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SOME NEW EVIDENCE ON THE TIMING OF CONSUMPTION
DECISIONS AND ON THEIR GENERATING PROCESS

By Luigi Ermini

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Abstract

While quarterly consumption data are known to be well fitted by an integrated first-order moving average process, $IMA(1,1)$, with a positive coefficient, it is found that monthly consumption data are well fitted by the same type of process, but with a negative coefficient. This sign reversal has three main implications. First, if the random walk hypothesis of consumption behavior is true, then the agents' decision interval must be greater than a month. In particular, this evidence rejects the possibility of continuously taken decisions of Hall's type. Second, quarterly data can be indistinguishably generated by temporal aggregation of either a random walk or an $IMA(1,1)$ process with negative coefficient. Third, if consumption decisions are generated as an $IMA(1,1)$ process at intervals shorter than a month, then the coefficient must be negative. The theoretical effects of temporal aggregation on $IMA(1,1)$ processes are also investigated, and some implications for empirical inference discussed.

Some new evidence on the timing of consumption decisions and
on their generating process (*)

1. Introduction

Working [1960] showed that a random walk process under temporal aggregation (TA) becomes a first-order moving average process in the first differences, IMA(1,1), with a positive coefficient whose magnitude depends on the interval of aggregation. The fact that quarterly U.S. data on consumption of non-durable goods and services are satisfactorily fitted, at least in a univariate framework, by an IMA(1,1) model with positive coefficient has led some authors to conjecture that consumption decisions may indeed be generated by the agents as a random walk, in line with the well known Hall [1978] model, but at a time interval smaller than a quarter; hence the observed quarterly data would be necessarily affected by the TA effect. Christiano et al. [1987], by adopting a continuous-time approach, determined that quarterly consumption data are consistent with the temporal aggregation of Hall-type decisions taken continuously. Ermini [1987], arguing that the "true" consumption decision interval is more plausibly in the vicinity of a month because of certain relationships between the timing of consumption decisions and income receipts, concluded that all reported tests on Hall theory of consumption behavior conducted with quarterly data are likely to be biased towards rejection, and showed that the bias reduces considerably if the null hypothesis is based on the IMA(1,1) representation and not on the random walk. In both investigations the estimated coefficient of the IMA(1,1) model fitted to quarterly consumption data was consistent with the theoretical value predicted by Working.

This paper presents some new empirical evidence indicating that, if the random walk model of consumption is true, then the true consumption decision interval must be strictly greater than a month. As described in Section 2, in fact, it is found that when fitting an IMA(1,1) model to *monthly* data of consumption of non durables and services, the coefficient of the first-order moving average is significantly *negative*. Since a random walk process cannot turn under TA into an IMA(1,1) with a negative coefficient, one implication of this finding is the strong rejection of the continuous-time approach of

(*) I thank Clive J.W. Granger for helpful comments and Mark Kamstra for assistance in computations.

Christiano et al [1987]; if the agents truly make Hall-type decisions continuously, then monthly data should fit an IMA(1,1) model with a positive coefficient as well.

Another implication is the following. Based on the theoretical effects of TA on IMA(1,1) processes presented in the Appendix, it is shown in Section 3 that a monthly IMA(1,1) with negative coefficient can turn under TA into a quarterly IMA(1,1) with positive coefficient. Particularly, it is shown that in some cases an IMA(1,1) with negative coefficient can be transformed into an IMA(1,1) with a positive coefficient consistent with the value that would be obtained, through Working's derivation, from aggregating a random walk. This fact indicates that the same observed quarterly series can be indistinguishably generated from the temporal aggregation of either a random walk model of consumption or an IMA(1,1) model of consumption. This result clearly weakens any empirical inference about the validity of Hall model in the case where the null hypothesis of a IMA(1,1) model is not rejected, as in Christiano et al [1987], against an alternative model.

