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ECONOMETRIC MODELLING OF CONSUMERS' EXPENDITURE IN ARGENTINA
1977(1) - 1990(4)

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**ECONOMETRIC MODELLING OF CONSUMERS' EXPENDITURE IN ARGENTINA 1977(1) - 1990(4)**

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## ABSTRACT

In this paper we develop a parsimonious conditional model of consumers' expenditure in Argentina for 1977-1990. Our set of information contains, besides expenditure, income and inflation. All variables are quarterly measured.

In specifying our model, we follow the econometric approach known as "general-to-particular" methodology. We address issues of empirical model design and evaluation, co-integration and exogeneity.

The empirical model is robust and has constant, well determined parameter estimates. These features are specially remarkable in a period of great macroeconomic instability which includes two hyperinflationary processes.

The validity of our conditional model supports the adequacy of the usual assertions about two prominent features of the Argentine economy: i) the great importance of current income in explaining consumers' expenditure; ii) the transmission of the volatility of inflation to the volatility of aggregate demand via consumers' expenditure.

ECONOMETRIC MODELLING OF CONSUMERS' EXPENDITURE IN ARGENTINA

1977(1) - 1990(4) (#)

Sebastián Galiani and Marcelo Sánchez

## 1. Introduction.

In this paper we estimate a conditional model for consumers' expenditure in Argentina. We follow the general-to-particular econometric methodology (see Hendry and Richard, 1982, 1983; and Hendry et al., 1984). In this approach, econometric models are viewed as successive simplifications of the underlying real world process generating data. Test statistics become a very important part of the selection criteria since the error processes on empirical models are derived via the specification of the model and its associated estimation procedure. This approach permits the researcher to choose an empirical model which both embodies the economic theory and allows for the presence of any significant factors not fully specified by the postulated theoretical model, such as matching its lag reactions to the autocorrelation structure of the associated observed time-series data.

Economic theory suggests us the inclusion of many variables in the consumption function, such as income, wealth, the rate of inflation, the rates

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of return of financial assets and the income distribution.[1] To construct our empirical model of consumers' expenditure we have considered not only the general prescriptions of the theoretical analysis, but also the idyosincratic characteristics of the Argentine economy and the restrictions on the available data.

In this paper we estimate a consumption function for the period 1977(1)-1990(4). One salient feature of the Argentine economy during this period was the presence of great macroeconomic instability which can be perceived in the huge volatility of the rate of inflation, consumption and the activity level, among other variables. Such macroeconomic instability was characterized by important fluctuations in relative prices which induced revisions of decisions (among them those about consumption) due to changes in perceived wealth.

The persistence of high inflation standards led the economy to adapt its contract network. This phenomenon manifested itself in the indexation of nominal contracts to some inflation index. The set of adaptation mechanisms and the behavior developed by the agents in this context are known as "high inflation regime", which provides a great volatility to both the rate of inflation and the real variables. In particular, the shortening of the contract period and the faster revision of expectations implied a greater impact of the shocks that the economy experienced, specially since the mid-seventies. A remarkable characteristic of this process was dollarization as a search of alternative liquid assets which did not depreciate with the acceleration of inflation.

In spite of generating an active search for information, greater rates of inflation induced a higher degree of uncertainty than that prevailing in periods of stability.[2]

An important determinant of consumption, as emphasized by the life-cycle theory, is wealth. A good approximation like that used in Brodin and Nymoen (1992) can not be obtained from the information available for Argentina. The

[1]For recent surveys of theoretical literature about the consumption function, see Abel (1990), Hall (1989) and, for a more comprehensive analysis, Deaton (1992).

[2]For a general account of the evolution of the Argentine economy under this high inflation context, see -among others- Frenkel (1989) and Heymann (1986).

real money stock may be sometimes an useful <u>proxy</u> of wealth, but we think this is not the case in a country, like Argentina during the period of analysis, which has experienced a process of dollarization whose magnitude we are not able to measure.

Most theories of the consumption function (such as those based on the permanent income and the life-cycle hypothesis) were formulated to reconcile a low short-run marginal propensity to consume with the relative stability claimed for the average propensity to consume over medium to long data periods. Davidson <u>et al.</u> (1978) implemented the lag mechanisms -postulated by those theories- which mediate the response of consumption to changes in income. They used an error-correction mechanism in the modelling of consumers' expenditure as a way of capturing adjustments in consumption which depended not on the level of income, but on the extent to which the latter deviated from an equilibrium relationship with the former (see Banerjee <u>et al.</u>, 1993).

Without suggesting the nonexistence of a long-run equilibrium relationship between consumption and income, the Argentine experience during the period under research indicates that it is likely for current income to be an important explanatory variable of consumers' expenditure due to the existence of liquidity constraints and the effects of changes in actual income on expected wealth.[3] According to Deaton (1992), the behavior with liquidity constraints can look less like the behavior of consumers under the permanent income or life-cycle hypothesis than like that of prudent economic agents in an uncertain world. As regards the effects of changes in actual income on expected wealth, Leijonhufvud (1973) explains that, when a change in the level of current income impinges on a household whose balances of cash and other liquid assets are low, wealth could be reduced via an increase in the implicit rate of discount, and so could consumption. If such situation generalizes, the system exhibits effective demand failure. With cash constraints operative, further disturbances will trigger deviation-amplifying multiplier processes.

As far as inflation is concerned, there are several channels through which it may influence consumers' expenditure. Deaton (1977) emphasizes the

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[3]As Granger (1993) states, econometric techniques presents difficulties in treating long-run relationships because new information accumulates very slowly. He also emphasizes the importance of learning about the short-run trajectories.

adverse effects on expenditure of unanticipated inflation. It has also been mentioned the adverse redistributive effects of inflation -and specially the inflation tax- on private consumption.[4]

Owing to the uncertain context engendered by Argentine high inflation standards, there was a shortening of the decision horizon which was more dramatic in financial markets. These markets behave in a different way under a high inflation regime as compared to a low inflation one. Under a high inflation regime it was observed that credit tended to vanish. Conversely, during periods of stabilization one assisted to a credit expansion which encouraged (specially durable) consumers' expenditure. This phenomenon motivates the inclusion -in our econometric model of consumers' expenditure- of lagged rates of inflation which capture the negative effects of inflation on the availability of credit via regime shifts.

Finally, important income and wealth effects resulted in Argentina from the widespread automatic indexation of wages to past inflation. This mechanism made it possible for the acceleration (desacceleration) in the rate of inflation to have a negative (positive) impact on the purchasing-power of wages.

The rest of this paper is organized as follows. Section 2 describes the data set and analyzes its properties. Section 3 contains our estimate of a consumption function for Argentina during the period of analysis. In section 4 we discuss exogeneity and invariance issues. Finally, conclusions are stated in Section 5.

## 2. Data set and its properties.[5]

The data set is quarterly, not seasonally adjusted, for the Argentine economy over the period 1977(1)-1990(4). There is no available quarterly information about consumers' expenditure (c) and income (y) from 1991 on.

[4]In a classical paper, Hendry and von Ungern-Sternberg (1981) remark the mis-measurement of real income due to ignoring the impact of inflation on the real value of nominal assets.

[5]The econometrics and graphs were done with PCGIVE 7 (see Doornik and Hendry, 1992).

Moreover, the variables have been logarithmically transformed. To operationalize **c** and **y** we have used the national accounts statistics published by the Central Bank of Argentina. In them, **c** is obtained as a residual from the other GDP components; it includes private and goverment consumption expenditures in goods and services.[6] With respect to income we have used the GDP.[7] Both variables are measured at 1970 market prices. As there is no available quarterly information of the implicit deflator of **c**, the rate of inflation ($\pi$) was calculated as the rate of change of the CPI published by the Instituto Nacional de Estadísticas y Censos.

First some basic properties of the data will be considered. Figure 2.1 shows the behavior of **c** and **y** over time and Figure 2.2 graphs their annual first differences, $\Delta_4 c_t = c_t - c_{t-4}$ and $\Delta_4 y_t = y_t - y_{t-4}$. As we can see, **c** shows

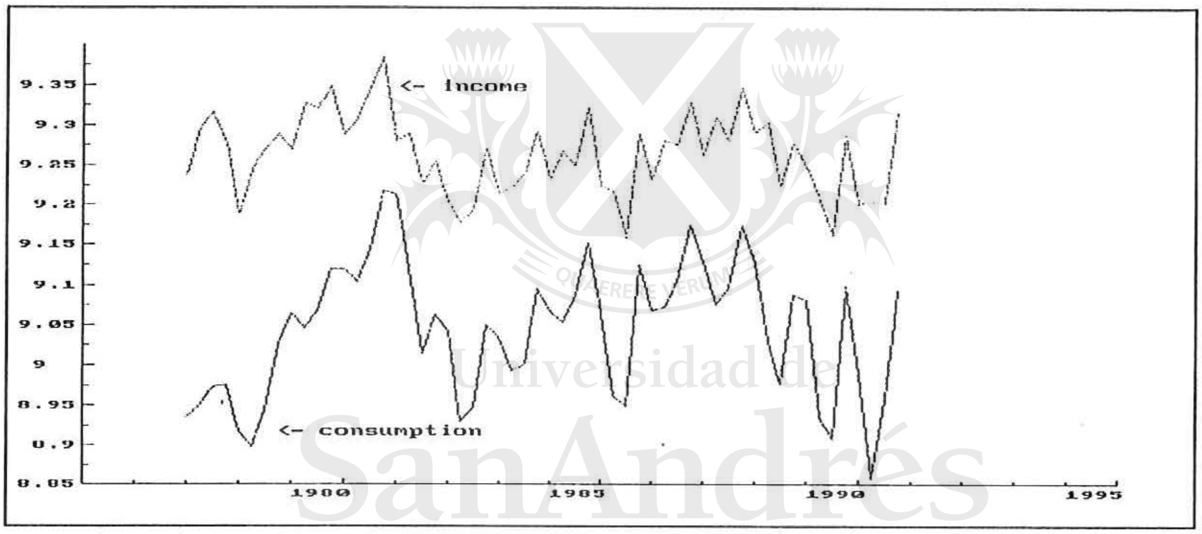


FIG. 2.1. Total consumption expenditure and GDP.

