Intertemporal consumption substitution and inflation stabilization:An empirical investigation

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Comments welcome

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Abstract

Exchange rate based inflation stabilization programs in developing countries often lead to an initial consumption boom followed by an eventual recession. To explain such phenomenon, theoretical models have focused on the role of intertemporal consumption substitution in response to temporary reductions in nominal interest rates. This paper assesses the empirical relevance of such mechanism for six high-inflation developing countries that have gone through repeated stabilization attempts. A simple monetary model is used to obtain estimates of the intertemporal elasticity of substitution, and dynamics simulations are carried out to test the predictive power of the model. The analysis concludes that, in several cases, temporary shocks appeared to have played a key role in generating a consumption boom.

1/ The authors are grateful to Allan Drazen, Graciela Kaminsky, Leonardo Leiderman, Enrique Mendoza, Peter Montiel, Jonathan Ostry, and seminar participants at the International Monetary Fund, the University of Maryland.

I. Introduction

Conventional wisdom holds that inflation can be brought down only at the cost of a recession. A tight monetary policy aimed at stopping inflation is thus viewed as a "powerful anticancer drug that has long-lasting and painful side effects" (Gordon (1982, p.11)). The contractionary effects of disinflationary policies are thought not to depend on whether the money supply or the exchange rate is used as the nominal anchor (Fischer (1986)). Typically, then, the key issues have been (a) the mechanisms through which such policies are contractionary, and (b) the associated "sacrifice-ratio" (i.e., the cumulative percent output loss per percentage point reduction in inflation). The belief that disinflation is always and everywhere contractionary is so ingrained that theoretical results to the contrary have been dismissed as mere intellectual curiosities (see Fischer (1986, p. 252)).

Recent evidence on disinflationary programs in chronic inflation countries, however, has challenged the traditional view. 2/ As Figure 1 illustrates, major exchange rate-based stabilization programs in chronic-inflation countries have usually led to a sharp increase in consumption at the beginning of the programs (see Kiguel and Liviatan (1992) and Végh (1992)). 3/4/ Later in the programs, consumption contracted. Furthermore, the late contraction has also been observed in successful programs, such as the Israeli one. Hence, this boom-recession cycle in exchange rate-based inflation stabilization programs in chronic inflation countries has emerged as a highly intriguing phenomenon, whose understanding is of great theoretical and practical relevance. Table 1, which documents these sizable fluctuations in real consumer spending during several stabilization episodes, shows initial increases in consumption ranging from 12 to 33 p

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2/ Naturally, the most celebrated challenge to conventional wisdom is Sargent's (1982) interpretation of the end of four hyperinflations. In contrast, this paper focuses on chronic inflation countries; that is, countries which have had long periods (even several decades) of high (relative to OECD countries) but relatively stable inflation.

3/ These programs comprise the Southern-Cone orthodox stabilization plans implemented during 1978 in Argentina, Chile, and Uruguay and the heterodox plans in Argentina (the 1985 Austral plan), Israel (1985), Brazil (the 1986 Cruzado plan), and Mexico (1987). The same pattern occurred in programs that took place in Latin America in the 1960's.

4/ It is worth noting that in OECD countries, consumption booms have also been observed in the first stages of the 1982 Danish and 1987 Irish exchange rate-based stabilization (see Giavazzi and Pagano (1990)).

ercent (the average duration of these programs is about seven quarters) followed (in six of the seven programs) by declines of comparable orders of magnitude.

Three main explanations have been advanced to explain the boom-recession cycle associated with exchange rate-based stabilization. 5/ An early paper by Rodriguez (1982), inspired by the Argentine 1978 program, focused on the expansionary effects of lower real interest rates. Specifically, in the presence of perfect capital mobility, a reduction in the rate of devaluation lowers the nominal interest rate. Since inflationary expectations do not change in the short-run (due to the assumption of adaptive expectations), the real interest rate falls thus stimulating aggregate demand. Given that inflation remains high, the real exchange rate appreciates, eventually throwing the economy into a recession. There are two main drawbacks to Rodriguez's (1982) hypothesis. 6/ At an empirical level, it cannot explain the consumption boom in those plans in which real interest rates rose. 7/ At a theoretical level, his results need not hold in a utility-maximizing framework, as shown by Calvo and Végh (1993).

A second explanation relies on wealth effects that may accompany the implementation of a stabilization plan. In Helpman and Razin (1987), the reduction in the proceeds from the inflation tax leads to a wealth effect due to the absence of Ricardian equivalence. In Drazen and Helpman (1988), the wealth effect results from the expectation that the fiscal deficit will be closed by a future reduction in government spending. For the case of Israel, Bruno (1993) argues that confidence in the program raised the value of government bonds. A problem with wealth effects, whatever their origin, is that they cannot explain the late recession.

5/ This brief review is not intended to be exhaustive; see also Dornbusch (1982), Obstfeld (1985), Drazen (1990), De Gregorio, Guidotti, and Végh (1992), and Roldos (1993),

6/ The assumption of adaptive expectations, a common criticism of Rodriguez's (1982) explanation, is actually not essential. The same results follow in a rational expectations framework, as long as the inflation rate is assumed predetermined because of, say, nominal rigidities (see Dornbusch (1982) and Calvo and Végh (1993)).

7/ Real interest rates rose in four of the seven episodes illustrated in Figure 1: the Austral, Cruzado, Israeli, and Mexico plans (see Végh (1992)).

A third explanation (the so-called "temporariness" hypothesis) is based on the existence of lack of credibility, modeled as temporary policy (see Calvo (1986) and Calvo and Végh (1993)). In the context of a cash-in-advance model, Calvo (1986) shows that a temporary reduction in the rate of devaluation leads to a consumption boom. In the presence of perfect capital mobility, the reduction in the devaluation rate implies a fall in the nominal interest rate. Since money must be used to purchase goods, the opportunity cost of holding money (given by the nominal interest rate) is part of the effective price of consumption. Hence, a temporary reduction in the nominal interest rate makes present consumption less expensive relative to future consumption, which induces consumers to substitute present for future consumption. 8/ When the interest rate falls back to its initial level, consumption falls. Temporary policy can be reinterpreted as non-credible policy. Specifically, if policymakers announce a permanent reduction in the devaluation rate but the public believes that the policy will be discontinued in the future, the same results obtain.

Calvo and Végh (1993) introduce a non-traded good and sticky prices into the picture and show that a non-credible stabilization generates an initial boom in both the traded and non-traded goods sectors. Inflation remains high due to the anticipation of a policy reversal, which leads to a sustained real exchange rate appreciation. The cumulative effects of the real exchange rate appreciation may throw the economy into a recession even before the end of the program. Hence, the temporariness hypothesis generates predictions that match most of the stylized facts associated with exchange rate-based stabilizations in chronic-inflation countries (see Végh, 1992).