In Section 3 it is also shown that there exists a specific interval of aggregation that can turn an IMA(1,1) with negative coefficient into a random walk; particularly, for the observed monthly IMA(1,1) this interval of aggregation is two months. This result raises some important questions for economic theory. On the one hand, if Hall model is true, the evidence presented here would indicate that consumers make their decisions at intervals close to two months. In this case the problem is to interpret the observed monthly IMA(1,1). Current life-cycle economic models do not provide insights on how agents behave between consumption decision times. Some results in this direction can be found in Ermini [1987b]. On the other hand, if agents truly make consumption decisions at intervals shorter than two months, then the problem is to suitably modify Hall model so to explain the generation of an IMA(1,1) with a negative coefficient. Life-cycle models of consumption behavior that generate IMA(1,1) processes have been proposed, among others, in Ermini [1987b] and Zin [1987]. However, the conditions under which the coefficient of the moving average component is negative are not clear. Moreover, it is worth mentioning the result in Backus et al [1988] that a positive risk-premium of equities over riskless assets apparently implies a consumption function following an IMA(1,1) process with *positive* coefficient.

2. The empirical results

Let C_t be a flow variable generated at every Δt_d (the decision interval) according to a random walk process, that is $\Delta C_t = \varepsilon_t$, with ε_t a zero-mean white noise process of variance σ_ε^2 . Let \bar{C}_t the (non-overlapping) temporally aggregated version of C_t , that is, for all integers t :

$$\bar{C}_t = \sum_{j=0}^{m-1} C_{mt-j}, \quad (1)$$

with $m\Delta t_d$ the interval of aggregation and m the sampling ratio. It is shown in Working [1960]¹ that $\Delta \bar{C}_t$ is a first-order moving average MA(1), with variance $R_m(0)$, first-lag autocovariance $R_m(1)$ and first-lag autocorrelation ρ_m given by:

$$R_m(0) = \frac{m(2m^2+1)}{3} \sigma_\varepsilon^2, \quad (2)$$

$$R_m(1) = \frac{m(m^2-1)}{6} \sigma_\varepsilon^2, \quad (3)$$

$$\rho_m = \frac{R_m(1)}{R_m(0)} = \frac{m^2-1}{2(2m^2+1)}. \quad (4)$$

For $m = \infty$ (equivalent to C_t being generated in continuous time and to \bar{C}_t being observed at any discrete interval), $\rho_m = 0.25$. Further, for all $m > 1$ ρ_m is strictly positive, which implies that the first-order coefficient of the moving average $\Delta \bar{C}_t = e_t + h e_{t-1}$ is also positive, since $\rho_m = h/(1+h^2)$.

Fig. 1 reports the autocorrelogram of the first differences of US quarterly consumption of services and non-durable goods (per capita, seasonally adjusted, constant 1982 dollars, from 1947-I to 1985-IV²). For consistency of notation, this series will be denoted as \bar{C}_t . The probability value of the first 18 lags to be white noise is 3.7%. By fitting an IMA(1,1) to the data, we get (t-statistics in parenthesis):

$$\Delta \bar{C}_t = 0.06 + e_t + 0.211 e_{t-1}, \quad (5)$$

(0.02) (2.67)

from which the estimated $\beta_m = 0.202$. The 95% confidence interval for the true ρ_m is (0.042, 0.362), so

¹ Working's paper concerns averages of stock variables (specifically, prices); that is, in our notation, $\bar{C}_t = (1/m) \sum_{j=0}^{m-1} C_{mt-j}$. To account for this difference, the expressions (2) and (3) have been suitably modified from the corresponding ones of the original paper.

² Source: Citibank database 1986. The series was obtained (in Citibank notation) as (GCN82 + GCS82)_x(GYDPC8)/(GYD82).

that from (4) all values of m greater than one (including $m = \infty$) are consistent with the data. Further, the probability value of the first 18 autocorrelation lags of the residuals e_t to be white noise rises to 10.5%. Thus, (5) is a "good" model for quarterly consumption, at least in a univariate framework³. In fact, from the apparent significance in Fig. 1 of at least the third-lag autocorrelation, higher order moving average terms should also be included. Fitting an IMA(1,3) to the data, we get for the first, second and third coefficient (t-statistics in parenthesis): 0.168 (2.14), 0.105 (1.32), 0.243 (3.08). (Fitting an IMA(1,6) confirms that the only significant coefficients are the first and the third). The probability value of the first 18 autocorrelation lags of the residuals to be white noise rises to 36.3%. However, with a log-likelihood ratio LR = 5.99 and a value of the central χ^2 distribution with 2 d.f. of 5.99 at 5%, the null IMA(1,1) is not rejected against the alternative IMA(1,3). The IMA(1,1) is also not rejected against more complex univariate ARIMA models, even without penalizing the latter ones for their number of parameters. For example, against an ARIMA(3,1,3) the log-likelihood ratio LR= 11.32 and the value of the central χ^2 distribution with 5 d.f. is 11.07 at the 5% and 12.83 at the 2.5% levels of significance.