[6]Although it will be desirable to model the data generation process of private consumption, the available annual information permits to conclude that the public consumers' expenditure remained a remarkably constant proportion of c during the period of analysis. This proportion is of slightly above 10% for the period 1961-1986, as estimated from the information about government expenditure provided by Dirección Nacional de Programación Presupuestaria del Ministerio de Economía: Sector público. Esquema de Ahorro-Inversión-Financiamiento: 1961-1986, 1988.

We were not able to obtain reliable information about how the households' expenditure is divided between durable goods, and non-durable goods and services. This is an unsolved problem of all studies of the consumption function in Argentina; see e.g. Leone (1980).

[7]Although to model the data generation process of private consumption we should have a measurement of personal disposable income, it is not available.

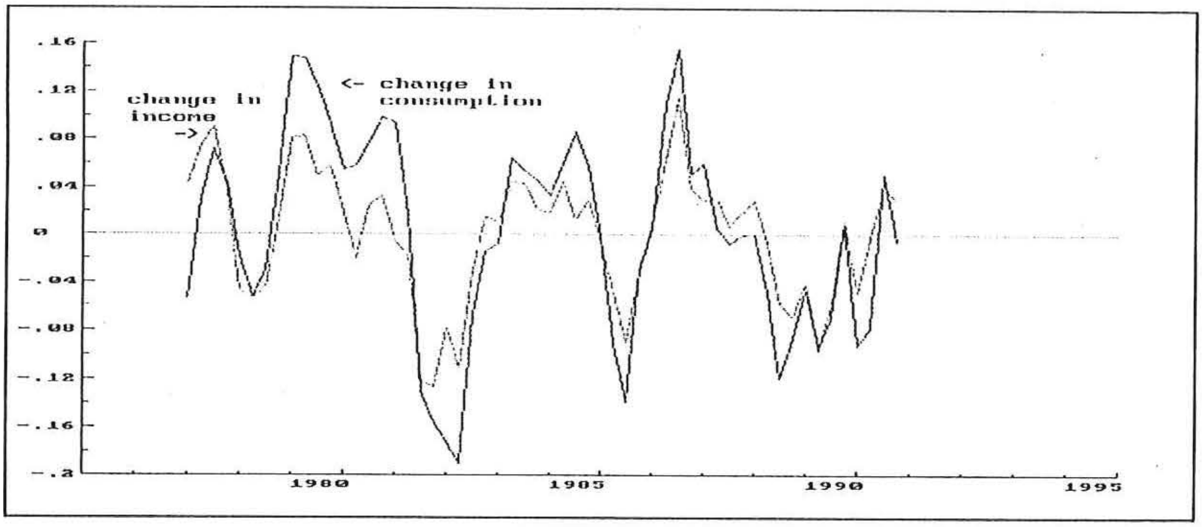


FIG. 2.2. Annual first differences of **c** and **y**.

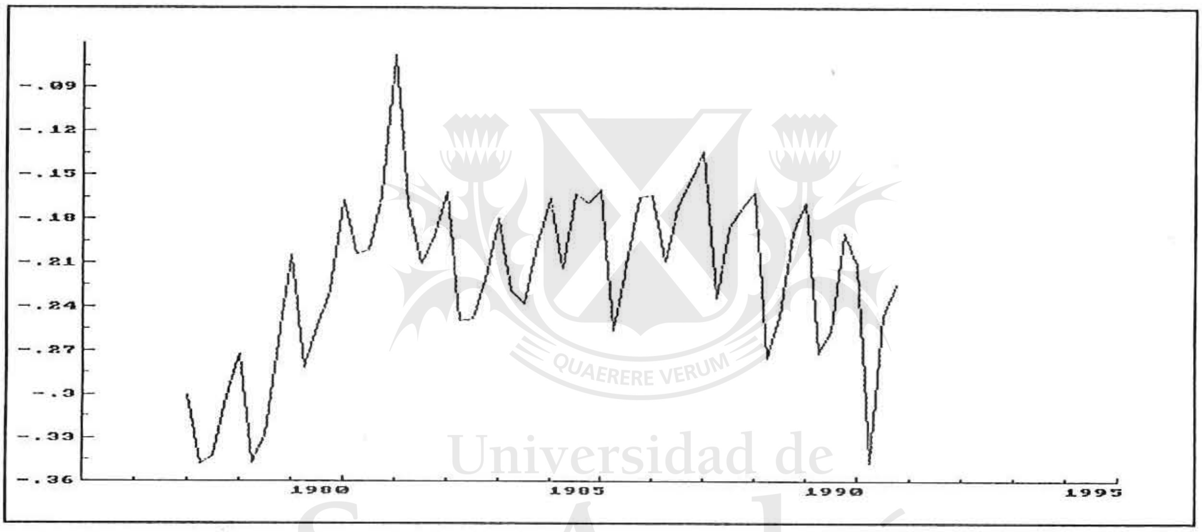


FIG. 2.3. Average propensity to consume.

fluctuations of greater amplitude than **y**.

Those changes in **c** are an unusual phenomenon in stable economies that possess a developed financial market. This fact is not predicted by any theory which supposes that agents optimize intertemporally and are not credit-rationed.[8] Additionally, this co-movement of **c** and **y** implies that the average propensity to consume has fluctuated significantly during the period, as shown in Figure 2.3. In fact, this ratio increased from 0.7 at the beginning of the

[8]This assertion does not hold if the current income-elasticity of permanent income is greater than one. For a good treatment of the volatility of consumption, see Deaton (1992, chapter IV).

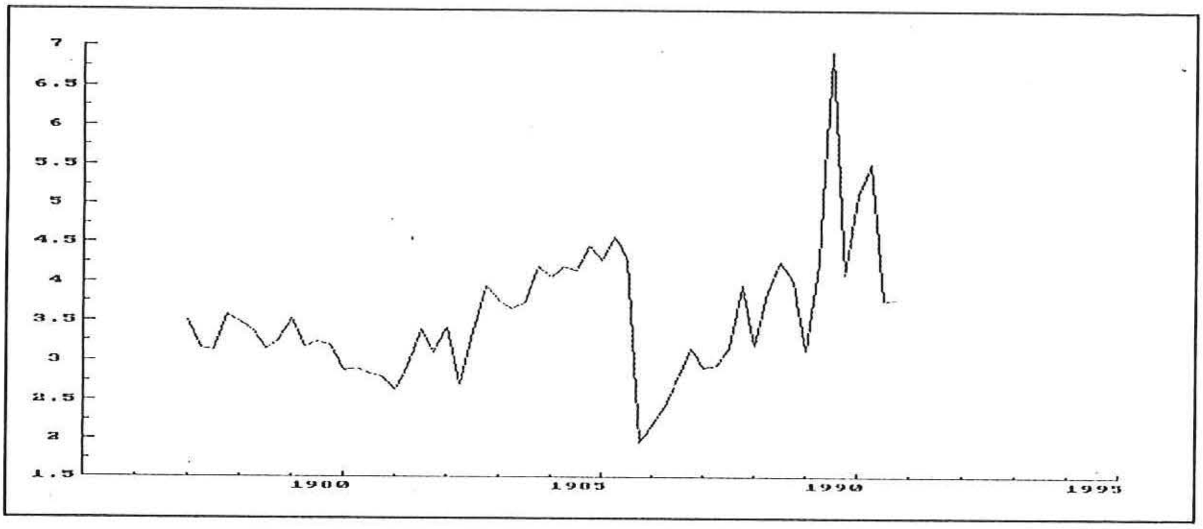


FIG. 2.4. The rate of inflation.

series to the peak of over 0.9 in 1981(1), then fell (but not continuously) to 0.7 again at the end of the series. This great variability of the average propensity to consume implies that a model explaining it shoud be well prepared to account for future fluctuations of consumers' expenditure.

As can be seen, **c** increased more than **y** during two stabilization periods known as the "tablita" and the Austral Plan, beginning in December 1978 and June 1985, respectively. Both plans induced disinflations and, associated to this, a rise in the purchasing power of wages and an expansion of consumers' credit.Conversely, the c-y ratio declined in the period after the collapse of the "tablita" (March 1981), including the occurrence of the debt crisis in 1982 (after which the country was dramatically credit-rationed), and after the end of the Austral Plan. During these episodes the rate of inflation accelerated.

