significant. The results are the same when applied to men only. The sign and significance of the coefficients are the same but the magnitudes are different. For example, the short and long-run price elasticities evaluated at the mean are -0.19 and -0.22, and are statistically significant. For women, price is not significant, and highly educated women drink less. Otherwise, the results differ only in magnitude from those of men.

Next, an instrumental variables (IV) regression allowing for the endogeneity of lead and lagged consumption and instrumenting with lead and lagged prices and income.⁵ Like the OLS estimator, this IV estimator controls for unobserved heterogeneity only through the inclusion of the demographic characteristics described above for each individual. The results for the full sample reject the myopic model in favor of future looking consumers. However, the implied interest rate is negative and insignificant. The short and long-run price elasticities evaluated at the mean are now larger in absolute value (-0.51 and -1.26), but they have very large standard errors. For women, the IV results are not supportive of a rational addiction model. The coefficient of lagged consumption has a negative sign, but it is insignificant. The coefficient of lead consumption is also insignificant. Including dummy variables for each individual in this IV regression, (i.e., applying fixed effects IV, as done by Becker, Grossman and Murphy (1994)), results in an insignificant F-statistic on the joint significance of the individual dummies. For this data set, it seems that controlling for the problem of endogeneity of lead and lagged consumption is more important than controlling for individual heterogeneity. Nevertheless, they report the fixed effects IV regressions. The full sample

⁴Regional variations in ethnic composition and cultural traditions across Russia translate into regional variations in alcohol consumption, see Simpura and Levin (1997).

⁵Individual prices of beer and vodka are used as instruments. The maximum number of lags used is three.

results yield an insignificant coefficient estimate of lagged consumption and a significant coefficient estimate of lead consumption. Price is also significant, while income is not. This rejects the myopic model in favor of future looking consumers, but the implied interest rate is negative. The results are the same for men but not for women. In the latter case, lagged and lead consumption, as well as price and income are insignificant. For women, except for the OLS estimates, the results are not supportive of the rational addiction model.

In order to check the sensitivity of our results, they performed IV estimation with robust standard errors by round. This is reported for rounds 8 and 11. For round 8, lagged consumption is insignificant, but forward consumption is significant for the full sample as well as for men and women. The implied interest rate is negative but insignificant for all cases. For the full sample, the short-run and long-run price elasticities at the sample mean are -0.295 and -0.333, but both have large standard errors. For round 11, both forward and lagged consumption are not significant for the full sample, but forward consumption is significant for men, while lagged consumption is significant for women. The implied interest rate is negative but insignificant for the full sample, as well as for men, but positive and insignificant for women. For the full sample, the short-run and long-run price elasticities at the sample mean are -0.108 and -0.147, but both have large standard errors.⁶

In sum, the results are sensitive to round by round estimation, and to estimation by gender. The pooled IV results suggest that there is no support for rational addiction in Russia among women. For men, the results are supportive of several implications of the RA model, but overall, we fail to endorse the model empirically on grounds of implausible negative estimates of the discount rate.⁷

4 CONCLUSION

The Becker and Murphy (1988) rational addiction theory is elegant and plausible to economists familiar with varying discount rates. It lends itself to a simple empirical model that can be tested with the availability of at least three years of data over several individuals or states. However, empirical application of the RA theory have faced some econometric and data difficulties. At the macro state level panel data, the advantages may be the large

⁶For the results for women in round 11, the roots of the second difference equation given in (3) are not real since $4\beta\theta^2 > 1$.

⁷Also, these results should be tempered by the fact that we did not deal with zero consumption of alcohol which could be due to quitting, starting to drink or measurement error, see Labeaga and Garcia (1996) and Jones and Labeaga (2003).

variation in prices across states. This is due to the differing tax rates among states which account for most of the variation in price. This contrasts to micro panel data on individuals where the variation in price is limited, or hard to come by, and is usually achieved by locating the individual's city or county and obtaining a price for that county or location. This means that the price variation occurs only at the location or county level and not at the individual level. This may hinder proper estimation of price effects and price elasticities important for policy makers attempting to estimate the effect of a tax hike on reducing consumption or raising revenues. One implication of the theory is that long-run price elasticities have to be larger in absolute value than short-run price elasticities. A condition that is generally met in practice with some exceptions. Stability of the second difference equation is important to check and was shown to be a problem for U.S. panel data on state liquor consumption when we run the regressions on a state by state basis or year by year basis. Finding lagged consumption positive and significant seem to be easy to achieve for these addictive commodities, although there were exceptions. Finding future consumption to be positive and significant was also achieved by most empirical studies, but the implied interest rate remain the fly in the ointment. Reasonable and significant estimates of the discount rate seem to be a crucial requirement underlying the theoretical model. Imposing prior, so called reasonable discount rates is not a legitimate scientific answer in this day and age, especially if we are seriously testing the theory.

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