Fig. 2 reports the autocorrelogram of the first differences of US *monthly* consumption of services and non durable goods (per capita, seasonally adjusted, constant 1982 dollars, from 1959 to 1985⁴). Not only do we observe a sign reversal of the first-lag autocorrelation, but also a marked insignificance of any other lag (some residuals of seasonality appear at lags 12 and 24). The probability value of the first 18 lags to be white noise is 0.5%. By fitting an IMA(1,1) to the data, we get:

$$\Delta \bar{C}_t = -0.045 + e_t - 0.221 e_{t-1} \quad (6)$$

(-0.004) (-4.03)

The corresponding estimated $\beta_m = -0.211$ and the 95% confidence interval for the true ρ_m is (-0.321, -0.101). Fitting, in analogy to the quarterly case, the IMA(1,3) to the data, we get: -0.234 (-4.17), 0.059 (1.02), 0.036 (0.65); further, the log-likelihood is practically zero. This confirms that monthly data in the first differences exhibit a marked first-order moving average component with a negative coefficient.

³ This model is dominated by bivariate consumption-income representations. In particular, it is rejected against the error-correction model that corresponds to being quarterly income and consumption co-integrated. However, the focus of the paper is not on deriving the best specification for quarterly consumption data, but on evidencing the presence of a significant positive first-lag autocorrelation of its first differences.

⁴ Same source of quarterly data. This series was obtained as (GMCN82 + GMCS82)x(GMYDP8)/(GMYD82).

Thus, the shift from quarterly to monthly data reverses the sign of the MA coefficient, or equivalently of the first-lag autocorrelation of consumption in the first differences.

To check the robustness of this sign reversal, the monthly sample was split in three subsamples of equal length, and an IMA(1,3) fitted to each. In all three cases the first-order coefficient only was significant and significantly negative, thus suggesting that this sign reversal phenomenon indeed reflects a consistent property of monthly consumption data. As a further check, the monthly data were temporally aggregated with $m= 2, 3, 4$, again obtaining a sign reversal.

Finally, the monthly series was subject to the Dickey-Fuller test for unit root and to the test for cointegration with monthly income, obtaining equivalent results to those already known for quarterly data (for instance, as in Engle and Granger [1987]).

3. Temporal aggregation of an IMA(1,1) process

In the Appendix the effects of the temporal aggregation (1) on IMA(1,1) processes are investigated. The main results from the Appendix are:

1. *Effect of TA on the process order:* if C_t is IMA(1,1) then also \bar{C}_t is IMA(1,1) for all $m > 1$.

Let ρ_1 be the first-lag autocorrelation of ΔC_t , and recall that ρ_m is the first-lag autocorrelation of $\Delta \bar{C}_t$. Then:

2. *Effect of TA on the first-lag autocorrelation:*

$$\rho_m = \frac{(m^2-1) + 2(m^2+2) \rho_1}{2(2m^2+1) + 8(m^2-1) \rho_1} \quad (7)$$

for all $m > 1$. For $b = 0$ (and hence $\rho_1 = 0$), (7) reduces to (4), the case of temporal aggregation of random walk processes. Some implications of these expressions are:

(i) $\lim_{m \rightarrow \infty} \rho_m = 0.25$ for all values of ρ_1 .

This is the same limit established by Working [1960] for the random walk process. This result shows, for example, that quarterly data can be consistent with a continuous-time IMA(1,1) process as well.

$$(ii) \max_{m, \rho_1} \rho_m = 0.25.$$

for all values of $m > 1$ and all values of ρ_1 in the admitted interval $(-0.5, 0.5)$. Therefore, if the first-lag autocorrelation of an IMA(1,1) process is greater than 0.25, this process cannot be produced by temporal aggregation of another IMA(1,1) process.

$$(iii) \rho_1 > 0 \Rightarrow \rho_m > 0.$$

This shows that monthly data cannot be the result of temporal aggregation of an IMA(1,1) process with positive coefficient. Thus, if consumption decisions are truly taken at intervals shorter than a month, they cannot follow an IMA(1,1) process with positive coefficient.