The comments in the precedent paragraph suggest us the convenience of including $\pi$, in addition to **y**, as an explanatory variable of **c**. This can be checked by comparing Figures 2.2, 2.3 and 2.4. The latter shows the evolution of the rate of inflation.

It is not obvious from Figure 2.1 that **c** and **y** are cointegrated. However, the existence of a long-run relationship between these variables is an important feature of both theoretical and empirical models. Prior to the substantive modelling exercise (carried out in section 3), we follow a commonly used approach to test whether there exists such a cointegration relationship. In testing for it, we shall evaluate if, in case the individual series were I(1), there exists some non-trivial function of them which is I(0).

To examine the order of integration of the variables we test for the existence of a unit root in the univariate representations of the individual series. Table 1 shows the values of the Sargan-Bhargava (Durbin-Watson) statistic and the augmented Dickey-Fuller statistic (ADF) for **c**, **y** and $\pi$ for the period 1970-1990. The number of lags used are just the necessary to remove residual serial correlation from the univariate representations. The tests for **c** and **y** do not reject the hypothesis for a unit root at the 1% level of confidence. Instead, in the case of $\pi$ the evidence is mixed. The ADF statistic lets us conclude that $\pi$ is a stationary series, but the Sargan-Bhargava statistic does not reject the hypothesis for a unit root at the 1% level of confidence.[9]

Table 1
Univariate Statistics for Testing Unit Roots

| Variable | CRDW | ADF(0) | ADF(5) |
|---|---|---|---|
| c | 0.58 | - | -2.54[1] |
| y | 0.73 | - | -2.40[2] |
| $\pi$ | 0.71 | -5.45[3] | - |

Notes: 1. With constant.
2. With constant and seasonals.
3. With constant and trend.

To evaluate if **c** and **y** are cointegrated we estimate the Engle-Granger static regression for them. If they were not, any linear combination of them can drift anywhere, and hence no long-run relationship exists between those

[9]Recursive estimates of the ADF statistic (using monthly data) showed in Ahumada (1992b) lead to conclude that the $\pi$ series is I(0) or I(1) depending on the period. What is more, she demonstrated that the series became explosive around mid-1989. Instead, the same method applied to our data let us see that the $\pi$ series became stationary in 1977(1) and was never non-stationary from then on. The difference between the two results could be related to the frequency with which the data are measured.

series. The result is the following:

(2.1) $$c_t = 0.976 \; y_t$$

$R^2$=0.999 $\sigma$=6% DW=0.86 T=54 k=1 ADF(0)=-2.845

The $R^2$ value, the unit coefficient on $y_t$, and the Dickey-Fuller (ADF) test (using seasonals) provide evidence in favor of cointegration. As we shall see later, our parsimonious model of consumers' expenditure (derived from a general-to-specific approach) will be an error correction mechanism. This fact brings additional support to the existence of an attractor of the temporal trajectories. The steady state predicted by the model will show homogeneity between **c** and **y**.

### 3. A Consumption Function:

In this section we estimate a conditional econometric model of Argentine consumers' expenditure following the general-to-particular approach. In addition to income, we include $\pi$ and seasonals as explanatory variables. Table 2 presents the estimates for an unrestricted autoregressive-distributed lag (ADL) model.

The ADL model in Table 2 shows no departure from the nulls of white-noise and innovation disturbances, as well as of parameter constancy within sample. Recursive Chow tests (provided by the PCGIVE 7 version and not shown here) do not cross the critical values from the F-distribution at the 5% probability level. Furthermore, $\sigma$ is slightly above 2%; it means that any new model will require a similar or smaller residual standard error as a necessary condition for encompassing this model.

This configuration provides us a statistically acceptable point of departure to search for a parsimonious conditional consumption function. However, the static long-run equation corresponding to the model in Table 2 does not show well determined long-run coefficients. The tests on the joint-significance of each variable's lag polinomial permit us to say that **c** and **y** are the only significative variables at the 1% level. On the other hand, one of the main arguments (based on the fact that the public sector appropiates part of personal income via the inflation tax) for including the rate of

Table 2
The Unrestricted Dynamic Model estimates for
Consumer's Expenditure: 1977(1)-1990(4) (*)

Left-hand side variable is $c_t$
Method: OLS. T=56 K=21 $\sigma$=2.39% $R^2$=0.948 DW=2.10
$FAR_{1-1}(1,34)$=3.54 $FAR_{1-4}(4,31)$=1.18 $Chi^2(2)$=0.25
$FARCH_{1-4}(4,27)$=0.27 FRESET(1,34)=0.84
VIT=0.24 JIT=3.25 Chow F(10,25)=0.31

| | j = lag, (quarter for the seasonals) | | | | | |
|---|---|---|---|---|---|---|
| | 0 | 1 | 2 | 3 | 4 | 5 |
| $c_{t-j}$ | - | 0.961<br>(0.182) | -0.294<br>(0.254) | 0.237<br>(0.254) | 0.115<br>(0.241) | -0.141<br>(0.170) |
| $y_{t-j}$ | 0.882<br>(0.140) | -0.745<br>(0.200) | 0.336<br>(0.230) | -0.275<br>(0.243) | -0.262<br>(0.253) | 0.236<br>(0.205) |
| $\pi_{t-j}$ | -0.015<br>(0.006) | 0.015<br>(0.007) | -0.004<br>(0.008) | -0.011<br>(0.007) | 0.013<br>(0.008) | 0.001<br>(0.007) |
| Constant | -0.461<br>(1.461) | | | | | |
| Qj | -0.006<br>(0.028) | -0.069<br>(0.025) | 0.009<br>(0.028) | | | |

(*) Standard errors are below the estimates, $\sigma$ is the percentage residual standard deviation, T the number of observations and k the number of coefficients. $FAR_{i-j}$ tests for *i*th to *j*th order autocorrelation. $Chi^2$ tests for residual normality. $FARCH_{i-j}$ tests for autoregressive conditional heteroscedasticity. FRESET tests for the omitted square of the fitted dependent variable. VIT tests for variance instability and JIT tests for joint instability. Chow F tests for parameter constancy over a ten-quarter forecast period.

'F' and 'Chi' denote distributions under the null and subscripts indicate the relevant alternative.

inflation as a steady-state determinant of private consumption is of no use here because our data of consumers' expenditure include government purchases of goods and services. These facts lead us to prefer the Engle-Granger static regression (2.1) as a long run equilibrium for c.

Looking at the unrestricted model in Table 2, we next eliminate the regressors that do not seem to be significative. Various orthogonalizing transformations were carried out on the model to create growth rates, the error correction term, and so on. Reestimating that model gives:

$$(3.1) \quad \Delta c_t = \underset{(0.016)}{0.002} + \underset{(0.100)}{0.829}\, \Delta y_t - \underset{(0.005)}{0.012}\, \Delta\pi_t - \underset{(0.004)}{0.017}\, \Delta\pi_{t-3}$$

$$- \underset{(0.060)}{0.146}\, (c_{t-2} - y_{t-2}) - \underset{(0.014)}{0.018}\, Q_1 - \underset{(0.009)}{0.105}\, Q_2 - \underset{(0.011)}{0.008}\, Q_3$$

$T=56$ $K=8$ $\sigma=2.28\%$ $R^2=0.918$ $DW=2.03$
$FAR_{1-1}(1,47)=0.01$ $FAR_{1-4}(4,44)=0.81$ $Chi^2(2)=3.52$
$FARCH_{1-4}(4,40)=0.43$ $FHET(11,36)=1.37$ $FRESET(1,47)=0.76$
$VIT=0.18$ $JIT=1.95$ Chow $F(10,38)=0.45$

where FHET tests for heteroscedasticity arising from the squares of the regressors.

The block of 13 restrictions was tested by an overall F-test (FNIN(13,35)=0.66). The model (3.1) also shows no departure from the nulls of white-noise and innovation disturbances, and parameter constancy.

Figures 3.1 and 3.2 show actual outcomes of, and the fitted values from, equation (3.1) over the whole sample and with a ten-quarter forecast sub-period, respectively. The performance is remarkable given the great variability of y and $\pi$ over the period. Note that $\sigma$ is only 2% and smaller than in the unrestricted ADL model.

Figure 3.3 plots the actual and forecast values for $\Delta c$ for 1988(3)-1990(4) with 95% confidence intervals for each forecast. Actual values lie well within the confidence intervals, confirming the validity of the conditional model (3.1).

