Asset Prices in Chile: Facts and Fads

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Chile enjoyed an unprecedented period of rapid and sustained growth in the 1986-98 period. Although average growth exceeded 7.1 percent a year, significant cyclical fluctuations also marked the period, both across sectors and across time. Growth in the tradable sectors exceeded 6.7 percent annually from 1986 to 1992 but declined to 5.7 percent a year during the following five years. In contrast, the nontradable sectors experienced comparable growth rates in both subperiods (7.5 percent yearly). The evolution of key relative prices, such as the real exchange rate and the real interest rate, has shown distinctively different patterns. Several papers have studied the key role played by these relative prices in inducing the massive resource reallocation between sectors observed in Chile and in providing adequate incentives to exports and growth (Morandé and Vergara, 1997).

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Although these relative prices are an important component of the growth process, other prices contain information that is useful for understanding different aspects of the development of the Chilean economy. In particular, the important role of asset prices in signaling investment opportunities has not been studied. Changes in asset prices should reflect investment decisions made by economic agents in response to the evolution of underlying determinants (the fundamentals), such as productivity gains, the business environment, access to foreign saving, deepening of domestic financial markets, and the strengthening of the growth process itself. One interesting aspect of asset prices is that, by their nature, they reflect the expectations of economic agents about the future. They are also extremely flexible and adjust very rapidly to changing expectations. Consequently, the evolution of asset prices—and the reallocation of financial capital they induce—should reflect changes in wealth and permanent income more accurately and rapidly than prices in other markets, where output and demand often react sluggishly to changes in economic conditions.

The links between macroeconomic development and asset prices, however, are not necessarily clear cut, from either a theoretical or an empirical point of view. In general, abrupt changes in real asset prices are associated with economic crises or, at least, financial turmoil. Nevertheless, real asset prices sometimes exhibit episodes of overvaluation (called fads or bubbles), in which prices deviate from what can be attributed to the economic fundamentals. Historical evidence in many countries shows that whenever a bubble develops, its eventual burst causes extensive damage to both asset markets and the rest of the economy (Kindleberger, 1989). Occasionally, however, bubbles are only apparent, and the change in asset prices that is perceived as abnormal in fact corresponds more to changes in the fundamentals than to speculative market psychology. Consequently, observers find a role for these prices as leading indicators of changes in the fundamentals themselves. If prices of real assets are perceived to be misaligned, it can be an indication, for example, that the real exchange rate or real interest rates are misaligned.

In this paper we study real asset prices in Chile to measure to what extent their changing patterns correspond to varying market fundamentals or policy variables and to determine whether there have been episodes of bubbles. Of particular interest are the role of fiscal and monetary policies, the implications of changes in access to foreign savings, the impact of domestic credit constraints, and the role of pension funds in inflating or deflating asset prices. We concentrate on the prices of stocks (equity), agricultural land, and real estate, which we deem representative of the evolution of aggregate assets in the economy for the 1978-98 period. The analysis of the links between asset prices and their fundamentals has been largely ignored by economists in Chile, with the exception of Meller and Solimano (1983) and Morandé (1992). Whereas the former paper is limited to a descriptive analysis of asset price indices in the 1979-83 period, the latter models the evolution of asset prices and their underlying determinants using time-series models.

Section 1 of this paper discusses a standard general-equilibrium model in which asset prices arise as the result of agents trying to maximize their utility intertemporally when income is subject to stochastic disturbances. Asset accumulation is used to smooth consumption; hence asset prices reflect not only an intertemporal substitution effect, but also a forecast of future wealth. Consequently, maximizing agents will arbitrage between the short and the long run, and among assets. These arbitrage mechanisms, based on expected future economic activity, provide the basis for a long-run relationship between different asset prices. Nevertheless, short-run dynamics usually reflect deviations from the long-run condition described above. Section 1 also presents the main economic explanations for such deviations.