At a theoretical level, therefore, the "temporariness" hypothesis offers intuitive and sharp predictions. Its empirical relevance, however, has received little or no attention. Since intert

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8/ The same results arise if money is introduced in the utility function and the cross-derivative between consumption and real money balances is assumed to be positive (Calvo (1986)).

emporal elasticities of substitution are believed to be low, many actually doubt that the "temporariness" hypothesis may have any empirical relevance at all. This paper attempts to fill that gap by examining the quantitative importance of the temporariness hypothesis in explaining the consumption cycle in exchange rate-based stabilization programs.

As suggested above, non-credible stabilization under predetermined exchange rates is supposed to affect consumption through temporary changes in nominal interest rates. Hence, an empirical assessment of such a transmission channel should proceed in two stages. In the first stage, the *existence* of this channel must be evaluated. For this channel of transmission to exist, (a) shocks to nominal interest rates should have a temporary component, and (b) there must be evidence of consumption smoothing; that is, the intertemporal elasticity of consumption substitution must be (statistically) different from zero. Clearly, if either condition were not satisfied, then temporary policy would be devoid of empirical content. These two conditions, however, are necessary but not sufficient for temporary policy to be quantitatively important. The second stage, then, is to evaluate the *quantitative importance* of the temporariness hypothesis by examining whether observed movements in nominal interest rates are large enough to generate, given the estimated elasticities of substitution, changes in consumption that roughly match the observed ones.

Six chronic-inflation countries--Argentina, Brazil, Chile, Israel, Mexico, and Uruguay--have been chosen for the empirical study undertaken in this paper. These chronic-inflation countries provide a highly suitable laboratory for an empirical examination of the temporariness hypothesis for two main reasons. First, such countries have been characterized by highly variable inflation and thus nominal interest rates (relative to industrial countries) stemming from "stop and go" policies. Indeed, as Table 2 shows, the variability of nominal interest rates in these countries is much higher than that in the United States. In Argentina, for instance, the variability of the quarterly nominal interest rate, as measured by the standard deviation is about 30 times that of the United States (see Table 2). Not surprisingly, quarterly changes in real per capita consumption are also anywhere between twenty-to-fifty times more variable in the sample countries than in the United States (see Table 2). Second, in these six countries, there have been seven major exchange rate-stabilization in which the consumption boom-recession cycle has been observed (see Figure 1). 9/ Since most of these programs were temporary, they provide clear-cut episodes in which the effects of temporary changes in nominal interest rates on consumption can be evaluated.

The paper first addresses the issue of the relative importance of temporary and permanent shocks to nominal interest rates. To that effect, we use Cochrane's (1988) methodology for decomposing the variance of nominal interest rates into its permanent and transitory components. This decomposition indicates that temporary shocks account for an important share (between 40 and 90 percent) of the total variance of nominal interest rates. In effect, in four out of the six countries, temporary shocks account for more than 75 percent of the variance of nominal interest rates. We thus conclude that temporary shocks, some of which surely stem from a history of stop and go policies, have been an important element in explaining the behavior of nominal interest rates. Not surprisingly, other variables affected by stabilization policies, such as the real exchange rate, have also been shown to be subject to substantial temporary shocks (see Calvo, Reinhart, and Végh (1994)).

The paper then addresses the existence of the transmission channel discussed above by estimating the intertemporal elasticity of consumption substitution for each of the six countries. Formally, we estimate the Euler equation that follows from a transactions costs model using Hansen's (1982) generalized method of moments (GMM). Our estimates suggest that the interte

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9/ With the exception of Mexico, where the late contraction in consumption has not occurred.

mporal elasticity of substitution is rather low but statistically different from zero. Specifically, the estimates for the six countries generally lie in a 0.05-0.8 range, and vary within countries depending on whether deposit rates or stock market returns are used. However, most often estimates tend to be low and clustered around 0.20.

Having established the existence of the necessary ingredients for the temporariness hypothesis to be empirically relevant, we proceed to assess its quantitative importance (i.e., its power in explaining the consumption pattern in actual episodes). Using the estimates of the elasticity of substitution and the Euler equations, we undertake dynamic simulations for the seven major stabilization programs mentioned above to see how well the predicted path of consumption replicates the actual path after the plans were implemented. We conclude that the model performs poorly for the Argentine and Uruguayan tablitas, as it misses the initial consumption boom entirely. The model performs better for the other five programs, especially for Brazil, Mexico, and Israel where the predicted cumulative changes in consumption roughly match the observed ones.

In conclusion, the empirical importance of the temporariness hypothesis in explaining the consumption pattern observed in exchange rate-based stabilizations in chronic inflation countries gets mixed results. The necessary conditions exist: the intertemporal elasticities are (statiscally) different from zero and temporary shocks are important. Hence, the temporariness hypothesis cannot be dismissed on these grounds. Given the low values of the elasticities of substitution, however, substantial changes in nominal interest rates are needed to generate observed changes in consumption.

The paper proceeds as follows. Section II derives the Euler equation. The relative importance of temporary shocks is assessed in Section III, which also presents the estimation results. Section IV contains the dynamic simulation of the model. Section V concludes.

II. The Model

As discussed in the Introduction, models that incorporate the "temporariness" hypothesis rely on the effect of temporary changes in nominal interest rates on consumption that take plac e through the cash-in-advance constraint. The cash-in-advance hypothesis, however, which im plies a fixed velocity, is too restrictive an assumption for empirical purposes. Therefore, mone y will be introduced into the model through a transactions costs technology (see, for instance, Kimbrough (1990) and Reinhart (1990)).

Consider a small open economy inhabited by a large number of identical, infinitely-lived individuals. The representative consumer maximizes

$$E_0 \sum_{t=0}^{\infty} \beta^t u(c_t), \qquad \textbf{(1)}$$

where E_0 is the expectations operator conditional on information available at time 0; u(.), the in stantaneous utility function, is increasing, concave, and twice-continuously differentiable; $\beta \in ($ $0,1)$ is the discount factor; and c denotes consumption of the only (non-storable) good.

There exist two assets in this economy: money (M) and an internationally-traded bond (B) (whose rate of return is i). The consumer must incur in transactions costs to purchase goods . Specifically, transactions costs (in terms of the consumption good) are given by

$$s_t = v(c_t, m_t), \quad v_c > 0, v_m < 0, v_{cc} > 0, v_{mm} > 0, v_{cm} < 0, v_{mm} v_{cc} - v_{cm}^2 \geq 0, \qquad \textbf{(2)}$$

where $m(\equiv M/P)$ stands for real money balances (P is the nominal price of the good). Thus, tra nsactions costs are increasing in consumption and decreasing in real money balances.

