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Fisherian Transmission and Efficient Arbitrage Under Partial Financial Indexation

The Case of Chile

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Partial financial indexation in Chile has produced a system in which most bank deposits are 30-day nonindexed deposits or 90-day indexed deposits. This paper uses data on the interest rates of these financial assets to test the joint hypothesis of rational expectations, efficient arbitrage, and a time-invariant liquidity premium. The data are also used to test whether the indexed/nonindexed interest spread is an accurate predictor of future changes in inflation, as the Fisher effect dictates. The significant implications of this empirical analysis for monetary policy are discussed. [JEL E44, E43, E52]

> All variations in the value of the circulating medium are mischievous: they disturb existing contracts and expectations, and the liability to such changes renders every pecuniary engagement of long date entirely precarious.
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> —John Stuart Mill (1848, p. 544)

THIS PAPER studies the mechanism of partial financial indexation currently operating in Chile. The analysis focuses on the behavior of interest rates on indexed and nonindexed bank deposits, to determine whether there is arbitrage in financial markets and whether interest rate

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differentials contain significant information for forecasting inflation. A variation of traditional market efficiency tests, as reviewed in Begg (1982), is used to explore the empirical relevance of the joint hypothesis that there is efficient arbitrage in financial markets, that expectations of inflation are formed rationally, and that indexed deposits bear a premium attributed to liquidity considerations or imperfect indexation. Simple econometric methods are also used to explore whether expectations of inflation are fully transmitted to nominal interest rates, as the Fisher effect indicates. Specifically, empirical tests are performed to show whether the interest spread between indexed and nonindexed deposits is an accurate predictor of future increases in inflation, following the work of Fama (1990) and Mishkin (1990).

The empirical analysis undertaken in this paper has significant implications for key operational issues related to the desirability of indexing the financial system and the conduct of monetary policy. In particular, the results of the econometric work give some support to the view that financial indexation diminishes the risks associated with inflation and does not necessarily induce an unjustified increase in interest rates. Moreover, the results suggest that the quoted real rate of return on indexed assets cannot be treated as a proxy for the relevant real interest rate of the economy, and that the interest rate spread between indexed and nonindexed assets is a better indicator of the public's perception of the stance of monetary policy.

The analysis of the Chilean experience is also useful as an empirical complement to the protracted debate on the effects of financial indexation, which is closely related to the recurrent controversy involving the costs of inflation.[1] Advocates of financial indexation argue that indexed financial assets, by neutralizing the effects of unanticipated inflation, eliminate income-redistribution effects induced by inflation and the inflation-risk component of nominal interest rates. Thus, indexation reduces the degree of uncertainty affecting credit transactions and facilitates the operation of financial markets under inflationary conditions. In

[1] The adverse effects of inflation on the distribution of income between creditors and debtors and on the degree of uncertainty affecting credit markets have long been a cause of concern (see Mill (1848, chap. XIII)). Two classic papers on indexation and the costs of inflation are Friedman (1974) and Gray (1976). For a textbook discussion of financial indexation, see Gordon (1978), and for wage indexation, see Parkin and Bade (1988). The advantages of financial indexation are explored from the perspective of the more recent intertemporal equilibrium approach by Calvo and Guidotti (1989). The risk of indeterminacy of the price level due to excessive inertia and other issues regarding the public's aversion to indexed contracts and the determination of the relevant price index are reviewed by Leijonhufvud (1981).

contrast, critics of financial indexation argue that indexing the financial system fuels the inflationary process by speeding up price changes.[2] Moreover, when coexisting with nonindexed labor and goods markets, financial indexation causes income-redistribution effects that affect those firms and households facing indexed debts with nonindexed revenue and income streams.[3]

Many of the arguments for and against financial indexation can be simply stated in terms of the Fisher effect, according to which the nominal interest rate, i, is equivalent to the ex ante real interest rate, r, plus expected inflation, p^e: $i = r + p^e$. For a given ex ante real rate of interest, the Fisher effect dictates that a higher nominal interest rate applies to financial transactions if prices are expected to rise. The case in favor of financial indexation can be presented by considering an environment in which errors in inflationary expectations are not made systematically and markets operate efficiently. Under these conditions, the Fisher effect implies that the evolution of interest rates in an indexed system cannot deviate *systematically* from the evolution of interest rates in a nonindexed system. The main advantage of indexation is that it flattens the term structure of interest rates, because risk premia associated with the variance of errors in forecasting inflation are eliminated, thereby easing the operation of financial markets under conditions of high and variable inflation.

The case against financial indexation is often based on the hypothesis that full Fisherian transmission never takes place, because expectations underestimate inflation systematically, or that there is no efficient arbitrage in financial markets to eliminate the inflation-risk component of interest rates. Thus, indexation forces nominal interest rates to be higher than under normal market conditions and induces persistent disturbances to the cost of credit and the demand for real money balances. The demand for money fluctuates because of the rise in interest rates and also because indexation sends destabilizing signals that fuel inflationary expectations even if fundamentals remain unaltered.

[2] A stronger version of this critique, as explained by Leijonhufvud (1981), argues that because indexation is equivalent to forcing price expectations to exhibit unitary elasticity, any small price change could cause an exploding inflationary spiral. However, advocates of financial indexation have shown that although changes in inflation are likely to be larger in indexed economies, the price level and inflation are well defined and stable as long as the supply of money is not fully indexed (see Parkin and Bade (1986)).

[3] Critics of indexation would argue that the social costs of the redistributive effects that occur when interest rates are not indexed are less important than those that occur under a partial indexation system. Issues related to income redistribution and its social cost are not addressed in this paper.

As these arguments show, the conflicting views about financial indexation are founded on different notions regarding the efficiency of credit markets and the manner in which individuals formulate expectations about future inflation. In economies where financial markets are well organized and competitive, and where market participants formulate expectations rationally, indexation eliminates the inflation-risk component of interest rates and cannot result in systematically higher interest rates than a system without indexation. The empirical tests undertaken in this paper attempt to establish to what extent Chilean financial markets fit this framework, by providing evidence on the degree to which there is efficient arbitrage, and by testing the forecasting power of interest rate differentials regarding future increases in inflation implicit in the Fisher effect.

The paper is organized as follows. The next section describes the operation of Chile's financial indexation mechanism, with a brief description of its role in the banking crash of 1983, and illustrates its impact on the structure of the deposit base of the banking system. Section II studies arbitrage conditions in the financial system, taking advantage of the existing regulations that force short-term nonindexed assets to coexist with medium- and long-term indexed assets. Section III analyzes the information contained in the interest rate spread between indexed and nonindexed time deposits regarding future changes in the inflation rate, and also provides some evidence on the real interest rate that is relevant for financial decisions. The last section presents conclusions.

I. Chile: Indexation and Financial Markets

The first part of this section describes how financial indexation operates in Chile, focusing on some of the imperfections of the mechanism being used. The second part offers a brief review of the characteristics of the Chilean financial system that help illustrate the effects of the indexation regime.

Financial Indexation in Chile

Financial indexation in Chile is based on the use of a unit of account known as the "Unidad de Fomento" (UF), or development unit. The UF operates as an exchange rate between Chilean pesos and development units that is linked to the inflation rate with a delay of approximately one month and serves to denominate all indexed financial transactions. A typical indexed time deposit operates as follows. The deposit is entered

by converting the amount in Chilean pesos into development units at the current UF exchange rate. Interest is paid on the balance denominated in UF on the basis of an annual rate compounded monthly. This interest rate will be referred to as the "premium over UF" throughout the rest of the paper. Upon maturity, principal and interest are converted back into Chilean pesos at the corresponding UF exchange rate. The deposit is indexed because the value of the UF grows each month at about the same rate as last month's inflation, and for this reason the premium over UF is viewed as a real interest rate.