$$(iv) \rho_1 = \frac{1-m^2}{2(m^2+2)} \Leftrightarrow \rho_m = 0.$$

This shows that there exist IMA(1,1) processes with negative ρ_1 that at a specific sampling ratio are aggregated into a random walk. For example, the IMA(1,1) with $\rho_1 = -0.25$, which is consistent with the monthly data, turns into a random walk with $m=2$. This shows that, had consumption data been available bimonthly and had it appeared as generated by a random walk, it would be impossible to infer whether agents make decisions bimonthly and thus follow Hall model, or make decisions more frequently, in which case they must follow an IMA(1,1) with negative coefficient.

$$(v) \text{ there exist some } m > 1 \text{ such that } \rho_1 < 0 \Rightarrow \rho_m > 0.$$

In other words, the observed sign reversal phenomenon between monthly and quarterly data is consistent with the theory. In particular, with $m = 3$, $\rho_m = (4+11\rho_1)/(19+32\rho_1)$ so that the 95% confidence interval for the true ρ_1 , $(-0.321, -0.101)$, reported in Section 2 for monthly data, produces a theoretical 95% confidence interval for the true ρ_3 , $(0.054, 0.183)$, which is a subset of, and hence is consistent with, the empirical 95% confidence interval reported in Section 2 for quarterly data.

$$(vi) \text{ there exist a sampling ration } m \text{ greater than one but less than infinite such that } \rho_1 < 0 \Rightarrow \rho_m < 0.$$

That is, the monthly IMA(1,1) with negative coefficient can be the result of temporally aggregated consumption truly generated by an IMA(1,1) with negative coefficient.

4. Conclusions

The analysis of monthly and quarterly consumption data reveals a phenomenon of sign reversal of the first-lag autocorrelation of consumption in the first differences: positive for quarterly data and negative for monthly data. This sign reversal has three main implications. First, if the random walk hypothesis of consumption behavior is true, then the agents' decision interval must be greater than a month. In particular, this results rejects the possibility of continuously made decisions of Hall type. Second, quarterly data can be indistinguishably generated by temporal aggregation of either a random walk or an IMA(1,1) with negative coefficient. Third, if consumption decisions are generated as an IMA(1,1) process at intervals shorter than a month, then the coefficient must be negative.

These implications raise some questions about the possibility of even finding the "true" model that generates consumption behavior. It seems, in fact, that different pairs (model, decision interval) are consistent, under suitable temporal aggregation, with both monthly and quarterly data. Some light on this problem may be shed by research efforts aimed at determining theoretical reasons for the negative coefficient of the monthly IMA(1,1).

Appendix: The effect of temporal aggregation on IMA(1,1) processes.

Let C_t follow the IMA(1,1) process

$$(1-B)C_t = \varepsilon_t + b \varepsilon_{t-1} \quad (1)$$

with B the backward operator (e.g., $BC_t = C_{t-1}$) and ε_t a zero-mean white noise process of variance σ^2_ε .

Let

$$\bar{C}_t = \sum_{j=0}^{m-1} C_{mt-j} = T(B) C_{mt} \quad (2)$$

where $T(B) = \sum_{j=0}^{m-1} B^j$ is the temporal aggregation operator. Then:

$$\Delta \bar{C}_t = \bar{C}_t - \bar{C}_{t-1} = T(B) (C_{mt} - C_{mt-m}) = T(B)(1-B^m) C_{mt} . \quad (3)$$

Multiplying and dividing by $(1-B)$, and considering that $(1-B^m)/(1-B) = T(B)$,

$$\Delta \bar{C}_t = |T(B)|^2 (1-B) C_{mt} = |T(B)|^2 (1+bB) \varepsilon_{mt}. \quad (4)$$

For simplicity of notation, let $\Delta \bar{C}_t = \eta_t$ and $u_t = (1+bB) \varepsilon_t$. Then (4) becomes:

$$\eta_t = |T(B)|^2 u_{mt}. \quad (5)$$

The autocovariances $R_u(\tau)$ of u_t are:

$$R_u(0) = (1+b^2) \sigma_\varepsilon^2 \quad (6)$$

$$R_u(1) = R_u(-1) = b \sigma_\varepsilon^2$$

$$R_u(\tau) = R_u(-\tau) = 0 \text{ for all } \tau > 1,$$

and the autocovariances of η_t from (5) are:

$$R_m(\tau) = (|T(B)|^2)^2 R_u(m\tau), \quad (7)$$

where

$$\begin{aligned} (|T(B)|^2)^2 &= m^2 + 2m \sum_{j=1}^{m-1} (m-j)[B^j + B^{-j}] + \\ &+ \sum_{j=1}^{m-1} \sum_{i=1}^{m-1} (m-j)(m-i)[B^{j+i} + B^{-j-i} + B^{j-i} + B^{i-j}]. \end{aligned} \quad (8)$$