At the beginning of period t, the consumer holds M_{t-1} and B_{t-1}, which he carried over fro m period t-1, and receives $i_{t-1}B_{t-1}$ as interest income. The nominal interest rate is given by $i=i^*$ $+\varepsilon$, where i^* and ε are the world interest rate and the rate of devaluation, respectively. The con sumer then chooses consumption and end-of-period money holdings (M_t) and bonds (B_t). Any consumption plan must satisfy the following sequence of budget constraints:

$$b_t = (1+r_{t-1})\, b_{t-1} + y_t - c_t - v(m_t, c_t) + \frac{m_{t-1}}{1+\pi_{t-1}} - m_t, \tag{3}$$

where b_t denotes the end-of-period real stock of internationally-traded bonds, y denotes the endowment of the tradable good, and $r_{t-1} \equiv (1+i_{t-1})P_{t-1}/P_t$ is the real interest rate.

The first-order conditions are, in addition to (2), 10/

$$u'(c_t) = \beta E_t[u'(c_{t+1})(1+r_t)\frac{p_t}{p_{t+1}}], \tag{4}$$

$$-v_m(m_t, c_t) = \frac{i_t}{1+i_t}, \tag{5}$$

where

$$p_t = 1 + v_c(m_t, c_t), \tag{6}$$

denotes the effective price of consumption. The effective price of consumption consists of the market price of the good (equal to unity) plus the transactions costs that result from purchasing an additional unit of the good, v_c.

Equation (4) is a stochastic Euler equation that describes the consumer's optimal consumption plan given the path of the real interest rate and the effective price. For our purposes, it is important to emphasize the role played by the effective price of consumption. Consider, for the sake of argument, a world of perfect foresight in which $\beta=1/(1+r)$. Then, if the time profile of the effective price of consumption is flat over time, it follows from (4) that consumption is also constant over time. Suppose now that the effective price of consumption falls to a lower level at time 0 but is expected to go back to its initial level at time T. Then, consumption increases at time 0, remains at that level until T-1, and falls at time T. 11/ The reason is that consumpti

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10/ We assume that, except for r_t (which involves next period's price level), the consumer knows at time t all variables dated with the subscript t.

11/ We are assuming that there are no wealth effects associated with the change in the time path of the effective price of consumption.

on is cheaper in the present (i.e, between time 0 and T) relative to the future (i.e., after time T). Thus, since the effective price depends negatively on the nominal interest rate (as shown belo w), equation (4) shows the link between <u>temporarily</u> lower nominal interest rates and higher co nsumption.

Equation (5) implicitly defines the demand for real money balances. Real money deman d depends positively on consumption and negatively on i/(1+i) (recall that $v_m<0$ and $v_{cm}<0$).

In order to estimate the model, we assume that the utility function and the transactions c osts technology take the following form:

$$u(c)=\frac{c^{1-1/\rho}}{1-1/\rho}, \qquad v(m,c)=k_0 c^{k_1} m^{-k_2}, \tag{8, 9}$$

where all parameters, ρ, k0, k1, and k2, are positive numbers. 12/

Under this specification, equations (4) and (5) become

$$E_t\left[(1+r_t)\left(\frac{c_t}{c_{t+1}}\right)^{1/\rho}\frac{p_t}{p_{t+1}}\right]=\frac{1}{\beta}, \qquad k_0 k_2 c_t^{k_1} m_t^{-(1+k_2)}=\frac{i_t}{1+i_t}, \tag{9, 10}$$

where

$$p_t \equiv 1+\frac{k_1}{k_2}\frac{i_t}{1+i_t}\frac{m_t}{c_t}. \tag{11}$$

Equations (9) and (10) will be the focus of our econometric analysis.

Equation (11) indicates that the effective price of consumption depends positively on th e nominal interest rate and negatively on velocity. Conceptually, cash-in-advance models can be viewed as a particular case of the model in which velocity is fixed.

Finally, note that equation (10) gives rise to the following money demand function:

$$\log(m_t)=\frac{1}{1+k_2}\log(k_0 k_2)+\frac{k_1}{1+k_2}\log(c_t)-\frac{1}{1+k_2}\log(\frac{i_t}{1+i_t}). \tag{12}$$

12/ The Cobb-Douglas specification for the transactions technology has been used by Dowd (1990).

We now have all the necessary ingredients to proceed with the empirical implementation of the model.

III. Empirical Evidence

The transmission mechanism outlined in the previous section, which links changes in the nominal interest rate to changes in consumption, rests on two building blocks. The first is non credible monetary policy, in the sense that changes in nominal interest rates associated with the disinflation program are perceived to be largely temporary and reversible. The second key in gredient is the presence of intertemporal substitution in consumption. The empirical relevance of these two key assumptions are, in turn, examined in the remainder of this section.

Data Preliminaries

The sample period was determined by the availability of the pertinent quarterly data. Th e definitions of the variables, the relevant sources, the period of coverage and the first and seco nd moments of the time series of interest are summarized in Table 2. Since in almost all the co untries in our sample disaggregated consumption data on durables, nondurables, and services a re not available, our consumption aggregate is total private consumption. The theoretical fram ework links consumption to a monetary aggregate that is primarily held for transactions purpos es, hence we employ narrow money, M1, in the estimation that follows. 13/ The nominal inter est rates used are on short-term deposits, with the exceptions of Israel and Uruguay where a sh ort-term lending rate is used. Quarterly stock market total return data were also used when ava ilable; the source for this data is International Finance Corporation.

13/ This restrictive definition of transactions balances is not always appropriate. In the case of Brazil, some c omponents of quasi-money are highly liquid (with overnight maturities) and hence, at least in principle, could be used for transactions purposes. In the case of Argentina and Uruguay, transactions are often carried out in U.S. d ollars. There are, however, no available time series on the amount of U.S. currency in circulation.

Table 2 highlights a number of relevant empirical regularities. First, in all the countries considered the variability in nominal interest rates is considerable in both absolute terms and re lative to industrial countries (the United States is included in Table 2 for comparison purposes) . Similarly, consumption appears to be much more variable than in most industrial countries, with a standard deviation 20-50 times larger than the comparable measure for the United States . Such features make these countries an ideal ground for testing the empirical relevance of the "temporariness" hypothesis, and more generally, the presence or absence of consumption smoo thing. Second, the first and second moments of Δx_t make plain that the constant velocity assu mption that characterizes cash-in-advance models is particularly inappropriate for these high-i nflation countries. With the exception of Chile, Table 2 shows that velocity is increasing rapid ly ($\Delta x_t < 0$) in the sample countries and that its variability is sizable. Hence, the transactions c ost model presented in the previous sections is likely to be more consistent with the data.