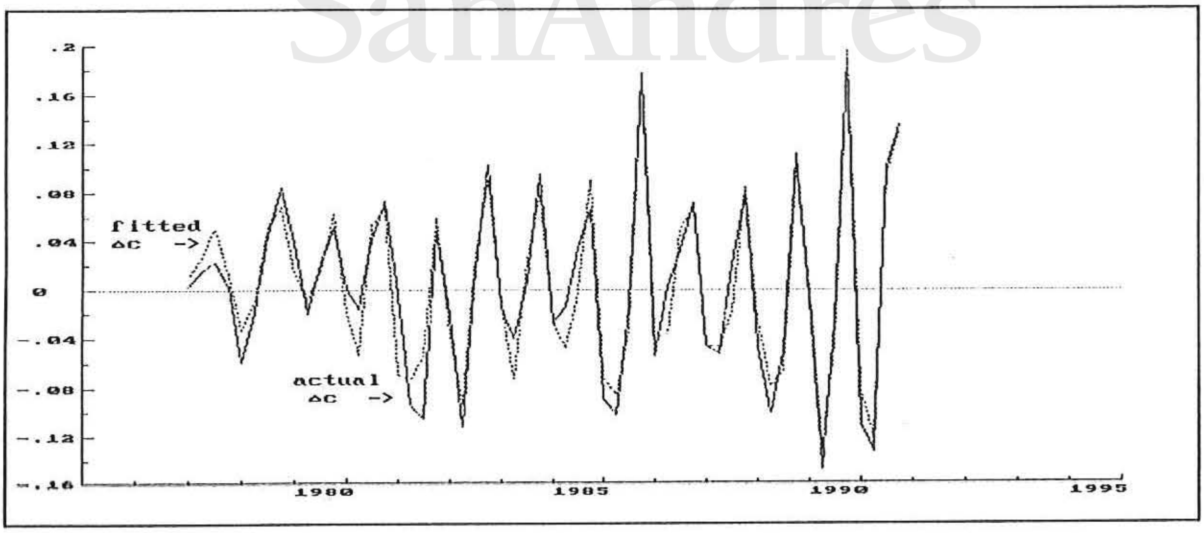


FIG. 3.1. Equation (3.1): Actual outcomes and fitted values over the whole sample.

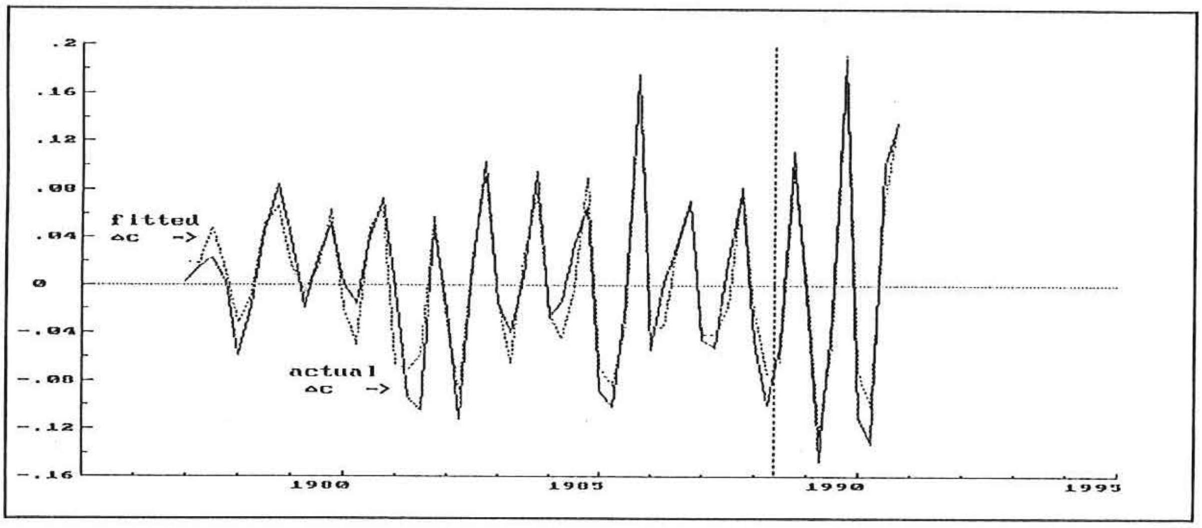


FIG. 3.2. Equation (3.1): Actual outcomes and fitted values with a ten-quarter forecast sub-period.

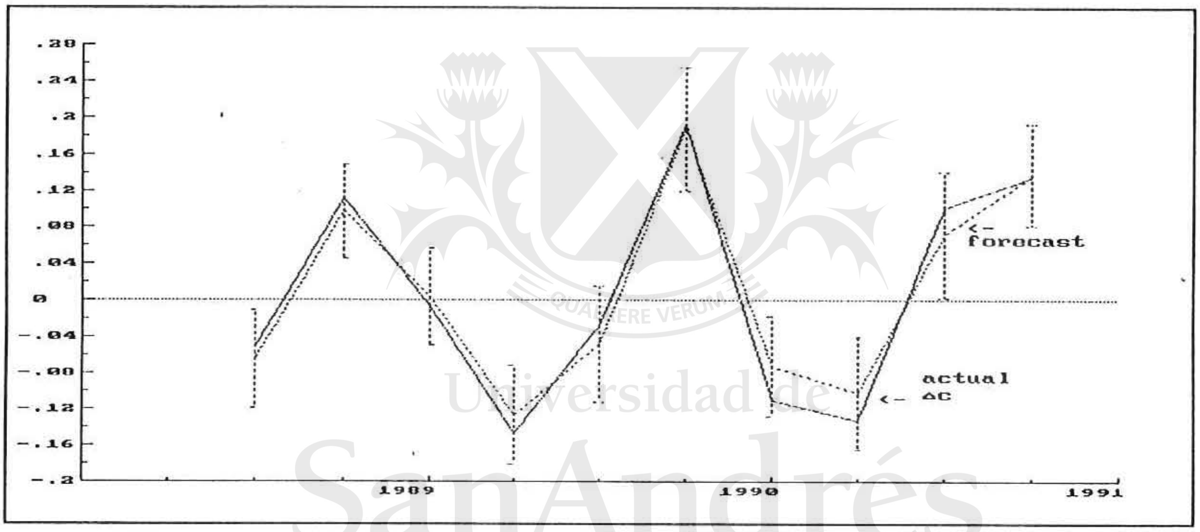


FIG. 3.3. Equation (3.1): One-step ahead forecast values of △c with ± 2 forecast standard errors.

To further evaluate the stability of the model parameters we have conducted several tests based on recursive least squares estimation. The estimated coefficients are constant; most of them, which relate to exogeneity analysis, are shown in Figures 4.1 to 4.3. There is a gain of significance over time for all variables, with the acceleration of inflation becoming significant only in mid-1989. It could lead to think that one may improve the model by simply eliminating $\pi$ as an explanatory variable. However, an ADL model run until 1989(1) without $\pi$ has shown instability. The fact that instability is eliminated by reincluding $\pi$ illustrates the role of this variable in stabilizing the conditional model.

It is worth noting the high impact elasticity for income (0.829). Adjustments are slow: the cumulated multipliers raise the income effect only to 0.875 in the course of the first four quarters. Furthermore, the terms concerning the acceleration of inflation have a negative sign, capturing the adverse effects on consumption mentioned above. Particularly, the presence of $\Delta\pi_{t-3}$ together with $\Delta\pi_t$ could be interpreted as showing the adverse effects of uncertainty on consumers' expenditure via the availability of credit.

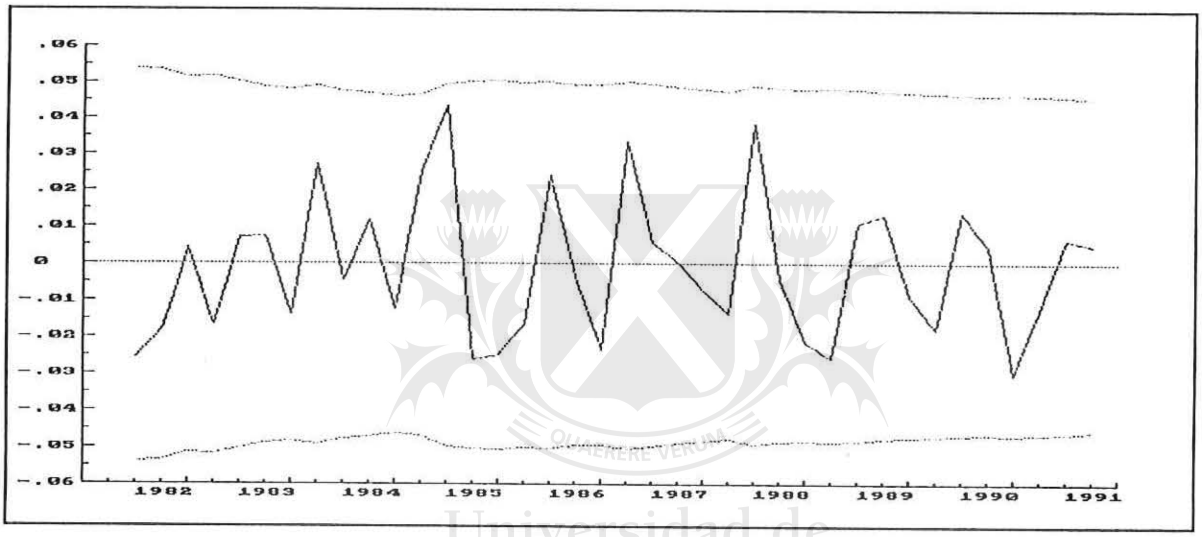


FIG. 3.4. Equation (3.1): One-step residuals and the corresponding calculated equation standard errors.