Thus:

(i) *variance*:

$$\begin{aligned} R_m(0) &= (|T(B)|^2)^2 R_u(0) = \\ &= m^2 R_u(0) + 2m \sum_{j=1}^{m-1} (m-j)[R_u(j) + R_u(-j)] + \\ &+ \sum_{j=1}^{m-1} \sum_{i=1}^{m-1} (m-j)(m-i)[R_u(j+i) + R_u(-j-i) + R_u(j-i) + R_u(i-j)]. \end{aligned}$$

From (6) the second summation is equal to $2(m-1)R_u(1)$. With regard to the third RHS term, the double summation of both $R_u(j+i)$ and $R_u(-j-i)$ is zero; the double summation of both $R_u(j-i)$ and $R_u(i-j)$ is equal to

$$\sum_{j=1}^{m-1} (m-j)^2 R_u(0) + \sum_{j=2}^{m-1} (m-j)(m-j+1) R_u(1) + \sum_{j=1}^{m-2} (m-j)(m-j-1) R_u(-1).$$

Since the second and third terms of this expression are equal, the third RHS term of the variance becomes for $m > 2$:

$$2 \sum_{j=1}^{m-1} (m-j)^2 R_u(0) + 4 \sum_{j=2}^{m-1} (m-j)(m-j+1) R_u(1),$$

and for $m = 2$ it simply equals $2 \sum_{j=1}^{m-1} (m-j)^2 R_u(0)$. Hence, collecting terms:

$$R_m(0) = R_u(0) \frac{2m^3+m}{3} + R_u(1) \frac{4(m^3-m)}{3} \quad \text{for } m > 2$$

$$R_2(0) = R_u(0) \frac{2m^3+m}{3} + R_u(1) 4m(m-1).$$

Letting $\rho_1 = R_u(1)/R_u(0)$ be the first-lag autocorrelation of the first differences of the original series C_t , and recalling (6), we get:

$$R_m(0) = (1+b^2) \sigma_e^2 \left[\frac{m(2m^2+1)}{3} + \frac{4m(m^2-1)}{3} \rho_1 \right] \quad \text{for } m > 2 \quad (9)$$

$$R_2(0) = (1+b^2) \sigma_e^2 \left[\frac{m(2m^2+1)}{3} + 4m(m-1) \rho_1 \right]. \quad (10)$$

For $b = 0$ (and hence $\rho_1 = 0$), (9) and (10) reduce to (2) of Section 1, the case of temporal aggregation of a random walk.

(ii) *first-lag autocovariance:*

$$\begin{aligned} R_m(1) &= (1T(B)1^2)^2 R_u(m) = \\ &= m^2 R_u(m) + 2m \sum_{j=1}^{m-1} (m-j) [R_u(m+j) + R_u(m-j)] + \\ &+ \sum_{j=1}^{m-1} \sum_{i=1}^{m-1} (m-j)(m-i) [R_u(m+j+i) + R_u(m-j-i) + R_u(m+j-i) + R_u(m-j+i)]. \end{aligned}$$

From (6), the first RHS term is zero; in the second RHS term the summation of $R_u(m+j)$ is zero and the summation of $R_u(m-j)$ is equal to $R_u(1)$. In the third RHS term the double summation of $R_u(m+j+i)$, $R_u(m+j-i)$ and $R_u(m-j+i)$ is zero, while the double summation of $R_u(m-j-i)$ is:

$$\sum_{j=1}^{m-1} (m-j)j R_u(0) + \sum_{j=1}^{m-2} (m-j)(j+1) R_u(1) + \sum_{j=2}^{m-1} (m-j)(j-i) R_u(-1).$$

For $m = 2$ only the first term of this expression is different from zero. Since $\sum_{j=1}^{m-2} (m-j)(j+1) =$