Consequently, addressing this behavior from an empirical point of view requires both a long-run and a short-run model. Section 2 provides an econometric approximation to such model, based on the literature on nonstationary time series. The estimated model successfully parameterizes the available data, allowing us to control for changes in fundamentals, weakly exogenous variables (such as international commodity prices), idiosyncratic variables (that is, those pertaining to a specific asset market), and policy variables.

Finally, we use the estimated model to study the dynamics of asset prices and simulate counterfactual scenarios, which we use to discuss the role that monetary and fiscal policies may have played throughout the 1990s in affecting asset returns. Arguably, the proposed counterfactual scenarios are simplistic. Our aim, however, is not to suggest that these are preferred policies, but to discuss the role that changes in such policies have had in affecting asset prices.

1. Determinants of Asset Prices

The literature on asset price determination is vast and diverse. Financial theories range from efficient-market models (Fama, 1970; Fama and French, 1988) to purely statistical event studies (Campbell, Lo, and McKinlay, 1997). Popular empirical models in the finance literature are the capital asset pricing model (CAPM) of Sharpe (1964) and Lintner (1965) and the arbitrage price theory (APT) of Ross (1976).

Economic theories, on the other hand, are based on the microeconomic behavior of agents who optimize dynamically within the context of an equilibrium model (Lucas, 1978). This type of model places a number of restrictions on the APT model and the CAPM. In particular, asset prices ought to reflect the valuation by market participants of the discounted stream of earnings derived from holding such assets. These earnings depend on a set of "fundamentals," which describe the functioning of the economy.

As a benchmark for the analysis, we develop a dynamic, stochastic general-equilibrium setting in order to characterize the long-run equilibrium restrictions for the relationship among asset prices. Later we discuss some of the main economic explanations for the existence of short-run deviations from the long-run equilibrium conditions. In this model, assets are used to smooth consumption, so that asset prices should reflect both intertemporal and intratemporal substitution effects. On the basis of their forecasts of future wealth, agents will arbitrage among assets to equalize their short- and long-run returns. This arbitrage mechanism, based on expected future economic activity, provides the basis for a long-run relationship between asset prices.

1.1 The Standard Model of Asset Prices

Consider a world in which a representative agent holds three assets—real estate, land, and equities—as a way to smooth consumption when income is subject to stochastic disturbances.[1] From the consumer´s optimization problem we derive the following Euler conditions:

1. Throughout this paper, "real estate" refers to owner-occupied housing and the land it occupies, whereas "land" refers only to agricultural land. "Housing" is used later in the paper to denote the rental price of housing.

$$1 = (1+i)\beta\,\mathrm{E}_t\left[\frac{u'(c_t+1)}{u'(c_t)}\right],$$

$$P_t^s = \beta\,\mathrm{E}_t\left[\frac{u'(c_t+1)}{u'(c_t)}\left(P_{t+1}^s + d_{t+1}\right)\right],$$

$$(1)$$

$$P_t^r = \beta\,\mathrm{E}_t\left[\frac{u'(c_t+1)}{u'(c_t)}P_{t+1}^r\right],$$

$$P_t^l = \beta\,\mathrm{E}_t\left[\frac{u'(c_t+1)}{u'(c_t)}P_{t+1}^l\right],$$

where $\beta = 1/(1+\rho)$ and ρ is the subjective rate of time preference, $u'(c_t)$ is the marginal consumption of the unique good, $\mathrm{E}_t\{\bullet\}$ denotes mathematical expectations conditional on the information available to the decisionmaker up to and including date t, and p_t^j is the real price in terms of the consumption good of an asset j (where the index j represents real estate r, land l, or equities s). The price of a share is quoted after dividends have been paid (the ex-dividend price). Variable d_t is the real output at time t given to holders of the shares, whose number is, for simplicity, normalized to one. Output is a random variable and generates the uncertainty in the model. Finally, assume that there exists a risk-free asset with a constant and exogenous return i. These Euler conditions show that the intertemporal marginal rate of substitution in consumption for the representative consumer is used to price payoffs on all securities traded in this economy.