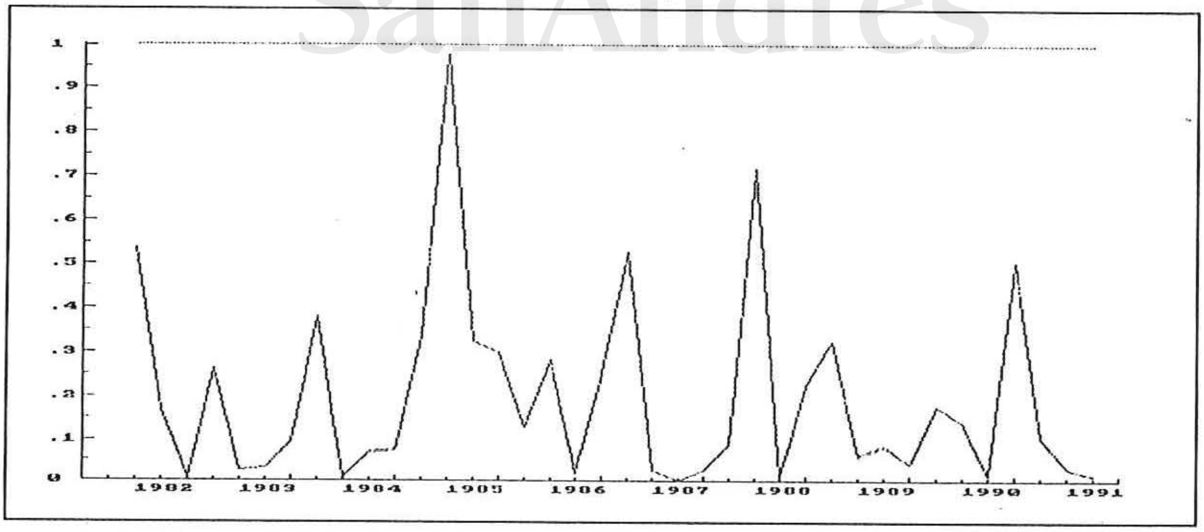


FIG. 3.5. Equation (3.1): Sequence of one-step Chow-statistics.

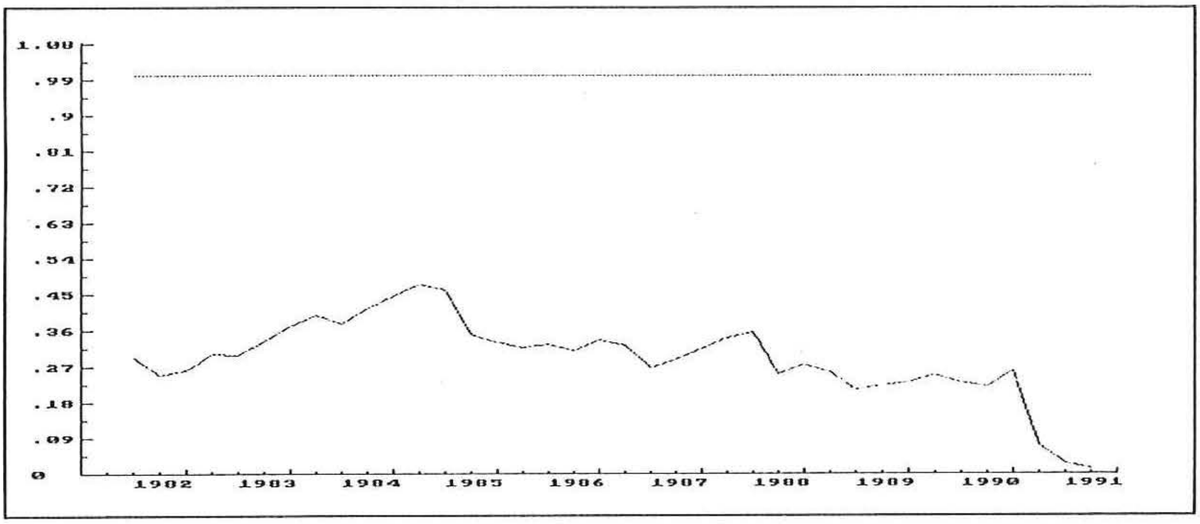


FIG. 3.6. Equation (3.1): Sequence of break-point Chow-statistics.

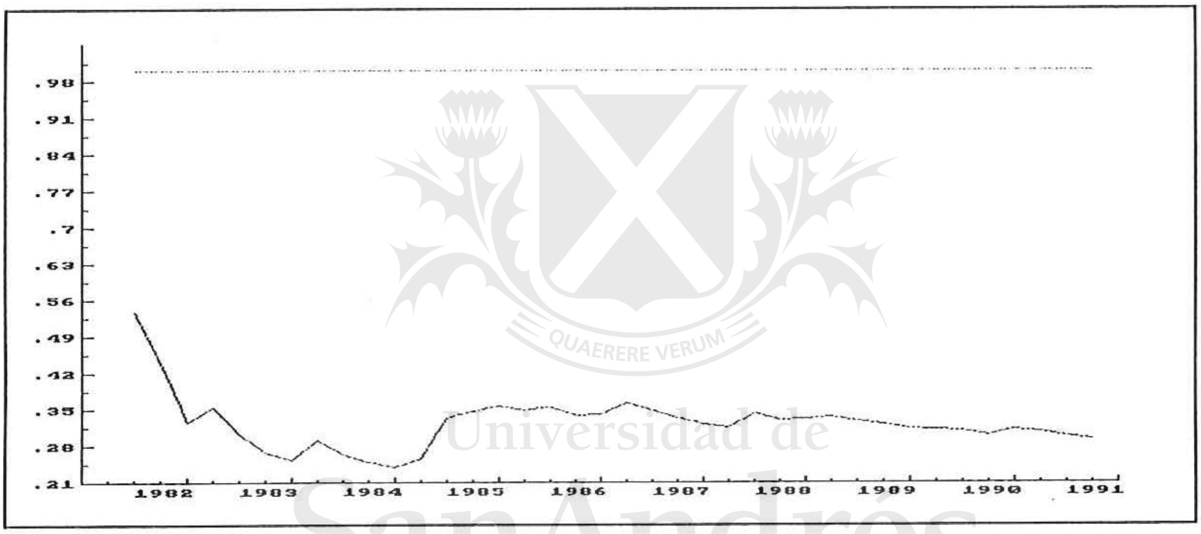


FIG. 3.7. Equation (3.1): Sequence of forecast Chow-statistics.

Figure 3.4 records the sequences of one-step residuals and corresponding calculated equation standard errors. The residual standard error is remarkably stable, given the large changes in data correlation structure. Finally, none of the Chow statistics shown in Figures 3.5 to 3.7 passes the critical values from the F-distribution at the 5% probability level.

The stability obtained suggested us the validity of the conditional model. In order to continue the evaluation of this model, next section further analyzes the issue of exogeneity.

## 4. Exogeneity and Invariance:

The literature considers several definitions of exogeneity, two of which are of special relevance here: weak exogeneity (WE) and super exogeneity (SE) (see Engle et al., 1983). WE sustains conditional inference and requires that the parameters of interest in a conditional model can be efficiently analyzed without specifying the marginal model for the potentially exogenous variables. SE is weak exogeneity combined with the invariance of the parameters of interest to a class of interventions that alter the marginal model, so that the parameters of the conditional model remain constant under a regime change (see Hendry, 1992). Thus, finding SE implies WE, and demonstrating SE relies on showing that the parameters of the conditional model remain constant even though the marginal model changes.

Before testing for SE, note that the coefficients of $\Delta y$ and $(c-y)_{t-2}$ (shown in Figures 4.1 and 4.2 together with ± twice their estimated standard errors) remain constant over time. This latter constancy argues against significant simultaneous-equations bias since the data correlations have not remained constant (see Hendry, 1992). Figure 4.3 shows that the coefficient of $\pi$ is also constant.