The timing and methodology by which the value of the UF are adjusted have important implications. The UF begins to be adjusted on the tenth day of month t, by a proportional amount each day, to ensure that by the ninth day of month $t + 1$ it has increased by as much as the price level did in month $t - 1$. Figure 1 depicts the ex post nominal yield that was actually paid on 90-day indexed time deposits, as a 3-month percentage rate that considers both the interest rate and the observed change in the UF, and the ex post 90-day inflation rate. The figure illustrates the degree of imperfection affecting the adjustment of the yield on indexed deposits to inflation because of the 1-month lag imbedded in the indexation mechanism.

Figure 1. *Ex Post Indexed Interest and Inflation, August 1986 to April 1989*
(90-day to 1-year deposits)

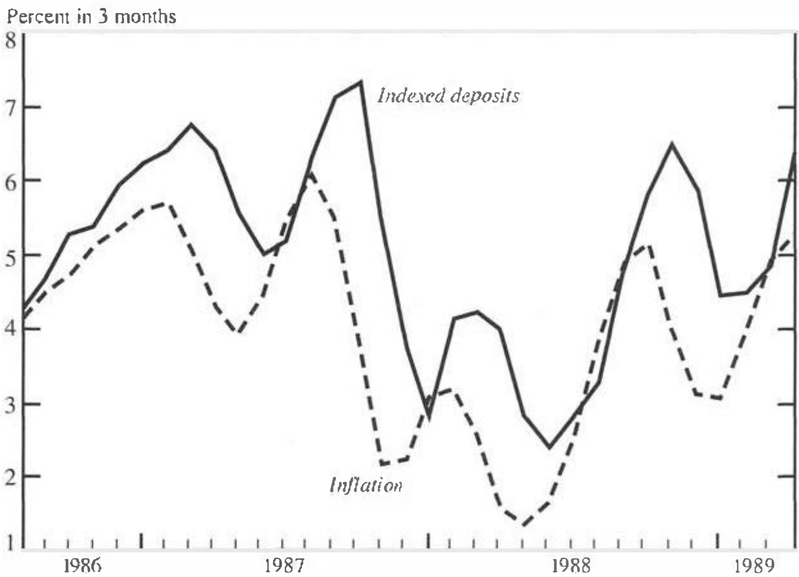
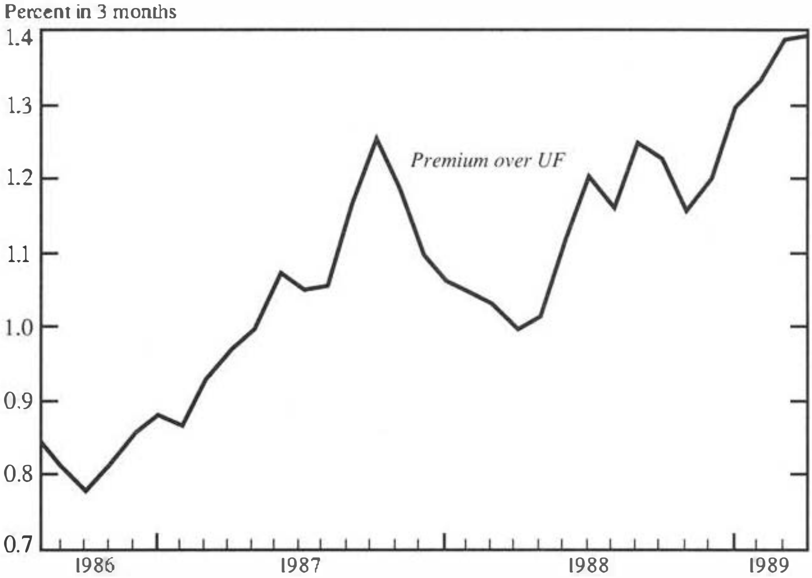


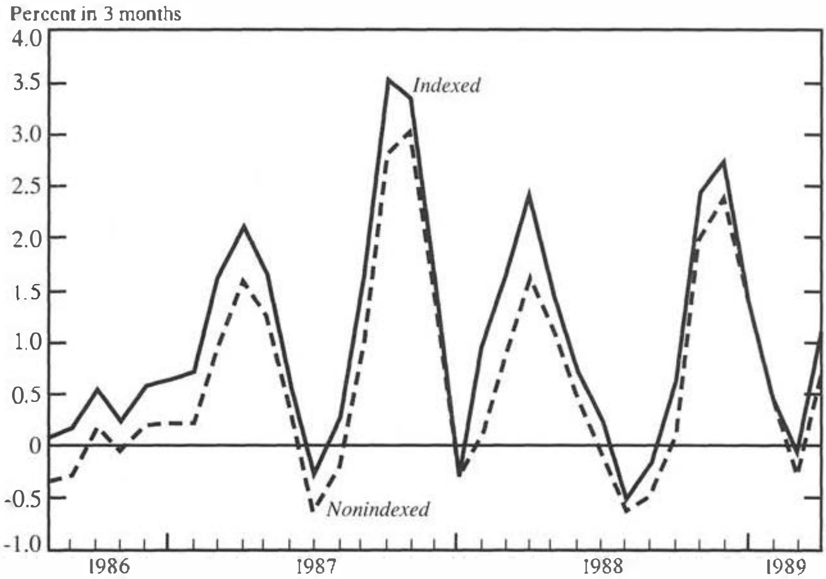
Figure 2. *Effective Premium on Indexed Deposits, August 1986 to April 1989*
(90-day to 1-year deposits)



The imperfection of the indexation mechanism is also reflected in the fact that, as shown in Figures 2 and 3, the premium over UF and the ex post real interest rate on indexed assets are not equivalent, as they should be under a perfect indexation system.[4] Figure 2 presents the premium over UF quoted on 90-day indexed time deposits as a percentage rate over 3 months. Figure 3 depicts the ex post real interest rates actually paid on 30-day nonindexed and 90-day indexed deposits as a percentage rate over 3 months (for 90-day indexed deposits, this ex post real interest rate corresponds to the difference between the two curves in Figure 1). As Figures 2 and 3 show, there is a significant discrepancy between the premium over UF and the ex post real interest rate paid on 90-day indexed deposits. For instance, in March 1989 a 90-day indexed time deposit offered a premium over UF of about 1.4 percent (5.7 percent on

[4] The ex post real interest rate on a 90-day indexed time deposit entered at date t is computed as the difference between the effective yield, which is the sum of the premium over UF quoted at t, plus the change in the value of the UF in the 90 days following t, and the inflation observed 90 days after t. The premium over UF is quoted in annual terms and compounds monthly. The rates in Figures 1–3 are returns over a period of 3 months.

Figure 3. *Ex Post Real Interest Rates, August 1986 to April 1989*
(30-day to 89-day and 90-day to 1-year deposits)



an annual basis), but the real return of this investment was in fact slightly negative. However, because the relevant real interest rates for economic decisions are the ex ante rates, which incorporate expectations of inflation rather than actual inflation, it is necessary to explore whether the premium over UF differs from the ex ante real interest rate on indexed deposits. This issue will be taken up later.