One implication of these equilibrium conditions is that markets are efficient, in the sense that agents will arbitrage away any deviation of prices from equilibrium. Arbitrage will dissipate any economic rent among assets and over time. This efficient market behavior, however, depends largely on the availability of information. As already stated, the previous setting presumes that uncertainty arises only from shocks to output. In the actual economy, however, information is not a public good. If information is costly, agents will find it optimal to acquire less than full information. Hence market efficiency is limited to arbitrage conditions derived under restricted information. Moreover, if some information is available to insiders and not traded in any market, the efficiency hypothesis is stronger, as it requires arbitrage even among noise traders.

1.2 Departures from the Standard Asset Pricing Model

The efficient-market hypothesis has provided the basis for extensive research on asset prices. The empirical evidence, however, shows that economic and financial theories of asset price formation cannot satisfactorily account for their evolution. In particular, several authors have documented important short-run deviations of asset prices with regards to the fundamentals; these deviations are popularly dubbed "bubbles" (Blanchard and Watson, 1982; Campbell and Shiller, 1987). Econometricians have also found that returns (that is, asset price changes) tend to exceed what rational expectations models would have predicted for any reasonable degree of risk aversion. This is the equity premium puzzle of Mehra and Prescott (1985) and Kocherlakota (1996).

The existence of bubbles can be characterized by studying Euler condition (1), for example, for p_t^s . This stochastic difference equation admits a class of homogeneous solutions of the form

$$P_t^s = \sum_{k=t+1}^{\infty} \mathrm{E}_t \left[\left(\prod_{j=t}^{k} x_{j+1} \right) d_{t+1} \right] + B_t = P_t^* + B_t \ , \tag{2}$$

where P_t^* is the fundamental price of the stock and B_t is the bubble. The fundamental price is obtained assuming

$$\lim_{k \to \infty} \mathrm{E}_t \left[\left(\prod_{j=t}^{k} x_{j+1} \right) p_{t+1}^s \right] = 0 \ . \tag{3}$$

In other words, we preclude the existence of bubbles by imposing the condition that the discounted value of the infinitely distant future prices be zero. A bubble exists if the exchange value of an asset exceeds its fundamental price, which is the present value of the services or dividends it yields over its economic life. This condition, referred to as the transversality condition, follows from the exclusion of Ponzi schemes and allows one to obtain a unique solution to equation (2).

Bubbles occur when the valuation placed on a specific asset becomes self-fulfilling. In this case, price behavior follows a self-rein—forcing process not related to any valuation based on the fundamentals, and the market price of the asset includes some extra amount over the equilibrium value. When this extra amount is positive, investors believe they can sell the asset for more than they paid for it. Famous past episodes of such bubbles are the Dutch tulip bulb mania in the seventeenth century, the Mississippi Bubble in Paris in the eighteenth century, and, more recently, the 1987 stock market boom and crash. The bubble ends—and the crisis starts—when some investors decide to exit the market and the premium placed on the assets rapidly disappears.

How does the analytical literature deal with bubbles? One view is that bubbles can be expected to occur in models where there is continuous entry of new, richer agents who buy the existing assets with their savings (Obstfeld and Rogoff, 1996). Alternatively, bubbles may arise from errors in the definition of the fundamentals. Flood and Garber (1980) and Obstfeld and Rogoff (1986), among others, show that, in rational expectations models with incomplete information (for example, omitted variables), prices can deviate substantially from the discounted future earnings stream of the asset and its future terminal value. A different view is represented by the noise trader approach, where investors, being not fully rational, take actions and hold beliefs subject to systematic biases separated from the fundamentals (Shleifer and Summers, 1990). An important question in this context is whether asset prices are a reliable measure of the fundamentals and whether bubbles—inferred from them as the residuals—adequately capture misalignments with respect to the equilibrium value.