<u>Temporary and Random Walk Components of Nominal Interest Rates</u>

Lack of credibility refers to agents' belief that the stabilization plan will eventually fail. Alternatively, it can be said that the program enjoys <u>short-run</u> credibility (since a decline in no minal interest rates reflects, in part, a decline in inflation expectations) but not <u>long-run</u> credib ility. Effectively, it implies that the decline in nominal interest rates that accompanies the slow er rate of devaluation is a <u>temporary</u> phenomenon. Hence, gauging the relative importance of t emporary shocks to nominal interest rates is a useful stylized fact to assess the empirical releva nce of the "temporariness" hypothesis. To quantify this issue, we employ Cochrane's (1988) m ethodology, which provides a measure of the persistence of shocks in a variable by examining t he variance of its long differences. Suppose that the variable x has the following representation:

$$x_t = \delta x_{t-1} + \varepsilon_t, \quad \text{where } \varepsilon_t \sim N(0, \sigma^2_\varepsilon). \qquad \textbf{(13)}$$

Then, for $\delta=1$, x follows a pure random walk and the variance of its k-differences grows linearly with the difference

$$\mathrm{var}(x_t - x_{t-k}) = k\,\sigma^2_\varepsilon. \tag{14}$$

If $\delta<1$ and x is a stationary process, the variance of its k-differences is given by

$$\mathrm{var}(x_t - x_{t-k}) = \sigma^2_\varepsilon (1-\delta^{2k})/(1-\delta^2). \tag{15}$$

Therefore, the ratio $(1/k)\mathrm{var}(x_t-x_{t-k})/\mathrm{var}(x_t-x_{t-1})$ is equal to one if x follows a random walk process and converges to zero if x is stationary. If x has both permanent (random walk) and temporary (stationary) components, the ratio will converge to the ratio of the variance of the permanent shock to the total variance of x. Thus, the closer that ratio is to unity, the lower is the relative importance of temporary shocks.

Table 3 summarizes the main results. The values of k range between one and nine years. In all the countries considered temporary shocks account for a sizable share (between 40 and 90 percent) of the variance of nominal interest rates. In effect, in the two countries where nominal interest rates and inflation have been most volatile, Argentina and Brazil, the random walk component of nominal interest rates accounts for 25 percent or less of the variance of changes in nominal interest rates and the degree of persistence in the cyclical (or temporary component) is relatively low. For Israel and Mexico, it is also apparent that the transitory component accounts for most (between 80 and 90 percent) of the variability in nominal interest rates. However, unlike Argentina, Brazil, and Chile, where temporary shocks die out relatively quickly (within five years), the extent of persistence in the cyclical component in Israel and Mexico is considerably greater. Temporary shocks play a lesser role, accounting for between 50 and 60 percent of the variance of nominal interest, in Chile and Uruguay. Other variables affected by stabilization policies, such as the real exchange rate, have also been shown to be subject to substantial temporary shocks. 14/ In line with the results summarized in Table 3, Calvo, Reinhart, and Végh (1994) show that temporary shocks account for about 50-65 percent of the variance of changes in the real exchange rate.

The finding that temporary shocks appear to have a predominant role in explaining the behavior of nominal interest rates in the sample countries is, perhaps, not very surprising given the history of intermittent high inflation episodes and repeated failed stabilization plans that have plagued most of these countries. Further, the empirical literature on credibility has shown that several of the programs under consideration, even some of the successful ones, enjoyed less-than-full credibility. For example, Baxter (1985) constructs a measure of credibility for the Chilean and Argentine Tablita episodes and Rojas (1991) for the Mexican plan. Both of these studies suggest that at different points in time during the course of the program agents attached a nonzero probability to the program's failure. This would, of course, imply that all or part of the initial nominal interest rate decline would be reversed in the future.

Evidence on consumption smoothing

The extent to which consumption responds to a temporary change in the nominal interest rate depends crucially on the elasticity of substitution between current and future consumption, ρ. Hence, if current and future consumption are close substitutes (ρ is high), then a given increase/decrease in the nominal interest rate will induce consumers to postpone/bring forward consumption by more than they would if the degree of intertemporal substitution were low.

The two first order conditions given by equations (9) and (12) include all the parameters that characterize both the transactions technology and consumer preferences. Ideally, to obtain estimates of all the relevant parameters, one would estimate these two equations simultaneousl

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14/ For evidence on the behavior of the real exchange rate during exchange-rate based stabilization programs, see Végh (1992), who documents a pattern of an initial appreciation of the real exchange rate, followed by a complete or partial reversal.

y using Hansen's (1982) generalized method of moments (GMM). However, the time series properties of the data in question complicate the estimation strategy. The variables appearing in the Euler equation (9) are stationary. 15/ However, the money demand equation (12) defines a relationship among I(1) variables in levels. As the vast literature on unit roots and cointegration establishes (see for instance, Engle and Granger (1987)), standard estimation techniques only yield consistent parameter estimates when these variables are cointegrated. If real balances, consumption, and nominal interest rates are cointegrated, the simultaneous estimation of the two first-order conditions (equations (9) and (12)) remains a feasible option. If these variables are not cointegrated, then such an estimation strategy would provide neither consistent estimates of the parameters of the transactions technology nor efficient estimates of the preference parameters (i.e. the intertemporal elasticity of substitution, ρ, and the subjective discount factor, β). The simultaneous estimation, along with the cross-equation restrictions that theory imposes, would contaminate the estimates of the key parameter of interest, ρ.

The foregoing discussion suggests that the next step is to test for cointegration among real balances, consumption and nominal interest rates. Johansen's (1988) cointegration tests were applied to the data and the results are summarized in Table 4. The top of Table 4 reports λ-max and trace statistics while the critical values at five and ten percent levels of confidence are reported below. In four of the six sample countries, we were not able to reject the null hypothesis of no cointegration. The exceptions were Argentina and Israel. The absence of cointegration in fairly standard money demand equations for some of the high-inflation countries included in our sample has been documented in Arrau, De Gregorio, Reinhart and Wickham (1990), who attribute the lack of cointegration (for Brazil, Chile, Israel, and Mexico) to an omitted variable that captures financial innovation and/or currency substitution. More recently, Easterly, Mauro, and Schmidt-Hebbel (1992) find that if dummies reflecting regime changes are introduced

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15/ The results of the unit root tests are not reported in the paper, but are available from the authors upon request.

, the null hypothesis of no cointegration can be rejected for Brazil, and Uruguay but not for Argentina, Chile, Israel, and Mexico. On balance, the mixed results of the cointegration tests suggest that an alternative estimation strategy will be required.

While the empirical relevance of the temporariness hypothesis does not depend importantly on the parameters that characterize the transactions technology, the role played by the intertemporal elasticity of substitution is crucial. 16/ Our approach calls for estimating the parameters ρ and β by fitting the first-order condition defined in equation (9) to the data using Hansen's (1982) generalized method of moments. The transactions parameters, k_1/k_2, were fixed at a level consistent with the theoretical priors ($k_1/k_2 > 0$) and estimates from the available empirical literature. However, to assess the sensitivity of the estimates of ρ and β to the assumed transactions technology, k_1/k_2 was allowed to vary across the fairly broad range of 0.25 to 8.00.