The Chilean indexation mecbanism was complemented with a set of regulations that govern the coexistence of indexed and nonindexed deposits. These regulations prevent full indexation of the money supply, as well as complete displacement of the Cbilean peso from the financial system by risk-averse agents seeking insurance against inflation.[5] Indexed time deposits are allowed only for maturities of 90 days or longer, whereas nonindexed interest-bearing deposits are legal for maturities of 30 days. The treatment is slightly different with respect to loans; indexed

[5]These regulations are consistent with some of the recent literature on monetary legal restrictions—for example, Smith (1988)—in which restrictions prevent the private sector from issuing close substitutes for money, so as to avoid large price fluctuations and adjustments in financial markets.

loans are allowed for terms as short as 30 days, but nonindexed lending operations are legal for maturities of less than 30 days.

Indexation and the Chilean Financial System

In general, Chile's financial markets are viewed as well organized and free from the distortions that result from excessive direct government intervention in many developing countries. Financial transactions are subject to strict supervision by the *Superintendencia de Bancos e Instituciones Financieras* (Superintendency of Banks and Financial Institutions), which strengthened its regulatory role after the major financial crisis that occurred in 1983.

The 1983 crash resulted from the accumulation of a large stock of nonperforming assets in the banking system, apparently as a result of high real interest rates, high debt-to-equity ratios of business firms, poor risk management of the bank's assets, and a significant decline in economic activity during 1981–82. The subsequent sharp depreciation of the peso in 1982–83 caused further damage by generating widespread defaults on loans contracted in foreign currency, and by causing significant operational losses to banks financing loans denominated in domestic currency with resources borrowed in foreign currencies. After 1983 the financial system started to recover slowly from the crisis, following the gradual improvement in economic conditions and the adoption of policies aimed at increasing liquidity and restoring solvency.[6]

The role that the indexation system may have played in generating the 1983 crash must be evaluated with caution, and a thorough discussion of the issue is beyond the scope of this paper. Financial indexation had been in place without causing any major disturbance long before the crisis started; it was maintained throughout and continued after the crisis ended.[7] For firms and banks involved in industries with declining relative prices, indexed loans and deposits were an important factor contributing to liquidity problems and inability to repay debts. In this way, the indexation system helped to speed up the rate at which the crisis spread. However, the prices of goods sold by the average firm increased at least

[6]These policies included the intervention and closing of banks, provision of massive liquidity support by the Central Bank, temporary public guarantee of bank deposits, establishment of credit lines in support of private debtors, and a recapitalization program based on voluntary sale of a fraction of bad loans to the Central Bank.

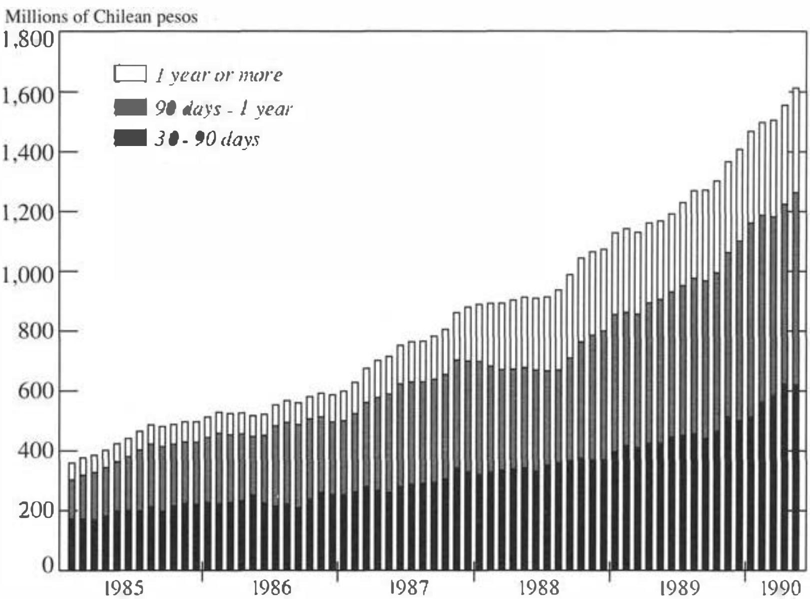
[7]Daily adjustments in the UF on the basis of the inflation from the previous month have been undertaken since 1977. The UF was first introduced in January of 1967 under a system of quarterly adjustments.

as fast as indexed debt commitments, and possibly faster because of the 1-month lag in the indexation system. Thus, it is likely that the widespread inability to service debts had more to do with the recession and poor banking practices than with indexation itself.[8] Moreover, by partially eliminating the risk of unanticipated inflation, financial indexation gave credibility to some of the policies used to attack the crisis, particularly the guarantee on bank deposits and the recapitalization programs.

There is evidence suggesting that financial indexation has induced important permanent changes in the structure of the Chilean financial system. In principle, partial financial indexation differentiates a subset of financial assets on the basis of the increased degree of protection against inflation that they provide. Consequently, risk-averse individuals are likely to reveal a preference for indexed assets in the allocation of their portfolio.

A review of the term structure of time deposits supports this hypothesis. Figure 4 illustrates the term structure of deposits in the financial

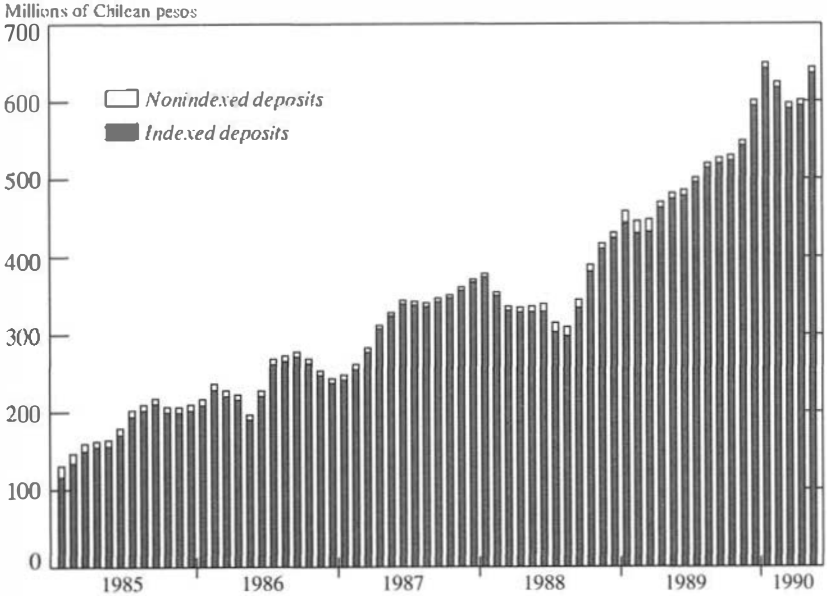
Figure 4. *Term Structure of Time Deposits, January 1985 to May 1990*
(Monthly averages)



[8]Unless it could be shown that a nonindexed system would have underestimated future inflation substantially in setting interest rates, resulting in lower interest rates than under indexation, it is likely that the chain of business failures would have followed in a similar manner even without indexation.

system using data on monthly average stocks of bank deposits for the period January 1986–May 1990. It shows that approximately 80 percent of all time deposits are concentrated in terms between 30 days and 1 year.[9] On average, 30- to 90-day deposits account for 38 percent of total time deposits, while 90-day to 1-year deposits account for 41 percent.[10] Figure 5 details the distribution of 90-day to 1-year deposits in terms of indexed and nonindexed deposits, and shows that the vast majority of them are indexed (about 98 percent on average).[11] Thus, the existing regime of partial financial indexation, operating in an environment of moderately high and variable inflation rates, has resulted in a market in which

Figure 5. *Structure of 90-Day Deposits, January 1985 to May 1990*
(Monthly averages)



[9]The secular growth in deposits with more than 1-year maturity reflects the growth of deposits from the pension funds, which should not be viewed as pertaining to the private sector but to nonbank financial intermediaries.