The recent turmoil in East Asian stock markets prompted a number of authors to focus on the connection among asset prices, credit constraints, and collateral when there are financial market distortions. Ito (1995) studied the co-movements of land and equity prices in Japan and found that positive credit shocks lead to loan expansions, which, in turn, prompt higher demand for fixed investment and land and a subsequent increase in land and equity prices. With the increase in stock prices, more funds are raised, and as a result, the price of land increases again, raising the value of collateral and allowing further increases in borrowing and investment. This multiplying process is due to lending constraints that arise either from incomplete financial markets (Kiyotaki and Moore, 1997) or from transactions costs (Stein, 1995). At some point, speculative activities start and prices begin to increase with expectations of further increases.

Finally, several recent papers stress the importance of information and incentives within the context of an unregulated financial sector, by relating bubbles to an agency problem (Allen and Gale, 1998), or by considering the importance of moral hazard by looking at implicit government guarantees (Krugman, 1998) or limited collateral (Edison, Pongsack, and Miller, 1998).

Empirical tests for bubbles, which were first employed before the theoretical literature was developed, focused initially on testing the excess volatility of asset prices. Motivated by the observation that fluctuations in the U.S. stock market seem too large to be explained by the fundamentals only, Shiller (1981) and Leroy and Porter (1981) provided statistical evidence that asset prices fluctuate markedly even when the fundamentals fluctuate little. Fama and French (1988), among others, provided evidence that returns are not only more volatile than implied by the fundamentals, but also predictable to a significant extent, contradicting the efficient-market hypothesis.

The predictability of asset returns is one of the most enduring questions in financial econometrics. An extensive literature concentrates on the exclusive use of past returns to forecast future returns, finding weak support for predictability. When other variables such as interest rates and dividends are introduced, however, the evidence becomes stronger. Campbell, Lo, and McKinlay (1997) show that, when changes in dividends are stationary and the convergence condition on the price series is imposed, there may be a stationary linear combination of prices and dividends (that is, they are cointegrated), even when dividends and prices are nonstationary. In this setting, the ability to predict the expected return on assets is not limited to the information from previous returns. This ability, furthermore, is determined by the microstructure of asset markets. In particular, if closing prices do not occur at the same time but are reported as if they did, a false impression of predictability can be created, and nonsynchronous trading effects can arise. Finally, the existence of persons with the monopoly right to post different prices in order to buy or sell shares translates into a cost associated with the bid-ask spread. Phillips and Smith (1980) showed that a significant fraction of abnormal returns can be eliminated when the previous spread is included. Blume and Stambaugh (1983) showed that the spread generates an upward bias in mean returns calculated with transactions prices. Finally, as Campbell, Lo, and McKinlay (1997) show, the movement of prices from the bid and ask prices can create spurious volatility and serial correlation in returns, even with no change in the economic value of the asset.

1.3 Asset Prices in Chile

Analysis of the evolution of asset prices in Chile is still incipient, largely because of the lack of appropriate time-series databases. An early paper by Meller and Solimano (1983) was the first attempt to test for bubbles in Chilean equity prices. They provided basic statistical evidence of large fluctuations in equity prices during the 1979-82 period and suggested that these were the result of the absence of adequate financial regulation. However, they did not develop a formal procedure to test such a presumption, nor did they provide any solid evidence of the existence of bubbles.

Morandé (1992) used quarterly data on key macroeconomic variables to investigate the extent to which the real prices of assets—land, real estate, and equities—were affected by structural reform and other macroeconomic policies in Chile during the late 1970s and early 1980s; he also investigated the timing of those effects. That paper stresses the dynamics of asset prices and the effects of exchange rate policies, tariffs, and capital inflows. Morandé found that there are important differences during the period under study with respect to the relevance of each policy for asset price formation. His results suggest that the fundamentals related to structural reform and key macroeconomic policies appear to be more important than portfolio considerations. The econometric analysis, however, was limited by the use of unrestricted vector autoregressive (VAR) procedures, which are able to characterize short-term dynamics but tend to omit long-term restrictions imposed by arbitrage conditions.