Given the constancy observed in the parameters of the conditional model (3.1), testing for SE requires specifying non-constant marginal models for

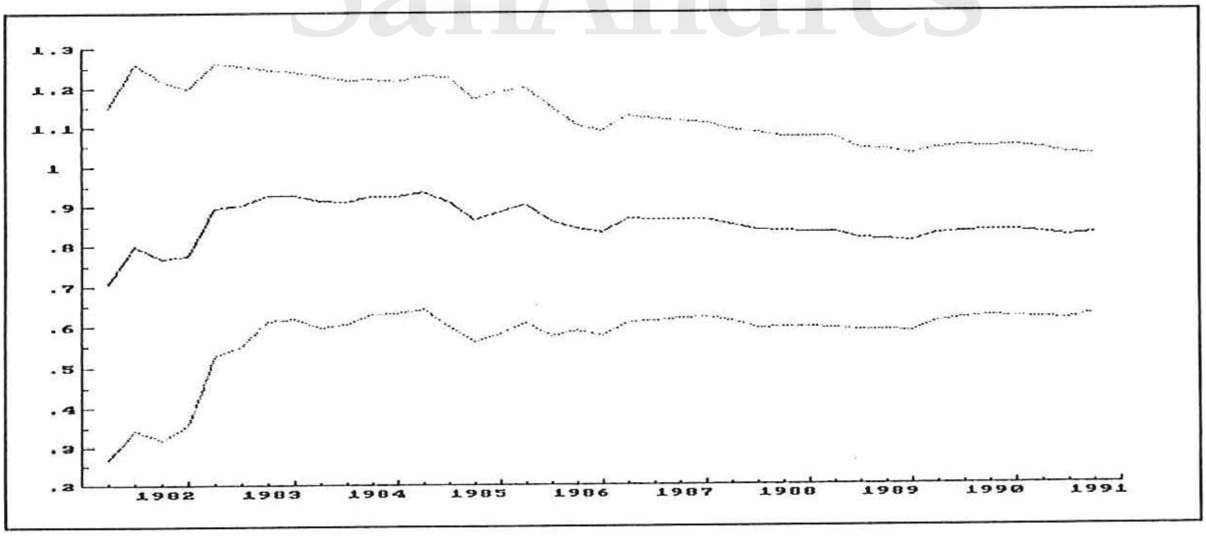


FIG. 4.1. Equation (3.1): Recursive estimation of the coefficient of $\Delta y$.

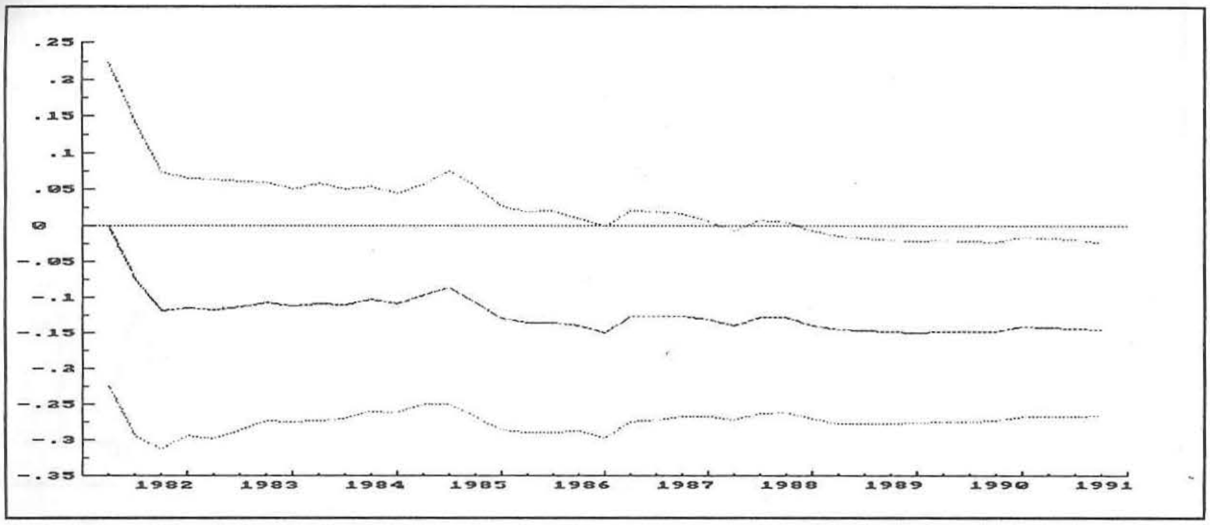


FIG. 4.2. Equation (3.1): Recursive estimation of the coefficient of the error-correction term $(c-y)_{t-2}$.

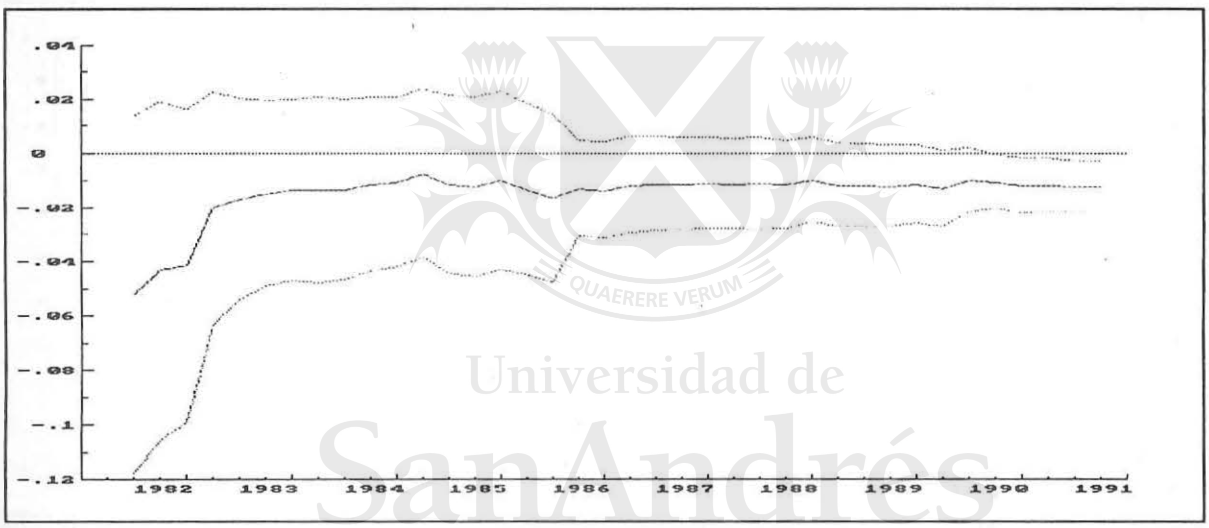


FIG. 4.3. Equation (3.1): Recursive estimation of the coefficient of $\Delta\pi$.

income and inflation.

Estimating over 1977(1)-1990(4), we obtain the following autoregressive model for income:

$$(4.1)\quad \Delta y_t = \underset{(0.001)}{0.031} + \underset{(0.112)}{0.262}\, \Delta y_{t-4} + \underset{(0.070)}{0.296}\, \Delta_4 y_t + \underset{(0.030)}{0.089}\, \mathbf{D85(4)}$$

$$+ \underset{(0.030)}{0.076}\, \mathbf{D89(4)} - \underset{(0.016)}{0.077}\, Q_1 - \underset{(0.012)}{0.016}\, Q_2 - \underset{(0.013)}{0.041}\, Q_3$$

$T=56 \quad K=8 \quad \sigma=2.85\% \quad R^2=0.776 \quad DW=1.68$
$FAR_{1-1}(1,47)=0.75 \quad FAR_{1-4}(4,44)=1.75 \quad Chi^2(2)=0.84$
$FARCH_{1-4}(4,40)=2.09 \quad FHET(9,38)=0.53 \quad FRESET(1,47)=0.42$

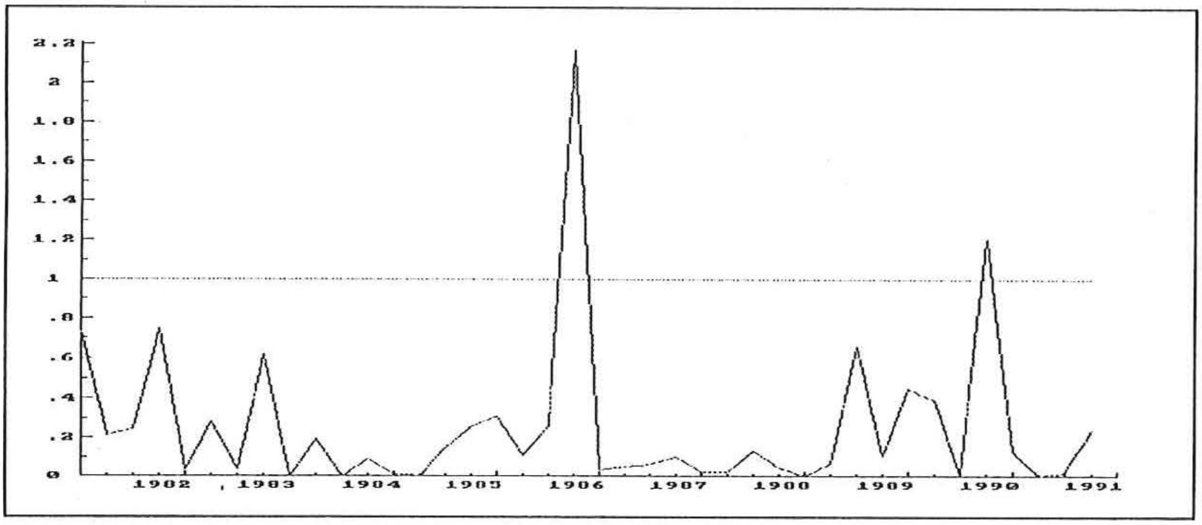

FIG. 4.4. Equation (4.1): Sequence of one-step Chow-statistics when **D85(4)** and **D89(4)** are excluded.