[10]Deposits in U.S. dollars are allowed for maturities of 30 days or longer, with adjustments in the exchange rate of the Chilean peso vis à vis the U.S. dollar that fluctuate around the difference between the movement of the UF and an estimate of foreign inflation. During the period January 1986–May 1990, these deposits were equivalent to less than one third of the total of deposits denominated in Chilean currency.

[11]A similar result is obtained from the data on total daily banking operations of time deposits.

depositors seek the best possible protection against inflation, which is found either in short-term, nonindexed deposits that are highly liquid, or in the most liquid indexed deposit available.

The fact that individuals hold substantial amounts of 30-day nonindexed deposits suggests that, in addition to liquidity considerations, they seem to regard them as relatively safe from the eroding effects of inflation. This may partly reflect the policy of setting an indicative nominal interest rate for 30-day deposits according to past and expected inflation plus a premium, which apparently was followed until June of 1987. However, as documented in the next section, efficient arbitrage with 90-day indexed deposits has also played a major role in ensuring that 30-day nonindexed deposits provide a good safeguard against inflation.

II. Efficient Arbitrage in Chilean Financial Markets

This section presents the results of some econometric tests that attempt to measure the degree of arbitrage present in Chilean financial markets. These tests focus on the effective rates of return paid on 30-day nonindexed and 90-day indexed time deposits. The rationale for dealing only with these two maturities is that, as noted above, the majority of time deposits in the Chilean financial system fall into these two categories.

Consider a perfectly competitive financial market in which there is no uncertainty. The public is endowed with perfect foresight regarding future values of short-term interest rates and the evolution of the UF, and arbitrage equalizes the ex post effective returns of 3-month investments on 30-day nonindexed deposits and a 90-day indexed deposit:

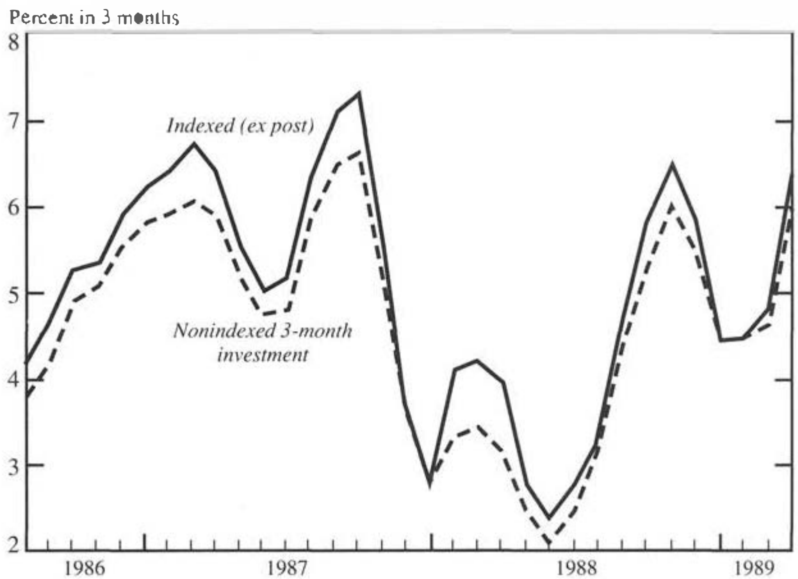
$$(1 + i_t)(1 + i_{t+1})(1 + i_{t+2}) = (1 + {}_tUF_{t+3})(1 + r_t). \tag{1}$$

The notation here is as follows: i_t is the monthly nominal interest rate on a 30-day nonindexed time deposit entered at date t, with i_{t+1} and i_{t+2} denoting the same rate as quoted 30 and 60 days after t; ${}_tUF_{t+3}$ is the percentage change actually observed in the UF in the 90 days following t; and r_t is the 90-day premium over UF quoted at date t on an indexed time deposit.[12] Thus, the left-hand side of equation (1) measures the actual effective yield of a 3-month investment in deposits with a maturity of 30 days, referred to here as *ex post nominal interest rate*, whereas the right-hand side measures the actual effective yield in Chilean pesos of a 3-month investment in a 90-day indexed deposit, referred to as *ex post indexed interest rate*.

[12]These rates are published in annual terms, but in evaluating equation (1), one must be aware that they are compounded monthly.

Figure 6 depicts the monthly evolution of the ex post nominal and indexed interest rates, as measured by the left- and right-hand sides of equation (1), respectively. The figure shows that both rates follow similar trends, although indexed deposits seem to be offered at a premium—the effective yield on indexed deposits is systematically higher than that on nonindexed deposits. Indexed deposits may need to be offered at a premium because the two financial assets are not perfect substitutes in terms of liquidity. For simplicity, this liquidity premium is assumed here to be a time-invariant constant. Moreover, indexed deposits may also offer a premium because the 1-month lag of indexation introduces an element of risk by making the real return on these deposits vary with the difference in the inflation rates between the month before the deposit is acquired and the month before it matures. When inflation rates are very volatile, or during turning points from increasing to decreasing inflation, this difference can be substantial and difficult to estimate. Thus, indexed deposits may pay a premium that follows a stochastic process determined by the agents' perception of the risk involved in the variance of monthly inflation due to imperfect indexation.

Figure 6. *Ex Post Nominal and Indexed Interest, August 1986 to April 1989*
(30-day to 89-day and 90-day to 1-year terms)



Even when the existence of a time-invariant liquidity premium is taken into account, the match of the two rates of return in equation (1) is not exact in actual data because the public and the banks do not know the future values of interest rates and the UF with certainty, and thus in practice expectations must replace the values of future variables in equation (1). Incorporating these arguments, taking logs of both sides of (1), and assuming that all rates involved are relatively small, the arbitrage condition can be rewritten as

$$i_t + E_t i_{t+1} + E_t i_{t+2} = E_t UF_{t+3} + r_t - LP. \tag{2}$$

Here, future short-term interest rates and the 3-month growth of the UF are expectations formed with the information available at date t,[13] and LP is a time-invariant liquidity premium on 90-day deposits.[14]

In order to perform formal econometric tests of arbitrage in Chilean financial markets using equation (2), it is necessary to introduce an assumption regarding the formation of expectations. The tests will analyze jointly efficient arbitrage, a time-invariant liquidity premium, and a particular hypothesis of how expectations are formed. The high degree of organization and development of Chilean financial markets suggests that both the banks and the public are potentially capable of forming an educated guess about future inflation by making extensive use of the information publicly available. Thus, the case of Chile provides a good basis for supporting the hypothesis that expectations of participants in financial markets are formed rationally.

If expectations are rational, forecasting errors are random variables that follow a stochastic process determined by how far in the future individuals must form expectations on the basis of the information available today. In the case that at period t expectations need to be formed for variables at $t + 1$, the errors follow a white-noise process, whereas in cases that require expectations for variables dated $t + 2$ or later, as in this paper, the errors follow a moving average representation.[15] Adding and

[13]This paper focuses on efficient arbitrage when expectations are formed on the basis of publicly available information, which in the finance literature is viewed as semistrong market efficiency since it ignores the role of "inside" information.

[14]Most of the econometric work discussed later in the paper only tests for the existence of this time-invariant liquidity premium, jointly with market efficiency and rational expectations, and cannot distinguish any form of stochastic premium associated with liquidity or the variance of inflation. Econometric results are viewed simply as evidence that the data can or cannot reject *this* joint hypothesis, despite the fact that other hypotheses may be observationally equivalent.