Recently, Budnevich and Langoni (1998) analyzed the behavior of real estate prices in Chile in the early 1990s. Using a panel data model, they established a long-run relationship between asset prices and their fundamentals. After analyzing real estate prices for six different counties in Santiago, they report a significant misalignment for some prices, thus casting doubts on the efficiency of the market.

In all of these studies of the Chilean case, government policies seem to play a key role in affecting asset prices. As noted, Meller and Solimano (1983) blamed the absence of financial regulation for the alleged emergence of a stock market bubble in the early 1980s. According to Morandé (1992), exchange rate policies and trade liberalization do affect asset prices in the short run. Likewise, Budnevich and Langoni (1998) assign a prominent role to monetary policies (via interest rates) and exchange rate policies. Certainly, macroeconomic policies played a key role in affecting the level of economic activity in Chile during the 1990s. In fact, the economy has twice experienced a significant contraction in aggregate demand as a result of restrictive monetary policies implemented with the goal of controlling inflationary pressures. To the extent that these policies relied on significant increases in interest rates, it is likely that they affected profitability in asset markets. Additionally, reforms in financial markets (such as pension funds) have been repeatedly blamed for the boom in asset prices observed throughout much of the last decade.

The Chilean experience constitutes an interesting case to study in order to improve our understanding of the evolution and determination of asset prices and their relationship with variables that proxy the evolution of fundamentals, such as profits, dividends, and rental prices for housing. Using a time-series cointegration framework, we attempt to capture not only the relationship between asset prices and their idiosyncratic determinants, but also their mutual influence in long-run equilibrium. This latter feature makes the econometric modeling consistent with the theoretical framework described above. In addition to policy variables (government expenditure and interest rates), the model considers external shocks, such as the terms of trade and capital inflows, which are deemed important in affecting firms' profitability and value.

2. Empirical Analysis

This paper focuses on the evolution of prices of three types of assets in Chile in the 1978-98 period: equities, land, and real estate. The period selected is the longest for which quarterly data are available (a detailed description of the data sources is contained in appendix A). Data on equities correspond to the General Share Index (IGPA) of the Santiago Stock Exchange, which is the only series spanning the entire period.[2] Data on land prices and real estate were constructed for this study according to a methodology developed by Morandé and Soto (1992). Land prices correspond to the hedonic price of a hectare of agricultural land, derived from the selling prices of farms located in the midsection of the country. To ensure homogeneity and representativeness, the data exclude forest lands and farms advertised as potentially divisible into small properties for residential use. Real estate prices correspond to the hedonic price of residential units offered within a selected set of counties in Santiago. The data exclude apartments and houses offered for commercial use.