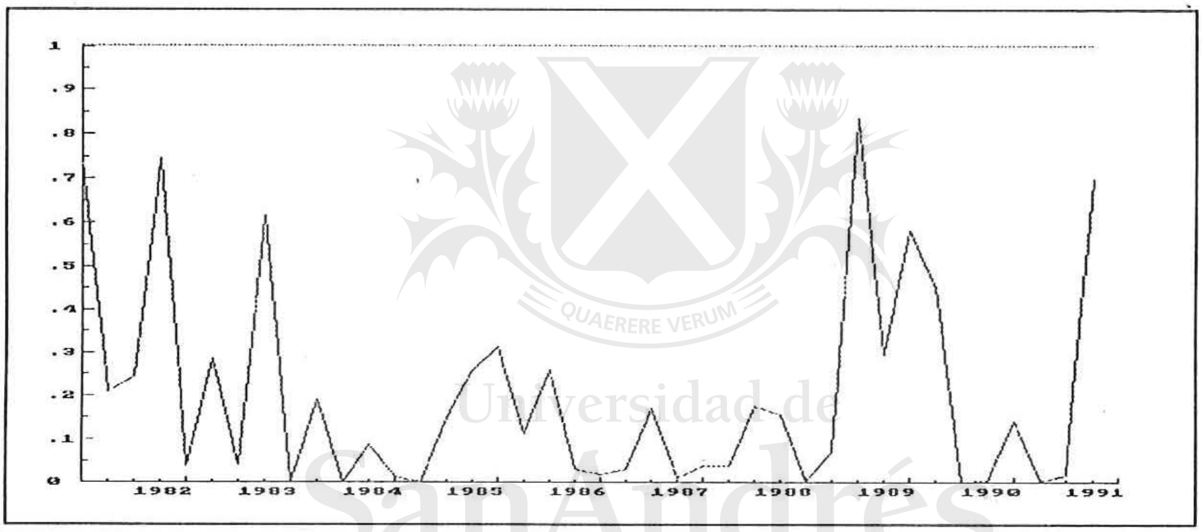

FIG. 4.5. Equation (4.1): Sequence of one-step Chow-statistics when **D85(4)** and **D89(4)** are included.

As can be seen, this is a well specified, stable marginal model for income. Its stability depends on including a set of intervention dummies -in quarters 1985(4) and 1989(4)- which were not necessary for the stability of the model (3.1). Instead, these dummy variables are crucial for achieving stability in (4.1). This is evident from the sequences of Chow-statistics shown in Figures 4.4 and 4.5, where the first excludes the dummies while they are included in the second. They are also joint-significant at the 1% level, as results from performing the corresponding joint F-statistic

(FNIN(2,48)=7.17).

The joint occurrence of structural breaks in the process of the conditioning variable $\Delta y_t$ and constancy of the coefficient of this variable in the consumption function indicate invariance with respect to the class of interventions in the sample period. To prove this, we must add the dummies in equation (4.1) to the model (3.1) and show that they are not significant (see Engle and Hendry, 1993). The result is:

$$
\begin{aligned}
(4.2)\quad \Delta c_t = {} & \underset{(0.010)}{0.003} + \underset{(0.105)}{0.835}\,\Delta y_t - \underset{(0.006)}{0.013}\,\Delta\pi_t - \underset{(0.004)}{0.017}\,\Delta\pi_{t-3} \\
& - \underset{(0.062)}{0.144}\,(c_{t-2} - y_{t-2}) - \underset{(0.029)}{0.006}\,\mathbf{D85(4)} - \underset{(0.030)}{0.007}\,\mathbf{D89(4)} \\
& - \underset{(0.014)}{0.018}\,Q_1 - \underset{(0.009)}{0.105}\,Q_2 - \underset{(0.011)}{0.008}\,Q_3
\end{aligned}
$$

T=56 K=10 $\sigma$=2.33% $R^2$=0.918 DW=2.04
$FAR_{1-1}(1,45)=0.03$ $FAR_{1-4}(4,42)=0.85$ $Chi^2(2)=3.45$
$FARCH_{1-4}(4,38)=0.42$ FHET(13,32)=1.02 FRESET(1,45)=0.96

The dummies are not joint-significant at the 1% level (FNIN(2,46)=0.04). These results suggest that income is super exogenous in the conditional model (3.1). We have also performed a test of weak exogeneity by adding the lagged error-correction term to (4.2) (see Urbain, 1992). The t-value is -1.293, which corroborates the exogeneity of income with respect to the long-run parameters.

As with income, we have estimated non-constant autoregressive marginal models for the rate of inflation. Nevertheless, in the case of $\pi$ we were not able to achieve a model with the goodness of fit and the stability properties of (4.1) by the use of intervention dummies.[10] These facts could lead to the conclusion that the marginal process of $\pi$ presented structural instability during the period.[11]

[10] A better fit to an univariate representation of $\pi$ (including several dummies) was obtained by Ahumada (1992a) using monthly data for the period 1977-1988.

[11] However, it would be desirable in further research to estimate a multivariate conditional model for $\pi$. A way could be a model of CPI inflation which conditions $\pi$ on the rates of change of wages, rate of exchange, prices of the 'flexprice' sector and public utility rates. Applying such a model to monthly data since 1983(2), Damill and Frenkel (1990) detect, according to the cumulative sum (CUSUM) of recursive residuals, the presence of instability in

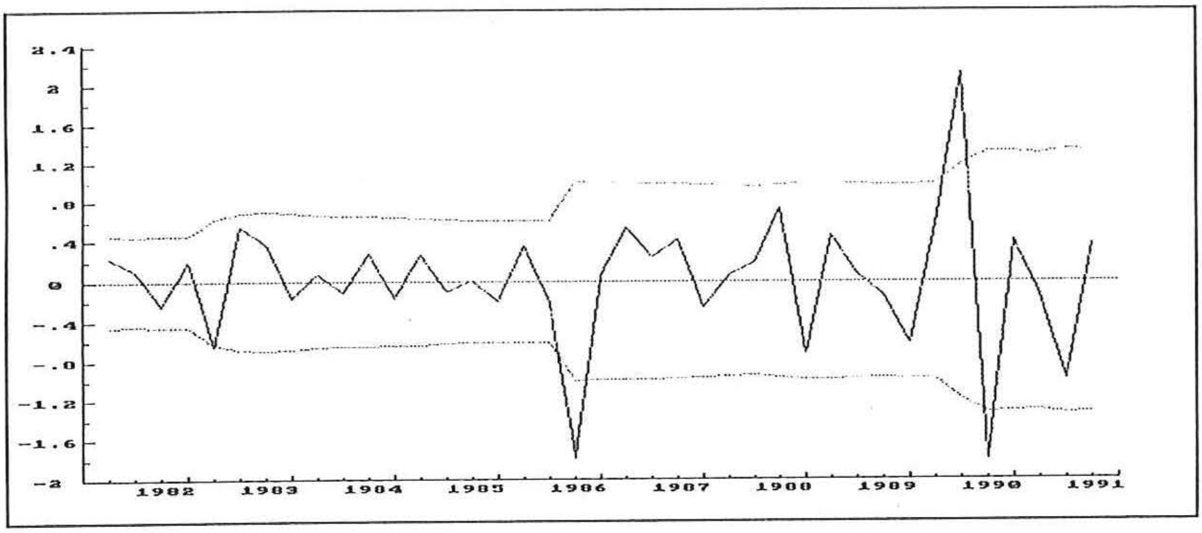


FIG. 4.6. Reversed regression for $\pi$: One-step residuals and the corresponding calculated equation standard errors.

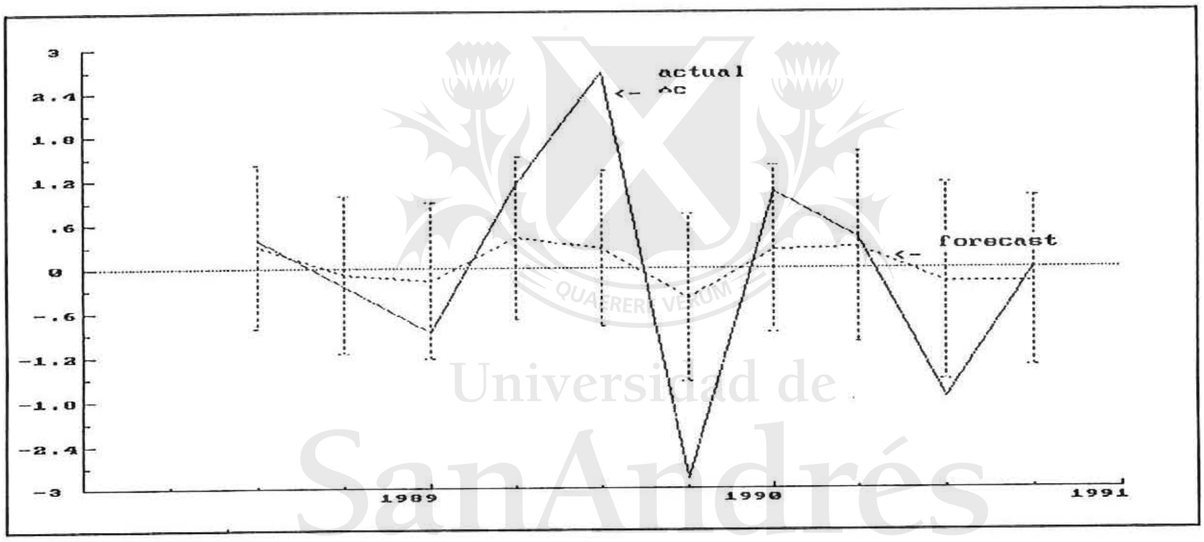


FIG. 4.7. Reversed regression for $\pi$: One-step ahead forecast values with ± 2 forecast standard errors.