[15]The author thanks Charles Adams for clarifying this point.

subtracting the actual values of expected variables to equation (2) results in the following expression:

$$i_t + i_{t+1} + i_{t+2} = {}_tUF_{t+3} + r_t - LP + u_t, \tag{3}$$

where u_t is a random variable that includes forecasting errors with regard to future 30-day nonindexed interest rates (e^i_{t+1} and e^i_{t+2}) and the error in predicting the 90-day growth of the UF (e^{uf}_{t+3}), $u_t = e^i_{t+1} + e^i_{t+2} - e^{uf}_{t+3}$. Given the 1-month lag affecting the adjustments in the UF, it can be shown that u_t should follow a first-order moving average process in order to be consistent with rational expectations.[16] The moving average error reflects the fact that innovations in monthly interest rates or the UF that occur between date t and date $t + 3$ constitute information not available at t that is relevant for explaining the relative returns of the two investments.

The first test of the joint hypothesis of efficient arbitrage, rational expectations, and a time-invariant liquidity premium is conducted as follows. Expression (3) is rewritten to define the dependent variable as the ex post differential of the effective yields of the two investments in time deposits being considered, which is referred to as $DR9030F$. If the data support the hypothesis, $DR9030F$ should follow a stochastic process characterized as a first-order moving average with a constant term:

$$DR9030F \equiv (i_t + i_{t+1} + i_{t+2}) - ({}_tUF_{t+3} + r_t) = -LP + u_t, \tag{4}$$

with $u_t = v_t + \theta v_{t-1}$, and v_t is i.i.d. The test consists of identifying the time-series process that characterizes $DR9030F$ in the Chilean data to see if it satisfies the properties consistent with efficient arbitrage, rational expectations, and a time-invariant liquidity premium on indexed deposits. This test will only support or reject the three hypotheses jointly.

The time-series process corresponding to $DR9030F$ is identified following the Box-Jenkins method. The plots of the autocorrelation and partial autocorrelations show that the process is not white noise, with a Q-statistic = 47.7 for 25 lags, but do not provide a clear indication of the nature of the process. The results of estimating a first-order moving average (MA(1)) process with a constant using monthly Chilean data for the period August 1986 to April 1989 are the following:[17]

[16] Since $e^{uf}_{t+3} \equiv e^{\pi}_{t+1} + e^{\pi}_{t+2}$, where e^{π} is the forecasting error with regard to inflation, it follows that, assuming that i and π are white noise, $\mathrm{cov}(U_t U_{t+1}) = \sigma^2_i + \sigma^2_{\pi}$, where σ^2_i and σ^2_{π} are the variances of the forecasting errors related to interest rates and inflation; and $\mathrm{cov}(U_t U_{t+k}) = 0$ for all $k > 1$.

[17] The numbers in brackets in all regression results are t-statistics; those marked with an asterisk (*) are significant at the 5 percent level. The hypothesis of zero autocorrelation of regression residuals for up to ten lags was tested using the method of Box and Jenkins.

$$DR9030F = -0.353 + 0.496 v_{t-1} + v_t$$
$$(-11.05)^* \quad (2.762)^*$$

$$\text{Adj. } R^2 = 0.234 \qquad DW = 1.793 \qquad SE = 0.183 \qquad F = 10.77.$$

(In this and succeeding regression results, adj. R^2 is the adjusted coefficient of determination, DW is the Durbin-Watson statistic, and SE is the standard error.)

These results show that the data give some support to the joint hypothesis stated above, but formally the hypothesis is rejected because the residuals, v_t, are not white noise. Estimation of an MA(1) process with a constant produces a time-invariant liquidity premium that is very close to the average of $DR9030F$ in the data, -0.35. Both coefficient estimates in the regression are statistically significant, and the estimate of θ is consistent with the sample autocorrelation for the first lag estimated at the identification stage ($\hat{\theta} = 0.5$ implies $\hat{\rho}_1 = 0.41$, which compares with $\hat{\rho}_1 = 0.48$ from the identification procedure). Despite these favorable results, the residuals, v_t, seem to follow a complicated stochastic process in which v_{t-3} is significant for explaining v_t (the partial autocorrelation coefficient for v_{t-3} is -0.51, with a standard error of 0.17). This pattern of autocorrelation could be indicative of the presence of quarterly seasonality in the data, or may suggest the existence of a stochastic premium associated with the riskiness of imperfectly indexed deposits, given wide changes in monthly inflation. Moreover, reflecting the ambiguity of the results of the identification procedure, estimating a second-order autoregressive process (AR(2)) produces a similar estimate of $-LP$ and statistically significant autoregressive terms.[18] Thus, in this case the data cannot distinguish clearly between an MA(1) and an AR(2) process.

An alternative test of the hypothesis under discussion is performed by estimating equation (3) directly:

$$(i_t + i_{t+1} + i_{t+2}) = \alpha_0 + \alpha_1({}_tUF_{t+3} + r_t) + u_t, \tag{5}$$

where $u_t = \theta v_{t-1} + v_t$, and v_t is white noise. To be consistent with the hypothesis mentioned, the coefficient estimates should be $\hat{\alpha}_0 < 0$, since $\alpha_0 = -LP$ is the time-invariant liquidity premium, and $\hat{\alpha}_1 = 1$, if there is efficient arbitrage under rational expectations. The error term, v_t, must be a serially uncorrelated random variable. Moreover, any variables added to equation (5) that are judged to be publicly available information

[18] However, by estimating an autoregressive moving average process (ARMA(2, 1)), it can be shown that the restriction that the two autoregressive terms are zero cannot be rejected by the data at the level of 5 percent significance (the corresponding F-statistic is $F(2, 27) = 3.363$).

dated t or earlier must exhibit coefficients that statistically are not significantly different from zero, since presumably these variables are part of the information set used to formulate the expectations. Variables such as current and past values of the inflation rate, the exchange rate, the UF, and the ex ante yields of indexed and nonindexed deposits were introduced to the equation to confirm that they do not convey significant additional information.

Estimation of equation (5) for the same sample period as the previous test produces the following results:

$$i_t + i_{t+1} + i_{t+2} = \underset{(-0.45)}{-0.05} + \underset{(41.64)^*}{0.94}({}_tUF_{t+3} + r_t) + \underset{(2.72)^*}{0.49}v_{t-1} + v_t$$

$$\text{Adj. } R^2 = 0.98 \qquad \text{DW} = 1.66 \qquad \text{SE} = 0.169 \qquad F = 879.3.$$

These results support the hypothesis that there is efficient arbitrage in Chilean financial markets in the sense that $\hat{\alpha}_1 \approx 1$. This implies that, everything else unchanged, changes in the effective return of indexed deposits are reflected in almost equally proportional changes in the effective return of nonindexed deposits. The regression as a whole has a very high explanatory power and a low standard error, and it produces residuals that follow a stochastic process according to which the autocorrelation between v_t and v_{t-3} is much weaker than in the previous case (the partial autocorrelation coefficient is -0.35, with a standard error of 0.18, and is not significantly different from zero at the 7 percent level). However, the regression fails to detect a significant time-invariant liquidity premium, and thus the joint hypothesis being tested is once again rejected by the data.