2.1 An Overview of Asset Price Evolution in Chile

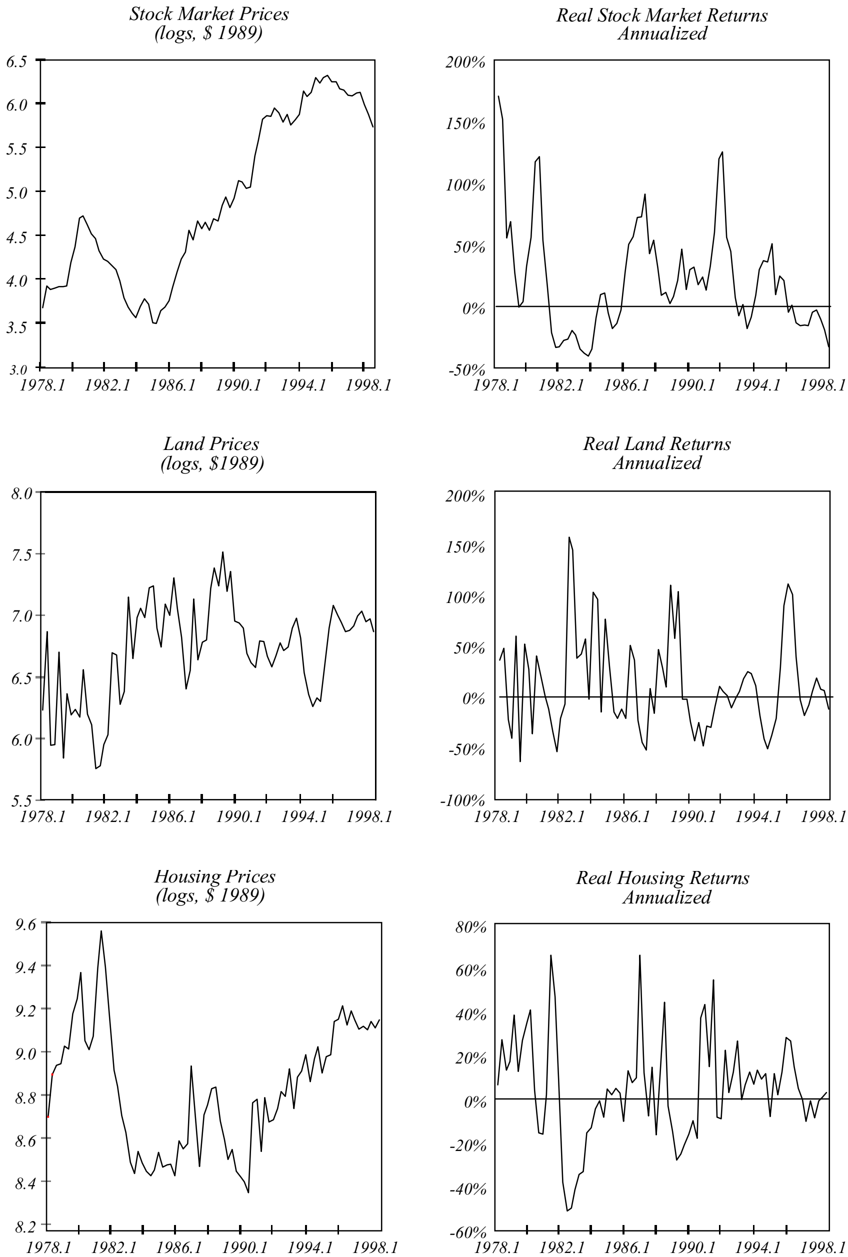
Two elements characterize the behavior of real asset prices in Chile, as depicted in figure 1. First, there has been a long-term upward trend in all prices, but changes in prices for different assets differ in magnitude and are clearly asynchronous. Upward trends in prices should not be surprising in a fast-growing economy, as rising profits are reflected in asset valuations. Second, asset prices are characterized by episodes of sudden and very significant fluctuations, both upward and downward, which are consistent with the onset and burst of bubbles. Examples of the latter may be the rise and fall in real estate prices in the late 1970s or that in land prices in the late 1980s. The evolution of equity prices reproduces to a significant extent the growth process of the Chilean economy, from the initial expansion in the late 1970s through the severe crisis of the early 1980s and the subsequent boom in the late 1980s and 1990s. Real estate prices display a similar response, but the cycles are exacerbated and exhibit a delay in adjustment when compared with the rest of the economy.

Returns on real assets are characterized by high volatility and, unlike in developed economies, the tendency to display large fluctuations over prolonged periods.[3] Persistence in returns is at odds with economic theory, which, as discussed in section 1, holds that intertemporal arbitrage should eliminate abnormal returns quite rapidly. The evidence in the Chilean case suggests otherwise. Equi-

2. There also exists a selective index (IPSA), which is deemed more representative of trading as it considers fewer but more regularly traded shares, but it is available only from 1981 onward. Both stock price indexes are, nevertheless, highly correlated (98 percent).

3. See Campbell, Lo, and McKinlay (1997).

Figure 1. Real Asset Prices and Returns



ties prices posted positive returns for twenty-nine quarters between 1985:4 and 1992:3, with only one occasion on which they were below 8 percent on an annualized basis. The annualized real return on real estate, on the other hand, remained negative for almost two years in the early 1980s (1982:1 to 1984:3), accumulating a total decline of more than 35 percent in real terms. The behavior of real returns largely reflects the link between asset prices and economic events and, in particular, macroeconomic policies, which were very active during the debt crisis of the early 1980s and after stabilization (the 1985-89 boom).