A final aspect of exogeneity is 'non-invertibility'. As we have seen, the parameters of interest in model (3.1) were stable. Therefore, the presence of 'regime shifts' which alter the correlation structure of the variables implies that the coefficients of the reversed models cannot be constant. The reversed regression for y showed joint instability at the 5% level, although it did not cross the critical values of the recursive Chow statistics and did

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mid-1989.

not present predictive failure. However, as the equation disturbances are autocorrelated ($FAR_{1-4}(4,44)=3.25$), the estimation results can be seriously affected (see Hendry, 1992).

Finally, let's consider the reversed model for $\pi$. We can reject the presence of autocorrelation ($FAR_{1-4}(4,44)=1.71$). Figure 4.6 shows the 1-step ahead residuals of this regression and Figure 4.7 graphs ahead forecast values of $\Delta\pi$ with ± two forecast standard errors. The instability is evident, which is consistent with finding super exogeneity for $\pi$ in (3.1).

## 5. Conclusions.

Our study on consumers' expenditure finds a relationship that remains stable over major policy changes from 1977(1) to 1990(4). The model focuses on just two variables: income and inflation. Not only does it represent adequately the short-run dynamics of consumers' expenditure, but also its long-run properties are analyzed by cointegration techniques.

This paper discusses exogeneity issues and applies several tests. The conditional model of consumers' expenditure appears constant, so the empirically non-constant univiariate marginal model for income implies the super exogeneity of this variable in the consumption function. An implication of super exogeneity is that the reversed regressions of the conditional model for its explanatory variables may be non-constant. That is the case in Argentina for the rate of inflation.

As we have demonstrated super exogeneity for income and inflation for the classes of interventions occurred during the period, we are able to make some statistically valid policy analysis (see, inter alia, Campos and Ericsson, 1988). In so far as government actions affect income and inflation, economic policy can (and will) influence the behavior of consumers' expenditure. However, more detailed discussion of the policy implications of our model requires specifying the policy at hand and analyzing the concrete mechanisms by which it influences income and inflation.

In conclusion, the validity of our conditional model supports the adequacy of the usual assertions about two prominent features of the Argentine economy: i) the great importance of current income in explaining consumers' expenditure; ii) the transmission mechanism of the volatility of inflation to the volatility of aggregate demand via consumers' expenditure.[12]

[12]The most important transmission channels were described in Section 1. The Appendix incorporates the purchasing-power of wages to our set of explanatory variables.

## Appendix

As we mentioned before, in the period of analysis the acceleration of inflation always entails the deterioration of the purchasing power of wages (PPW). We have estimated a conditional model incorporating this variable to our data set. This variable is defined as the logarithm of the quarterly average of the monthly wages per worker in the industrial sector (provided by the Labor Ministry), deflated by CPI of the following period. The parsimonious model is the following:

$$\text{(A.1)} \quad \Delta c_t = \underset{(0.016)}{-0.014} + \underset{(0.098)}{0.866}\,\Delta y_t - \underset{(0.005)}{0.010}\,\Delta\pi_t + \underset{(0.022)}{0.058}\,\Delta PPW_{t-3}$$
$$+ \underset{(0.023)}{0.102}\,\Delta PPW_{t-4} - \underset{(0.060)}{0.195}\,(c_{t-2} - y_{t-2}) - \underset{(0.014)}{0.008}\,Q_1$$
$$- \underset{(0.009)}{0.096}\,Q_2 - \underset{(0.011)}{0.001}\,Q_3$$

$T=56$ $K=9$ $\sigma=2.21\%$ $R^2=0.924$ $DW=2.00$
$FAR_{1-1}(1,46)=0.00$ $FAR_{1-4}(4,43)=1.73$ $Chi^2(2)=2.08$
$FARCH_{1-4}(4,39)=0.95$ $FHET(13,33)=1.34$ $FRESET(1,46)=0.44$
$VIT=0.32$ $JIT=2.08$ Chow $F(10,37)=0.19$

The performance of this model is very similar to (3.2), as can be seen from the statistics below (A.1) and the Figures A.1 and A.2. The model (A.1) passes successfully all the recursive statistics concerning stability,

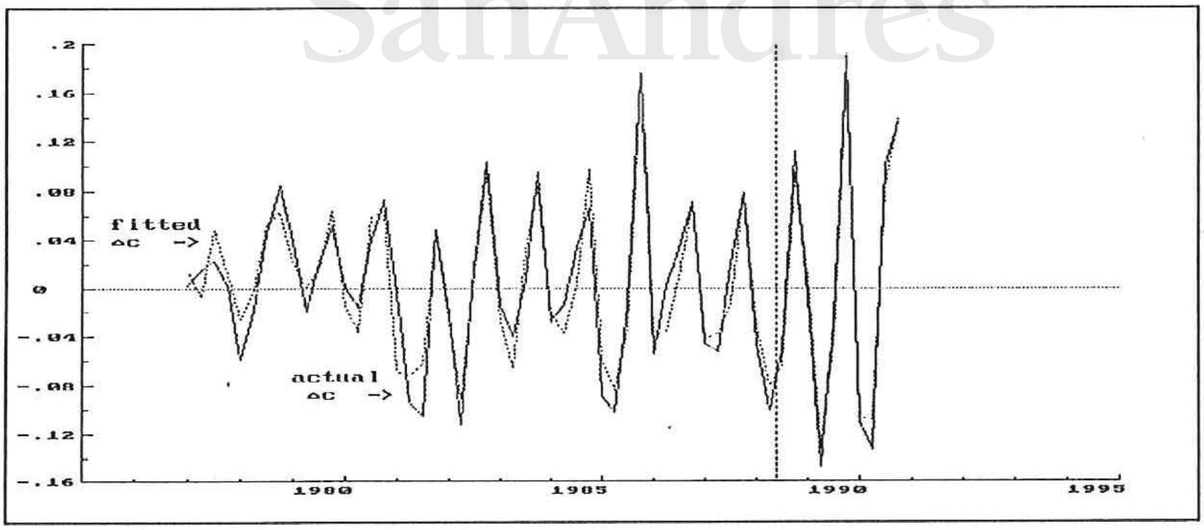


FIG. A.1: Equation (A.1): Actual outcomes and fitted values with a ten-quarter forecast sub-period.

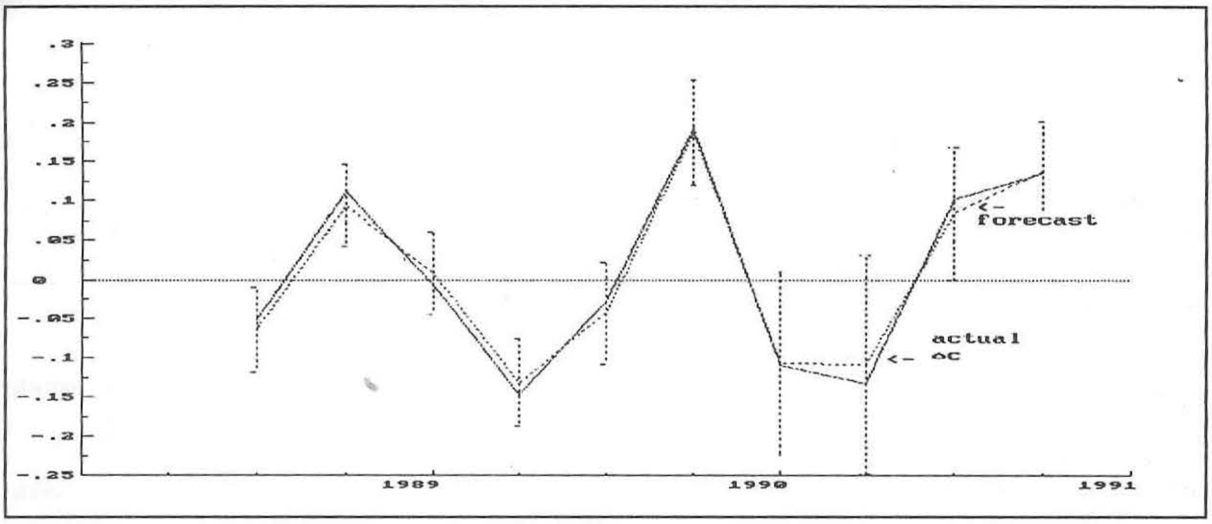


FIG. A.2: Equation (A.1): One-step ahead forecast values of △c with ± 2 forecast standard errors.

although it performs poorer than (3.2) because the coefficient of $\triangle PPW_{t-4}$ shows some increasing tendency. As with (3.1), there is a gain of significance over time for all variables, with $\triangle\pi$ and $\triangle$**PPW** becoming significative toward the end of the period.

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