The residuals in the estimated equation are partly forecast errors which, by the assumption of rational expectations, are orthogonal to any variable in the agents' information set at time t. Since all available information is used to forecast future consumption and prices, the number of instruments in agents' information set exceeds the number of parameters to be estimated. The excess of instruments over parameters yields a set of overidentifying restrictions that can be used to test the underlying consumption model. Our estimation strategy also takes into account that the residuals, u_t, may be serially correlated for a variety of reasons. Hall (1988) makes plain that time aggregation problems in the consumption data introduce a first-order moving average process with a known parameter in the error term. 17/ Further, Hayashi and Sims (1983) argue that current values of consumption, c_t, may not be observed before expectations of future consumption, c_{t+1} are formed. This may have the effect of making today's forecast error correlated with last period's yet unobserved error. For these reasons we correct for a first-order movi

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16/ All that is required is that $k_1/k_2 > 0$, a condition which is easily satisfied when the demand for real money balances depends positively on consumption and negatively on the nominal interest rate.

17/ See also Working (1960) for an illustration of how the moving average parameter is calculated.

ng average process in the error term. In addition, given that our samples encompass periods of very high inflation as well as periods of moderate-to-low inflation, we allow for the presence of more general forms of heteroskedasticity in the disturbances.

As Hall (1988) indicates, the presence of a first-order moving average in the error term disqualifies variables measured at time t from being used as instruments to forecast c_{t+1}, i_{t+1}, and x_{t+1}. The most recent permissible instrument is one lagged two periods. The instruments employed are stationary variables and the vector of instrumental variables is given by $z'_t = [\text{constant}, \Delta c_{t-1}, \Delta c_{t-2}, \Delta x_{t-1}, \Delta x_{t-2}, \Delta i_{t-1}, \Delta i_{t-2}, r_{t-1}, r_{t-2}]$. There are nine instruments and, therefore, nine orthogonality conditions. With two parameters being estimated, there are seven over identifying restrictions.

Table 5 presents the results when the real interest rate measure used in estimation was based on deposit rates while appendix Table A.1 presents the results for a subset of the countries using stock market returns. Our main findings can be summarized as follows. First, the estimates of the parameters are economically meaningful as $\rho > 0$ and $\beta > 0$, and the hypothesis that $\beta < 1$ (say 0.99 or 0.98) cannot be rejected. Second, there is evidence of consumption smoothing. With the exception of Uruguay for values of $k_1/k_2 \leq 2$ the intertemporal elasticity of substitution is significantly different from zero when deposit rates are used. 18/ Third, while statistically significant, the degree of intertemporal substitution is low; our estimates of ρ lie in the 0.05-0.80 range with most estimates clustered around 0.20. Only in the case of Chile, when stock returns are used and $k_1/k_2 \leq 2$, does ρ exceed unity. Fourth, for most of the countries in our sample (with the exception of Chile) the estimates of ρ and β are not very sensitive to changes in the transactions costs parameters, k_1/k_2, irrespective of which measure of real returns is used. Las

18/ When stock returns are used, the estimates of the parameters are similar, as $\rho > 0$ and $0 < \beta < 1$. However, the point estimates of ρ for Argentina and Mexico are estimated less precisely.

tly, the J-statistics are small relative to the degrees of freedom, indicating that the overidentifying restrictions are not rejected by the data. 19/

Comparison with existing estimates

We next compare our results to those studies that have either introduced money into the consumer choice problem, or have applied the Euler equation approach to developing country data. The earlier work of Giovannini (1985) and Rossi (1988) employed Hall's linearized Euler equation which results from a one-good non-monetary model. 20/ Giovannini's estimates for Argentina, Brazil and Mexico paralleled those of Hall (1988) for the United States, generally finding that ρ was not significantly different from zero. 21/ However, this framework embodies a number of restrictive assumptions: first, it does not allow for monetary considerations in consumption decisions; second, it limits intratemporal substitution across commodities. There is some evidence that relaxing either or both of these assumptions enhances a model's ability to fit the data. Eckstein and Leiderman (1992) estimate a Sidrauski type model for Israel and find that their model is not rejected by the data; their estimates of the intertemporal elasticity of substitution are comparable to ours and are clustered in the 0.15-0.50 range, depending on the instrument set used. Poterba and Rotemberg (1987) fit a similar model to United States data; they estimate ρ at 0.15. Using such a model, Arrau (1990) finds much higher intertemporal elasticities of substitution for Chile and Mexico, 1.50 and 3.00, respectively, although these are estimated with little precision.

A different line of research has applied the Euler equation approach to developing country data, but rather than introducing money into the household optimization problem, this work h

19/ The only exception is Argentina, when stock returns are used and $k_1/k_2 \leq 2$.

20/ The Hall specification of the Euler equation would also apply in a multiple commodity setting if the consumption and price indices were good proxies for the true utility-based indices. However, existing indices do not employ such a methodology.

21/ Rossi (1988) maintains the Hall framework but allows for liquidity constraints. He finds, for a panel of nine Latin American countries, that ρ is about 0.10 but statistically significant.

as emphasized the role of disaggregation across commodities. Ostry and Reinhart (1992), who disaggregate among traded and non-traded goods for a panel of four Latin American countries (including Brazil and Mexico), estimate the intertemporal elasticity of substitution to lie in the 0.37-0.43 range. 22/ Similarly, disaggregating among food and non-food components, Atkenson and Ogaki (1992) estimate ρ to be about 0.27 for India and 0.40 for the United States.

In sum, the previous discussion appears to suggest that most of the evidence from developing countries points to a rather low degree of substitution between current and future consumption, as most estimates of ρ are less than 0.80. However, several of these studies, including the present one, conclude that the elasticity of substitution is significantly different from zero, which implies that the empirical relevance of the "temporariness" cannot be dismissed on those grounds alone.

IV. Dynamic simulations under perfect foresight

As shown in the previous section, a specification such as equation (9) yields parameter values that make economic sense, overidentifying restrictions that are not rejected by the data and, in the majority of instances, estimates that are robust to variations in transactions technology. However, these criteria say little about whether the simple model presented in Section II is capable of replicating the observed consumption pattern. To address this question, we now turn to dynamic simulations. 23/

In the following exercise we are particularly interested in determining whether the model captures the consumption booms that have characterized the early phases of the stabilization programs under study. The assumption of perfect foresight implies that actual values for real an

22/ Their estimates are significant at standard levels of confidence.