$\sum_{j=2}^{m-1} (m-j+1)j$, rearranging terms, and using (6) and the definition of ρ_1 , we obtain:

$$R_m(1) = (1+b^2) \sigma_e^2 \left[\frac{m(m^2-1)}{6} + \frac{m(m^2+2)}{3} \rho_1 \right] \quad \text{for } m > 2 \quad (11)$$

$$R_2(1) = (1 + b^2) \sigma_e^2 \left[\frac{m(m^2-1)}{6} + 2m \rho_1 \right] \quad (12)$$

For $b=0$, (11) and (12) reduce to (3) of Section 1. Finally, let $\rho_m = R_m(1)/R_m(0)$ be the first-lag auto-correlation of the first differences of the temporally aggregated series \bar{C}_t . Then:

$$\rho_m = \frac{(m^2-1) + 2(m^2+2) \rho_1}{2(2m^2+1) + 8(m^2-1) \rho_1} \quad \text{for } m > 2 \quad (13)$$

$$\rho_2 = \frac{(m^2-1) + 12\rho_1}{2(2m^2+1) + 24(m-1)\rho_1} = \frac{1 + 4\rho_1}{6 + 8\rho_1} \quad (14)$$

Note that the value $\rho_2 = (1+4\rho_1)/(6+8\rho_1)$ in (14) can also be obtained directly from (13) with $m = 2$.

Finally, with regard to the effect of temporal aggregation on the model order, the proposition that an IMA(1,1) under TA remains IMA(1,1) for all m greater than one can be easily verified from $R_m(\tau) = 0$ for all lags $\tau > 1$. This result is equivalently obtained from the formula reported, for example, in Weiss [1984].

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Figure 1. - Autocorrelogram of first differences of quarterly consumption

LAG	COVARIANCE	CORRELATION	-1	9	8	7	6	5	4	3	2	1	0	1	2	3	4	5	6	7	8	9	1	
0	1208.81	1.00000																						*****
1	263.58	0.21805												.										****
2	122.537	0.10137												.										**
3	281.008	0.23247												.										*****
4	102.692	0.08495												.										**
5	-145.73	-0.12055												.	**									.
6	29.2349	0.02418												.										.
7	97.9367	0.08102												.										**
8	-157.39	-0.13020												.	***									.
9	-6.0069	-0.00497												.										.
10	-17.235	-0.01426												.										.
11	0.66905	0.00055												.										.
12	45.2551	0.03744												.										*
13	-96.109	-0.07951												.	**									.
14	-26.943	-0.02229												.										.
15	26.6412	0.02204												.										.
16	-44.137	-0.03651												.	*									.
17	90.1093	0.07454												.										*
18	129.381	0.10703												.										**
19	-82.282	-0.06807												.	*									.
20	166.69	0.13789												.	***									.
21	61.1987	0.05063												.										*
22	70.1732	0.05805												.										*
23	-79.492	-0.06576												.	*									.
24	98.5961	0.08156												.										**

'.' marks two standard errors

Figure 2. - Autocorrelogram of first differences of monthly consumption

LAG	COVARIANCE	CORRELATION	-1	9	8	7	6	5	4	3	2	1	0	1	2	3	4	5	6	7	8	9	1		
0	870.568	1.00000																							
1	-201.63	-0.23161												*****		.									
2	34.0358	0.03910												.		*	.								
3	57.9028	0.06651												.		*	.								
4	-58,763	-0.06750												.		*	.								
5	37.6007	0.04319												.		*	.								
6	8.58664	0.00986												.		.	.								
7	19.1208	0.02196												.		.	.								
8	59.8783	0.06878												.		*	.								
9	-14.648	-0.01683												.		.	.								
10	68.9066	0.07915												.		**	.								
11	93.2503	0.10711												.		**	.								
12	-34.929	-0.04012												.		*	.								
13	13.4403	0.01544												.		.	.								
14	-55.593	-0.06386												.		*	.								
15	-31.203	-0.03584												.		*	.								
16	-13.672	-0.01571												.		.	.								
17	-90.591	-0.10406												.		**	.								
18	62.1711	0.07141												.		*	.								
19	-4.425	-0.00508												.		.	.								
20	78.6705	0.09037												.		**	.								
21	-13.03	-0.01497												.		.	.								
22	-50.98	-0.05856												.		*	.								
23	43.5516	0.05003												.		*	.								
24	-188.12	-0.21609												.		**	.								

'.' marks two standard errors

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