If equation (5) is re-estimated imposing the restriction that the liquidity premium is fixed at -0.35 (the average of $DR9030F$), the results show that the data cannot reject the restriction, and that changes in the ex post indexed interest rate are still reflected in close to one-to-one changes in the ex post nominal interest rate.[19] However, the joint hypothesis of efficient arbitrage, rational expectations, and a time-invariant liquidity

[19] The t-statistics for α_1 in the unrestricted and restricted estimates of equation (5) are unusually high. Augmented Dickey-Fuller statistics were computed for the variables in the regression, and the presence of a unit root that could account for large t-statistics was not detected. The regressions do not involve an identity, as Figure 6 shows, but the fact that movements in the ex post nominal interest rate mimic so precisely movements in the ex post indexed interest rate casts some doubts on whether these interest rates are fully determined by market forces in the sample; as noted in the text, in June 1987 the Central Bank ended a policy of setting indicative interest rates for nonindexed deposits according to a rule based on the UF premium.

premium is rejected once again because the residuals are not white noise. In this case the regression output is summarized as follows:

$$i_t + i_{t+1} + i_{t+2} = \underset{}{-0.35} + \underset{(162.63)^*}{0.99({}_tUF_{t+3} + r_t)} + \underset{(2.80)^*}{0.50\nu_{t-1}} + \nu_t$$

$$\text{Adj. } R^2 = 0.98 \qquad \text{DW} = 1.80 \qquad \text{SE} = 0.182 \qquad F = 1513.9.$$

The F-test for the restriction that $\alpha_0 = -0.35$ is $F_{(1,30)} = 5.889$, and this implies that the restriction cannot be rejected by the data at the level of 1 percent significance. The residuals, ν_t, display a pattern of autocorrelation between ν_t and ν_{t-3} similar to that found in the first test (the partial autocorrelation coefficient for ν_{t-3} is -0.49, with a standard error of 0.176).

Given that the first test and the restricted version of the second test detect third-order serial autocorrelation of the residuals, a third test of the joint hypothesis under study is performed by imposing a pattern of quarterly seasonality in the form of a multiplicative seasonal MA term at lag 3, SMA$(3)_t$. An empirical justification for this, as Figure 6 illustrates, is that the differential between ex post indexed and nominal interest rates tends to be lower in the second and fourth quarters, and higher in the first and third quarters. One theoretical interpretation of this seasonality pattern is that it could reflect the existence of a stochastic premium on indexed assets. As noted before, the return on indexed deposits is affected by the difference in the inflation rate between the month before the deposit is entered and the month before it matures. When monthly inflation rates fluctuate sharply, as has been the case in Chile according to Figure 1, this embodies a significant risk. Another interpretation of seasonality in the interest differential is that it represents changes in monetary policy that are not part of the information used by individuals to formulate their expectations. This hypothesis is not tested here, but it is interesting to note that Chile relies heavily on open market operations in indexed assets to manage monetary policy.

The estimated time-series process for $DR9030F$ with the seasonality component is

$$DR9030F_t = \underset{(-13.64)^*}{-0.37} + \underset{(4.84)^*}{0.68\nu_{t-1}} - \underset{(-4.98)^*}{0.88\text{SMA}(3)_t} + \nu_t$$

$$\text{Adj.} R^2 = 0.475 \qquad \text{DW} = 2.034 \qquad \text{SE} = 0.152 \qquad F = 15.46.$$

In contrast to the previous results, the introduction of quarterly seasonality eliminates completely the pattern of correlation in the error terms.[20]

[20]The restriction that the seasonal adjustment parameter is zero is clearly rejected by the data at the 1 percent and 5 percent significance levels ($F_{(1,30)} = 15.21$).

Thus, if the additional assumption of quarterly seasonality is taken into account, the data cannot reject the joint hypothesis of efficient arbitrage, rational expectations, and a premium on 90-day indexed time deposits consistent with time-invariant and stochastic components. This is a very weak hypothesis, however, since the stochastic element embodied in the seasonal moving average has many interpretations, and could be viewed simply as a rejection of rational expectations and market efficiency.

To summarize, the results of the tests performed here provide some evidence suggesting that efficient arbitrage takes place in Chilean financial markets, in the sense that changes in the ex post indexed interest rate are associated with changes in the ex post nominal interest rate of almost identical magnitude. There is also some evidence that market participants tend not to make systematic errors in forecasting short-term interest rates or the UF. However, the joint hypothesis of efficient arbitrage, rational expectations, and a time-invariant liquidity premium is not supported by the data—the relevant residuals display negative autocorrelation at the third lag, or, when the residuals are white noise, the time-invariant premium is not statistically significant. Although this joint hypothesis fails, the results lend some support to the view that financial indexation minimizes the risks associated with inflation, since no inflation-risk premium is detected in the effective yield paid on nonindexed deposits. The only premia detected in the tests are paid on indexed deposits, and are viewed as stemming from liquidity preference and imperfect indexation. The negative constant in the stochastic process of $DR9030F$ suggests that indexed deposits pay a time-invariant premium of approximately 1.5 percent a year. The coefficient of quarterly seasonality estimated in the last regression is also negative, suggesting—as one possible interpretation—the presence of a stochastic premium on indexed deposits.

Thus, the evidence from Chilean data suggests that indexation, at work in an environment where there is evidence of efficient arbitrage, facilitates financial operations by eliminating the inflation-risk component of nonindexed interest rates. Given rational expectations, efficient arbitrage, and a full Fisher effect, the elimination of financial indexation would only alter interest rates to the extent that inflation-risk premia are re-introduced. It is therefore important to search for further evidence on Fisherian transmission and rational expectations.

III. The Interest Spread and Increases in Inflation

It has been argued here that indexation does not induce systematic, unjustified increases in interest rates in response to inflation when expectations are rational and there is efficient arbitrage, because the Fisher

effect in a nonindexed system would produce similar increases. The tests conducted in the previous section suggest that Chilean financial markets conform well to some elements of the hypothesis of efficient arbitrage and rational expectations, but do not produce evidence on the validity of the Fisher effect, nor do they give a clear description of the relationship between the returns paid on indexed and nonindexed deposits and expectations of future inflation. In this section, a series of tests are undertaken to explore the empirical relevance of the Fisher effect and one of its key implications under the assumption of rational expectations—namely, that the differential between the premium over UF of an indexed deposit and the nominal interest rate of a nonindexed deposit should be a precise indicator of future changes in inflation. These tests also have the operational value of helping to establish whether the UF premium can be treated as equivalent to the relevant real interest rate, and thus whether it is proper to use it as an instrument or indicator in the design of monetary policy.

The informational content of the spread between indexed and nonindexed interest rates can be extracted by estimating a functional relationship that follows from the Fisher effect. This approach is a variation of the methodology applied recently to study inflation forecasts based on the term structure of nominal interest rates by Fama (1990) and Mishkin (1990).

The Fisher equations for 30-day nonindexed and 90-day indexed time deposits can be expressed as follows:

$$E_t P_{t,30} = i_{t,30} - R_{t,30} \tag{6}$$

$$E_t P_{t,90} = (r_{t,90} + E_t UF_{t,90}) - R_{t,90}, \tag{7}$$

where E_t denotes a rational expectation conditional on publicly available information available at date t; $P_{t,30}$ and $P_{t,90}$ are the inflation rates for the 30 and 90 days that follow t; $R_{t,30}$ and $R_{t,90}$ are the ex ante real interest rates of each type of deposit; $i_{t,30}$ is the monthly nominal interest rate quoted at date t on a 30-day nonindexed deposit; $r_{t,90}$ is the quarterly premium over UF quoted at date t on a 90-day indexed deposit; and $UF_{t,90}$ is the growth of the UF during the 90 days following t.

Following Mishkin (1990), equations (6) and (7) are combined in an expression that provides a framework useful for extracting the information about expectations of future inflation contained in the interest spread. This is done by imposing the conditions that expectations are formed rationally and that ex ante real interest rates fluctuate around constant averages over time. Subtracting (6) from (7) and imposing these two conditions yields the following result:

$$P_{t,90} - UF_{t,90} - P_{t,30} = \alpha_0 + \alpha_1(r_{t,90} - i_{t,30}) + \nu_t, \tag{8}$$

where

$$v_t = e^{\pi}_{t,90} - e^{uf}_{t,90} - e^{\pi}_{t,30} - e^{r}_{t,90} + e^{r}_{t,30}.$$

This expression must be treated with caution because the dependent variable on the left-hand side has a special interpretation. The dependent variable is the *uncovered* increase in inflation (*UNCINFCH*) during the period from 30 to 90 days after t. It is considered uncovered because it represents the residual change in inflation after taking into account the protection that indexation gives to 90-day deposits (that is, *UNCINFCH* is equal to the difference between the inflation rates minus the correction in the UF).