23/ An alternative strategy, pursued in Reinhart and Végh (1994), is to assume a constant path (except for discrete jumps) of the nominal interest rate and obtain a closed-form solution for the consumption path. This closed-form solution can then be used to obtain numerical estimates for the changes in consumption that might be observed in practice. In contrast, the dynamic simulations presented here take into account the actual paths of both the nominal and the real interest rate to compute the estimated consumption path throughout the stabilization plan.

d nominal interest rates and velocity were used to generate the forecasts. Rearranging terms in equation (9) it is easy to see that

$$c_{t+1} = c_t[\beta(1+r_t)\frac{P_t}{P_{t+1}}]^{\rho}. \qquad \mathbf{(16)}$$

Since the forecast is dynamic, predicted consumption at t+1 will depend on predicted consumption at t. Actual and predicted levels are set equal to unity in the quarter immediately preceding the stabilization plan. The predicted series reflects the cumulative changes predicted by equation (9). The number of quarters over which the actual and predicted series are compared was determined by the duration of the stabilization plans, the longest plan being the Israeli (which is still in place), and the shortest being the Cruzado plan (which was abandoned after four quarters). The simulation use the estimates of β and ρ reported in Table 5 (fourth column) for the case $k_1/k_2=2$. 24/ Charts 2 and 3 plot actual real per capita consumption against the dynamic forecasts.

The actual consumption boom (see Tables 1 and 6) is measured as the percent change in consumption from its level in the quarter preceding the beginning of the plan to its peak level during the duration of the program. For example, the peak in consumption during the Austral plan occurs six quarters into the plan (see Chart 3) and represents an increase of 16.2 percent (see Table 6). The predicted boom is similarly measured as the percent change in consumption from its pre-plan level to its predicted peak level during the duration of the program. For instance, for the Austral plan the predicted peak occurs one quarter into the plan (see Chart 3) and represents an increase of 10.0 percent (see Table 6). 25/

24/ With the exception of Chile, where there is broader variation in the estimate of ρ, the choice of transactions technology has little impact on the simulated changes in consumption.

25/ In the cases of the Argentine and Uruguayan tablitas, where the model predicts a sustained decline, the predicted "peak" (Table 6) is defined as the highest level during the forecast horizon, which occurs one quarter into the program.

As Table 6 and Charts 2 and 3 illustrate, the model's ability to mimic reality is mixed. The model performs poorly for the Argentine and Uruguayan Tablitas, as it misses the upturn in consumption entirely. For the remaining cases, the fit is better. As shown in Table 6 for the Chilean Tablita, the predicted increase in consumption is 29 percent versus an actual increase of 33 percent. For the Cruzado plan the prediction of a 11.4 percent increase compares favorably with an actual increase of 12.5 percent. For the Austral plan, the predicted rise is about 10 percent and the actual increase is 16 percent. The model overpredicts for both Israel and Mexico. For Israel and Mexico, the predicted increases in consumption are about 33 percent and 17 percent, respectively, versus actual increases of 24 percent and 14 percent. In sum, in five of the seven programs the model's predictions range anywhere between 59 and 136 percent of the actual increase in consumption.

The correlation between the dynamic forecasts and actual consumption, which provide an indication of the model's tracking ability, also present a mixed picture. As shown in Table 7, the correlations are positive, statistically significant and greater that 0.75 (irrespective of which rate of return measure is used) for four of the episodes; the Chilean Tablita, the Cruzado, Israeli, and Mexican plans. Not surprisingly (see Chart 2), the correlation between actual and predicted consumption is either relatively low or actually negative for the Argentine and Uruguayan tablitas, respectively. For the Austral plan, the model predicts reasonably well the initial rise in consumption (see Chart 3) but fails to capture much of the subsequent variability. Consequently, the correlation coefficient is negative, although not statistically significant.

V. Conclusions

Temporary policy in small open economies offers intuitive and clear predictions. Specifically, a temporary reduction in the nominal interest rate makes present consumption less expensive relative to future consumption, which induces consumers to substitute present for future c

onsumption thus generating a consumption boom. When the interest rate rises back to its initia
l level, consumption falls. Thus, policy temporariness introduces a transmission mechanism w
hereby monetary policy has real effects. To date, however, there is surprisingly little empirical
evidence examining the relevance of such policies in explaining the observed cycle in consum
er spending.

This paper has attempted to fill that gap by first examining if the conditions for the *exist
ence* of this channel are met. Estimates of the Euler equation that follow from a transactions c
osts model suggest that, in general, the intertemporal elasticity of substitution is low but statisti
cally different from zero. In addition, the variance decomposition of nominal interest rates into
its permanent and transitory components indicate that temporary shocks account for an import
ant share (between 40 and 90 percent) of the total variance of nominal interest rates. We thus c
onclude that nominal interest rates have been subject to substantial temporary shocks. Howeve
r, these two conditions (intertemporal elasticities of substitution significantly different from zer
o and the presence of temporary shocks to nominal interest rats) are necessary but not sufficien
t conditions for temporary policy to be quantitatively important.

To evaluate the *quantitative importance* of this channel of transmission we examine whet
her observed movements in nominal interest rates are large enough to generate, for the estimat
ed elasticities of substitution, changes in consumption that roughly match the observed ones.
Using the estimates of the elasticity of substitution and the Euler equations, we undertake dyna
mic simulations for seven major stabilization programs to see how well the predicted path of c
onsumption replicates the actual path immediately after the plans were implemented. We then
compare the predicted booms in consumption to the actual booms. While the model misses ent
irely the actual boom in two of the seven programs, in the remaining five programs the predicte
d booms range from 59 to 136 percent of the actual booms. We conclude, then, that for several

cases temporary declines in nominal interest rates appear to have played a key role in generating consumption booms.

It is important to note that, due to lack data, the analysis has not explicitly incorporated durable goods, which are believed to have played a key role in the consumption booms that have accompanied exchange rate-based stabilization. The presence of durable goods is likely to increase the estimates of the intertemporal elasticity of substitution, as found by Fauvel and Sampson (1991) for Canada. Furthermore, durable goods would also allow for intertemporal price substitution, which should accentuate the predicted effects of the model. Hence, if anything, this paper should be viewed as providing a lower bound for the importance of temporary shocks to nominal interest rates in explaining consumption booms.

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TABLE 1
Inflation Stabilization Programs and the Consumption Boom-Bust Cycle

Program	First Quarter	Inflation: Beginning of the Program (Percent Per Quarter)	Inflation: Lowest During the Program (Percent Per Quarter)	Consumption: Initial Boom (Percent Change)	Consumption: Later Contraction (Percent Change)
Tablitas:					
Argentine	79:Q1	27.8	14.5	17.5	-23.2
Chilean	78:Q1	9.4	-0.4	33.0	-26.0
Uruguayan	78:Q4	10.2	1.9	16.0	-20.4
Heterodox Plans:					
Austral	85:Q3	105.6	7.1	16.2	-15.9
Cruzado	86:Q2	51.0	2.2	12.5	-9.0
Israeli	85:Q3	48.7	1.9	23.9	-7.4
Mexican 1/	88:Q1	29.5	1.7	13.9	n.a.

Notes: Inflation at the beginning of the program refers to the inflation rate prevailing in the quarter before the start of th e program (for instance, 85:Q2 for the Israeli plan). The consumption boom is measured as the percent change in consu mption from its level during the first quarter of the plan to its peak level during the duration of the program. The subse quent contraction is measured as the percent change between the peak and trough levels.
1/ Data ends in 92:Q4. Preliminary evidence that suggests consumption peaked in 93:Q1 and has been declining ever si nce.

TABLE 2
Data, Sources and Descriptive Statistics

Sample, sources and number of observations	Variable	Mean	Standard Deviation
Argentina			
1978:Q1-1989:Q4, Central Bank,	Δc_t	-0.002	0.046
CEPAL, and International Financial	r_t	-0.057	0.153
Statistics, 48	i_t	0.331	0.244
	Δx_t	-0.027	0.173
Brazil			
1975:Q1-1987:Q4, Central Bank,	Δc_t	0.004	0.040
Cojunctura Economica, and	r_t	-0.027	0.088
International Financial Statistics, 52	i_t	0.136	0.072
	Δx_t	-0.002	0.118
Chile			
1975:Q1-1989:Q4, Central Bank,	Δc_t	0.005	0.044
Arrau and De Gregorio (1992),	r_t	0.019	0.044
International Financial Statistics, 60	i_t	0.119	0.104
	Δx_t	0.001	0.067
Israel			
1979:Q1-1991:Q2, Israel	Δc_t	0.004	0.040
Bureau of Statistics, and	r_t	0.056	0.104
International Financial Statistics, 50	i_t	0.245	0.205
	Δx_t	-0.010	0.102
Mexico			
1980:Q1-1992:Q1, Central Bank and	Δc_t	0.001	0.019
International Financial Statistics, 49	r_t	-0.016	0.042
	i_t	0.092	0.037
	Δx_t	-0.003	0.079
Uruguay			
1976:Q1-1989:Q4, Central Bank and	Δc_t	0.002	0.030
International Financial Statistics, 56	r_t	0.034	0.036
	i_t	0.158	0.030
	Δx_t	-0.008	0.081
Memo items:			
United States			
1947:Q1-1991:Q4, Department of	Δc_t	0.004	0.001
Commerce and Board of Governors	i_t	0.014	0.008
of the Federal Reserve, 180			

Notes: The consumer price index is used to construct measures of real balances, m_t, and ex-post real interest rates, r_t. C onsumption, c_t, and m_t, are on a per capita basis. The ratio of money to consumption, the inverse of velocity, is denoted by x_t and the Δs refer to rates of growth. The nominal interest rate is denoted by i_t and nominal and real interest rates a re expressed as quarterly returns; for example, the average quarterly nominal interest rate for Mexico is 9.2 percent, whi ch translates to about 42 percent per annum.

TABLE 3
1/k Times the Variance of k-differences of
Nominal Interest Rates

$(1/k)(\sigma_k^2/\sigma_1^2)$ for various k (Quarters)								
4	8	12	16	20	24	28	32	36
Argentina								
0.62	0.56	0.33	0.18	0.11				
(.010)	(.016)	(.022)	(.026)	(.030)				
Brazil								
0.56	0.37	0.21	0.19	0.23				
(.013)	(.022)	(.028)	(.032)	(.031)				
Chile								
0.44	0.49	0.50	0.49	0.50				
(.010)	(.016)	(.020)	(.023)	(.027)				
Israel								
1.78	2.02	1.69	1.71	2.14	1.71	0.82	0.11	0.13
(.012)	(.019)	(.026)	(.032)	(.036)	(.038)	(.038)	(.034)	(.034)
Mexico								
1.13	1.41	1.63	1.98	2.24	2.19	1.26	0.34	0.24
(.011)	(.019)	(.025)	(.030)	(.034)	(.035)	(.034)	(.024)	(.007)
Uruguay								
1.44	1.12	1.03	0.77	0.83	0.61			
(.010)	(.016)	(.022)	(.026)	(.030)	(.034)			

Notes: Standard errors, which appear in parentheses, were tabulated from Monte Carlo simulations. The simulations we re conducted for the relevant number of observations for each country from 600 trials.

TABLE 4
Cointegration Tests for Real Balances, Consumption,
and the Nominal Interest Rate

Country	Argentina	Brazil	Chile	Israel	Mexico	Uruguay
λ-max	26.083**	15.427	12.887	19.323*	14.584	14.048
Trace	33.827**	21.362	25.380	33.112**	26.059	16.612

Critical values for λ-max:

at 95 % confidence level = 20.967
at 90 % confidence level = 18.598

Critical values for trace:

at 95 % confidence level = 29.680
at 90 % confidence level = 26.785

Notes: The sample periods are those defined in Table 1. The number of lags in the vector autoregression is three and a constant term was the only predetermined variable. The critical values are those under the null of no cointegration. Th e source of the critical values is Osterwald-Lenum (1992); a ** indicates significance at the 95% confidence level, while a * denotes significance at the 90% level.

TABLE 5 The Parameters of Consumer Preferences: Estimates Using Deposit Rates of Interest

k_1/k_2	0.250	0.500	1.000	2.000	5.000	8.000
			ARGENTINA: 1978:Q1 to 1989:Q4			
ρ	0.194	0.194	0.196	0.196	0.191	0.185
	(0.080)	(0.077)	(0.075)	(0.075)	(0.072)	(0.069)
β	1.068	1.056	1.031	1.023	0.997	0.984
	(0.034)	(0.033)	(0.032)	(0.032)	(0.033)	(0.034)
J	7.441	7.274	6.891	6.780	6.498	6.416
	(0.384)	(0.401)	(0.440)	(0.450)	(0.483)	(0.492)
SSR	0.126	0.125	0.122	0.127	0.127	0.124
			BRAZIL: 1976:Q1 to 1988:Q4			
ρ	0.128	0.111	0.094	0.080	0.071	0.068
	(0.046)	(0.042)	(0.036)	(0.031)	(0.027)	(0.026)
β	1.023	1.021	1.017	1.013	1.008	1.007
	(0.035)	(0.040)	(0.048)	(0.057)	(0.066)	(0.069)
J	1.082	1.082	1.082	1.082	1.082	1.080
	(0.993)	(0.993)	(0.993)	(0.993)	(0.993)	(0.993)
SSR	0.059	0.059	0.060	0.060	0.060	0.060
			CHILE: 1976:Q1 to 1989:Q4			
ρ	0.795	0.731	0.628	0.492	0.310	0.236
	(0.186)	(0.166)	(0.137)	(0.102)	(0.062)	(0.046)
β	0.993	0.993	0.995	0.998	1.005	1.012
	(0.007)	(0.007)	(0.008)	(0.010)	(0.016)	(0.021)
J	7.553	7.447	7.305	7.152	7.047	7.097
	(0.374)	(0.384)	(0.398)	(0.413)	(0.424)	(0.419)
SSR	0.112	0.106	0.099	0.090	0.079	0.077

Notes: Standard errors are in parentheses. The number that appears in parentheses below the J-statistic is the probabilit y value under null hypothesis of the orthogonality of the residuals. SSR is the sum of squared residuals.