Expression (8) indicates that, if expectations are rational, the ex ante real interest rates vary around constant means, and the Fisher equations (6) and (7) hold, then *UNCINFCH* must be equal to the sum of the average differential in ex ante real interest rates, $\alpha_0 \equiv \bar{R}_{90} - \bar{R}_{30}$; the spread in the quoted returns of indexed and nonindexed deposits, $\alpha_1 \equiv 1$; and a random error, v_t. The error term is a combination of expectational errors pertaining to the 30- and 90-day inflation rates ($e^{\pi}_{t,30}$ and $e^{\pi}_{t,90}$), the change in the UF ($e^{uf}_{t,90}$), and the discrepancy between the period-by-period and mean values of the ex ante real interest rates ($e^{r}_{t,30}$ and $e^{r}_{t,90}$). To be consistent with rational expectations, this disturbance must have zero mean and constant variance, and may follow a second-order moving average process, because the data correspond to monthly observations and expectations need to be formed for variables up to three months ahead—in this case, an expectation for inflation in the third month after t is needed, whereas in the tests of efficient arbitrage only inflation expectations up to the second month ahead were required.

The economic interpretation of the movements implicit in equation (8) is the following. For a given value of α_0, a widening of 1 percentage point in the differential between the 90-day premium over UF and the 30-day nominal interest rate indicates that individuals expect inflation in 90 days to increase by 1 percentage point in addition to the expected growth in the UF (about 4 percentage points on an annual basis). Therefore, if the estimate of equation (8) produces an error term with an MA(2) representation and shows that the hypothesis $\hat{\alpha}_1 \neq 0$ is rejected and the hypothesis $\hat{\alpha}_1 = 1$ cannot be rejected, then the data would indicate (1) the spread of interest rates is a precise indicator of future increases in uncovered inflation as a joint test of the Fisher effect and rational expectations; and (2) the average of ex ante real interest rates is constant over time. This second result would suggest that the UF premium is not a good proxy for the relevant real interest rate because, as Figure 2 illustrates, the UF

premium exhibits an increasing trend and does not fluctuate around a time-invariant mean.[21]

Equation (8) was estimated by ordinary least squares using data from August 1986 to January 1990, July 1987 to January 1990, and August 1986 to December 1987. The last sample is dominated by observations from the period during which the Central Bank provided indicative interest rates for 30-day nonindexed time deposits, and consequently, the policy regime may have intervened with the Fisherian transmission of market-determined, rational expectations of inflation to interest rates. To allow for quarterly seasonality, the equation for the period August 1986–January 1990 was also estimated with a multiplicative seasonal moving average term at the third lag. To test the null hypothesis that $\hat{\alpha}_1 = 1$, two procedures were followed: a standard "t"-test, and an F-test that combined the sum of squared residuals of unrestricted and restricted versions of the model. The results of all estimates are summarized in Table 1.

With one exception, the results show that the hypothesis that the coefficient $\hat{\alpha}_1$ is significantly different from zero but insignificantly different from unity cannot be rejected by the data, and that the residuals follow an MA(2) process. Surprisingly, the exception is not the period of indicative interest rates but the period from July 1987 to January 1990. For the sample August 1986 to January 1990, the unrestricted estimate for $\hat{\alpha}_1$ is 1.38, with a standard error of 0.26. This implies that the hypothesis $\hat{\alpha}_1 = 0$ is rejected with less than 1 percent significance, and the hypothesis $\hat{\alpha}_1 = 1$ cannot be rejected with less than 1 percent significance. Moreover, estimating the model for all the sample periods with the restriction that $\alpha_1 = 1$ produces F-statistics according to which the data cannot reject the restriction at 1 percent significance. In general, these results suggest that the joint hypothesis of Fisherian transmission and rational expectations cannot be rejected by the data.

Figure 7 depicts the actual and predicted values of *UNCINFCH* produced by the unrestricted model for the sample August 1986 to January 1990. This figure illustrates clearly the predictive power of the regressions based on the spread between indexed and nonindexed interest rates to forecast the evolution of the uncovered increase in inflation. One of the interesting operational implications of this result is that the market-determined interest spread between indexed and nonindexed deposits is an accurate indicator of the public's perception of future changes in

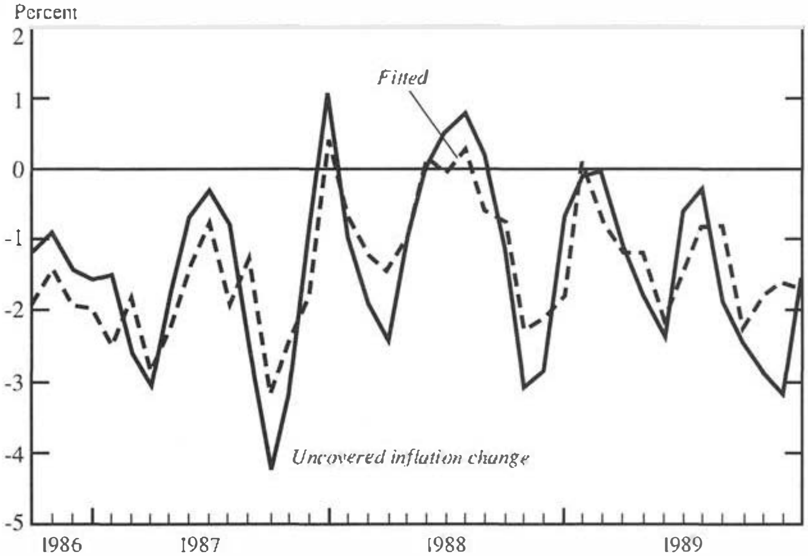
[21] This conclusion would not follow if the UF moved exactly with inflation month by month. In this case, the dependent variable in equation (8) collapses to $-P_{t,30}$, and the UF premium becomes identical to the ex ante real interest rate.

Table 1. *Estimates of Uncovered Inflation Change Equations*

Sample	$\hat{\alpha}_0$	$\hat{\alpha}_1$	SE	Adj. R^2	DW	t-test $\hat{\alpha}_1 = 1$	F-test $\hat{\alpha}_1 = 1$
				Unrestricted			
August 1986–January 1990	−0.75 (−4.45)*	1.38 (5.31)*	0.79	0.59	1.61	1.47	—
August 1986–December 1987	−0.40 (−0.72)	1.55 (2.31)**	0.79	0.51	1.35	0.82	—
July 1987–January 1990	−0.74 (−4.25)*	1.89 (6.05)*	0.79	0.66	1.77	2.86*	—
Seasonally adjusted August 1986–January 1990	−0.77 (−4.02)*	1.29 (3.69)**	0.78	0.59	1.67	0.84	—
				Restricted			
August 1986–January 1990	−0.92 (−7.23)*	1.00	0.82	0.31	1.63	—	3.89***
August 1986–December 1987	−0.83 (−4.55)*	1.00	0.75	0.43	1.26	—	2.89***
July 1987–January 1990	−1.02 (−6.53)*	1.00	0.87	0.34	1.62	—	6.62***
Seasonally adjusted August 1986–January 1990	−0.91 (−7.53)*	1.00	0.78	0.38	1.71	—	0.41***

Note: Numbers in brackets are t-statistics for the null hypothesis that the corresponding coefficient is not significantly different from zero. One asterisk denotes significance at the 1 percent level, two asterisks denote significance at the 5 percent level, and three asterisks denote that the restriction in question cannot be rejected at a level of 1 percent significance. All regressions also include first- and second-order moving average components. The residuals have been identified to be white-noise processes using the method of Box and Jenkins, except for the period August 1986 to December 1987, when the residuals exhibit second-order serial autocorrelation.

Figure 7. *Uncovered Inflation Change*
(Actual and predicted values)



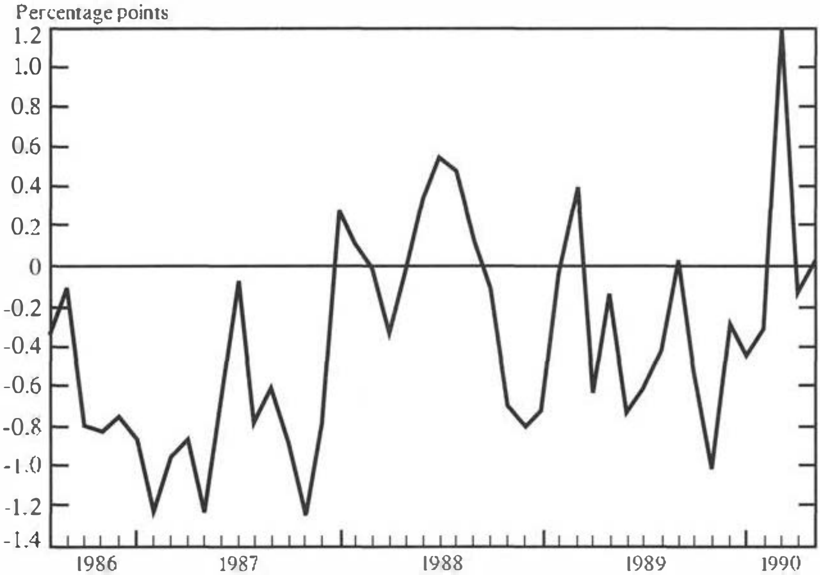
inflation. Policymakers have ready access to this interest spread and can use it as an indicator to assess the stance of monetary policy.[22]

Consider the interest spread between the quarterly UF premium and the monthly nominal interest rate depicted in Figure 8. According to the model, a declining spread is an indicator that monetary policy is recognized as tight, or that uncovered inflation is expected to fall, whereas an increasing spread indicates the opposite.[23] During the first and third quarters of 1989 the interest spread widened, and these two periods were also the two peaks of the uncovered acceleration of inflation during that year (see Figure 7). In the first quarter of 1990, the interest spread declined at first, as the Central Bank announced the introduction of an

[22] Note that the spread of interest rates is a useful indicator of the stance of monetary policy, but is not an instrument nor a target. The size of the spread per se does not indicate the extent of the adjustments that are necessary in instruments and targets, it only informs the authorities of the market's perception regarding monetary policy stance.

[23] The issue here is not how tight the Central Bank designs monetary policy, but how individuals in financial markets perceive it. Thus, it does not suffice that the authorities design a theoretically sound anti-inflationary policy; the reputation and credibility they command in financial markets also play a rucial role.

Figure 8. *Interest Rate Spread on Time Deposits, August 1986 to May 1990*
(90-day indexed minus 30-day nonindexed deposits)



adjustment program, but then it widened to reach 1.2 percent in March (4.9 percent annually). According to the model, the increase in the spread indicated that individuals viewed the adjustment policies either as transitory or insufficient to halt the increasing trend of inflation, possibly reflecting expectations of a fiscal expansion or large inflows of foreign capital attracted by a large favorable differential between domestic and foreign interest rates.

The data also support the hypothesis that ex ante real interest rates fluctuate around time-invariant means. With the exception of the estimate for the period August 1986 to December 1987, in which the estimate of the constant term is not statistically different from zero, the regressions show that the average 90-day ex ante real rate is more than ¾ of 1 percentage point higher than the average of the similar 30-day rate. As discussed previously, this evidence casts serious doubts on the use of the UF premium as an indicator of the relevant real interest rate because it clearly does not fluctuate around a time-invariant mean.

Equations (6)–(8) also provide an alternative test that can be used to substantiate further the claim that the UF premium is not equivalent to the ex ante real interest rate. These equations imply that, for both

rates to be statistically equivalent, the imperfection of the indexation mechanism should not be statistically significant. Thus, estimating equation (8) with a dependent variable that is just the negative of the 30-day-ahead inflation rate should produce similar results as those listed in Table 1. Estimation of this modified equation shows that this hypothesis is rejected. The coefficient $\hat{\alpha}_1$ is significant, but at a value of -0.42, the constant term rises to -1.6, and the first-order autocorrelation coefficient is higher than unity, which questions the stationarity of the residuals.

The UF premium is not the relevant ex ante real interest rate because, as explained earlier, in calculating the latter agents consider the discrepancy between the inflation of the month before they enter into a credit contract and their expectations of inflation in the last month before it expires. When the inflation rate is stable, this discrepancy is minimal and the UF premium reflects the relevant ex ante real rate, but when inflation fluctuates, this is no longer a correct approximation. The fluctuations affecting monthly inflation are due in part to seasonal or random factors, but they also reflect the underlying trend of the inflation rate. What the tests show is that, for Chile, the 1-month lag in financial indexation and the volatility of the month-to-month inflation rate cause discrepancies between the UF premium and the ex ante real interest rate that cannot be accepted as accidental. Thus, the results support the view that the spread between indexed and nonindexed interest rates is a better indicator of the stance of monetary policy than the UF premium itself.

IV. Concluding Remarks

This paper analyzes the Chilean mechanism of financial indexation and explores some aspects of its influence on the operation of the financial system. The results of a number of empirical tests support the view that indexation facilitates financial intermediation under inflationary conditions and does not automatically produce higher interest rates than a system free of indexation. In particular, the econometric evidence suggests that there is efficient arbitrage in Chilean financial markets and that market participants tend not to make systematic mistakes in forecasting future interest rates or future inflation. The evidence also supports the existence of the Fisher effect. The spread between indexed and nonindexed interest rates is found to convey significant information regarding future changes in inflation, and the average levels of ex ante real interest rates are estimated to be approximately constant over time.

The findings of this paper have four operational implications:

• *Efficient arbitrage is present in Chilean financial markets.* An increase in the effective rate of return on a 90-day indexed deposit is reflected in almost a one-to-one increase in the effective return on 3-month investments in 30-day nonindexed deposits.

• *Inflation changes are consistent with Fisherian transmission and rational expectations.* The difference between the 90-day UF premium and the 30-day nominal interest rate is a precise indicator of the expectations of economic agents regarding future changes in inflation (adjusted to account for the protection that imperfect indexation gives to 90-day deposits).

• *Indexation is imperfect.* The 1-month lag under which indexation is currently undertaken implies that the 90-day UF premium should not be regarded as the real interest rate relevant for economic decisions or for the purposes of designing monetary policy. The spread between indexed and nonindexed interest rates is a better indicator of monetary conditions.

• *Indexation eliminates the inflation risk from nonindexed financial contracts.* Because the data suggest that Fisherian transmission takes place in an environment where expectations are not systematically erroneous and efficient arbitrage exists, financial indexation minimizes the inflation risk (which is not detected in 30-day nonindexed deposits) and does not systematically force a full adjustment of nominal interest rates to inflation that would not have occurred otherwise.

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