TABLE 5 (continued)
The Parameters of Consumer Preferences:
Estimates Using Deposit Rates of Interest

k_1/k_2	0.250	0.500	1.000	2.000	5.000	8.000
			ISRAEL: 1980:Q1 to 1991:2			
ρ	0.217	0.214	0.207	0.196	0.172	0.157
	(0.024)	(0.024)	(0.023)	(0.022)	(0.020)	(0.018)
β	0.990	0.990	0.991	0.993	0.997	0.999
	(0.015)	(0.015)	(0.016)	(0.018)	(0.020)	(0.018)
J	6.871	6.896	6.940	7.010	7.126	7.173
	(0.442)	(0.440)	(0.435)	(0.428)	(0.416)	(0.411)
SSR	0.051	0.051	0.051	0.050	0.048	0.047
			MEXICO: 1981:Q1 to 1991:3			
ρ	0.186	0.182	0.175	0.161	0.127	0.105
	(0.043)	(0.043)	(0.042)	(0.040)	(0.035)	(0.031)
β	1.022	1.023	1.023	1.024	1.027	1.029
	(0.013)	(0.012)	(0.014)	(0.015)	(0.019)	(0.023)
J	4.298	4.309	4.330	4.368	4.457	4.498
	(0.745)	(0.743)	(0.741)	(0.737)	(0.727)	(0.721)
SSR	0.013	0.013	0.013	0.013	0.013	0.013
			URUGUAY: 1977:Q2 to 1989:Q3			
ρ	0.043	0.046	0.052	0.138	0.165	0.167
	(0.111)	(0.110)	(0.108)	(0.093)	(0.084)	(0.075)
β	1.173	1.157	1.133	1.025	1.019	1.021
	(0.621)	(0.531)	(0.402)	(0.049)	(0.033)	(0.031)
J	8.948	8.944	8.936	8.495	7.861	7.333
	(0.256)	(0.257)	(0.257)	(0.291)	(0.345)	(0.395)
SSR	0.049	0.049	0.049	0.050	0.051	0.052

Notes: Standard errors are in parentheses. The number that appears in parentheses below the J-statistic is the probabilit y value under null hypothesis of the orthogonality of the residuals. SSR is the sum of squared residuals.

TABLE 6
Actual and Predicted Increases in Consumption 1/

Program	Change in Consumption Actual	Change in Consumption Predicted	Share of Predicted to Actual
Tablitas:			
Argentine	17.5	-2.5	--
Chilean	33.0	29.0	0.88
Uruguayan	16.0	-0.2	--
Heterodox Plans:			
Austral	16.2	9.6	0.59
Cruzado	12.5	11.4	0.91
Israeli	23.9	32.6	1.36
Mexican	13.9	17.3	1.24

Notes: Predicted changes in consumption are based on the parameter estimates reported in Table 5 for k_1/k_2. The latest included data point for Mexico is 92:Q4.

TABLE 7
Dynamic Simulations: Goodness of Fit

Program	Number of Quarters	Pairwise Correlation of Actual and Forecast Consumption	
		Deposit Rate	Stock Returns
Tablitas:			
Argentine	16	0.48 (0.23)	0.60 (0.21)
Chilean	15	0.86 (0.14)	0.88 (0.13)
Uruguayan	15	-0.73 (0.19)	n.a.
Heterodox Plans:			
Austral	9	-0.43 (0.34)	-0.45 (0.34)
Cruzado	10	0.77 (0.23)	0.80 (0.22)
Israeli	20	0.91 (0.10)	n.a.
Mexican	20	0.95 (0.08)	0.94 (0.09)

Notes: Standard errors are in parentheses.

TABLE A.1
The Parameters of Consumer Preferences:
Estimates Using Stock Returns

k_1/k_2		0.250	0.500	1.000	2.000	5.000	8.000
				ARGENTINA: 1978:Q1 to 1989:Q4			
ρ		0.203	0.331	0.410	0.385	0.357	0.342
		(0.366)	(0.335)	(0.312)	(0.243)	(0.216)	(0.204)
β		0.973	0.969	0.960	0.931	0.921	0.917
		(0.034)	(0.034)	(0.017)	(0.020)	(0.021)	(0.021)
J	14.817	14.186		11.854	8.231	8.219	8.330
		(0.038)	(0.048)	(0.105)	(0.313)	(0.313)	(0.304)
SSR		0.091	0.091	0.100	0.119	0.117	0.116
				BRAZIL: 1976:Q1 to 1988:Q4			
ρ		0.250	0.167	0.121	0.097	0.081	0.077
		(0.071)	(0.044)	(0.030)	(0.024)	(0.020)	(0.019)
β		0.993	0.991	0.989	0.987	0.985	0.984
		(0.017)	(0.026)	(0.035)	(0.044)	(0.052)	(0.054)
J	6.410		6.410	6.412	6.410	6.410	6.410
		(0.493)	(0.493)	(0.493)	(0.493)	(0.493)	(0.493)
SSR		0.047	0.049	0.050	0.050	0.051	0.051

Notes: Standard errors are in parentheses. The number that appears in parentheses below the J-statistic is the probabilit y value under null hypothesis of the orthogonality of the residuals. SSR is the sum of squared residuals.

TABLE A.1 (continued)
The Parameters of Consumer Preferences:
Estimates Using Stock Returns

k_1/k_2	0.250	0.500	1.000	2.000	5.000	8.000
			CHILE: 1976:Q1 to 1989:Q4			
ρ	2.535	3.058	2.465	1.328	0.516	0.337
	(0.756)	(0.698)	(0.428)	(0.236)	(0.095)	(0.063)
β	0.996	0.996	0.996	0.997	1.003	1.007
	(0.007)	(0.002)	(0.002)	(0.003)	(0.008)	(0.013)
J	7.037	4.061	1.960	4.247	6.126	6.546
	(0.425)	(0.773)	(0.962)	(0.751)	(0.525)	(0.478)
SSR	0.112	0.106	0.099	0.090	0.079	0.077
			MEXICO: 1981:Q1 to 1991:3			
ρ	0.031	0.034	0.038	0.041	0.037	0.032
	(0.147)	(0.135)	(0.116)	(0.089)	(0.052)	(0.036)
β	1.030	1.026	1.023	1.021	1.023	1.028
	(0.207)	(0.163)	(0.121)	(0.095)	(0.087)	(0.095)
J	6.724	6.726	6.728 6	6.727 6	6.722	6.717
	(0.458)	(0.458)	(0.458)	(0.459)	(0.458)	(0.459)
SSR	0.015	0.015	0.015	0.015	0.013	0.013

Notes: Standard errors are in parentheses. The number that appears in parentheses below the J-statistic is the probabilit y value under null hypothesis of the orthogonality of the residuals. SSR is the sum of squared residuals.