Nevertheless, macroeconomic variables cannot account for another characteristic of asset prices in Chile, namely, the fact that real returns do not evolve in similar, synchronized fashion as suggested by theory. The differential evolution of asset prices and returns can be easily assessed by looking at simple descriptive statistics such as those presented in table 1. It can be seen that the long-run average annual real return on equities is markedly higher than that on land or real estate (21.7 percent versus 12.6 percent and 4.7 percent, respectively). When the sample is split into the two decades, the result is unaffected. Naturally, differences in the riskiness of these assets should account for a fraction of the difference in returns, but it would be unjustified to attribute all differences to that factor. An additional explanation for this difference in returns is that, unlike equities, land and housing yield an implicit service in the form of rental prices. For example, real rental prices (on housing units of similar characteristics to those surveyed to derive real estate prices) averaged 14 percent in real terms in the sample, which would increase the return on real estate to 18.7 percent.

The return on equities has been quite high by international standards. The annual nominal return of the New York Stock Exchange index was 13.6 percent in the 1976-92 period (Harvey, 1994). However, the average return on real estate (4.7 percent) is slightly below the U.S. level for the 1975-95 period (5.3 percent), as reported by Taylor (1998).

A second interesting feature is that, although prices of all three types of assets increased in the period, the returns in the first and second decades present contrasting behaviors. Whereas equity and land returns declined markedly in the 1988-98 period, real estate returns increased significantly (from 3.3 percent in the first decade to 6.1 percent in the second). Although part of this phenomenon could reflect the adjustment induced by arbitrage, other elements also played an important role. For example, a significant fraction of the decline in land returns could be the result of the continuous appreciation of the Chilean peso since 1988 and the fact that land prices considered in this study were obtained from farms that produce tradable goods. Between 1988 and 1997, the real exchange rate appreciated by 30 percent.

A third element that characterizes asset returns is their high degree of volatility, as is apparent from the extreme values, standard deviations, and coefficients of variation presented in table 1. Annual returns as high as 170 percent and price drops of over 50 percent are impressive when one considers that changes in returns—as measured, for example, by the coefficient of variation—are usually a measure of risk or the risk premium (Engle, 1982).

Finally, the volatility of all asset prices fell significantly in the last decade when measured in absolute terms. A similar phenomenon is seen when volatility is normalized with respect to returns (using the coefficient of variation), except for land returns, whose volatility has clearly increased. This is consistent with the stability of the Chilean economy and the increased depth of financial markets.

Table 1. Statistical Properties of Annual Returns on Real Asset Prices

	Equities			*Land*			*Real estate*		
Measure[a]	*1978:1-1998:2*	*1978:1-1987:4*	*1988:1-1998:2*	*1978:1-1998:2*	*1978:1-1987:4*	*1988:1-1998:2*	*1978:1-1998:2*	*1978:1-1987:4*	*1988:1-1998:2*
Mean	21.7	25.5	17.7	12.6	16.4	8.8	4.7	3.3	6.1
Maximum	170.7	170.7	125.5	157.0	157.0	111.0	66.0	66.0	54.7
Minimum	–39.5	–39.5	–32.1	–64.1	–64.1	–51.1	–51.2	–51.2	–27.7
Standard deviation	43.5	52.1	32.6	46.9	51.1	42.4	23.0	26.7	18.6
Coefficient of variation	2.0	2.1	1.8	3.7	3.1	4.8	4.9	8.1	3.1

Source: Authors' calculations based on data sources in appendix A.
a. Mean, maximum, and minimum are in percentages per year.

2.2 Characterizing the Persistence of Shocks in Asset Prices

Assessing the persistence of shocks is important in econometrics because failure to account for the presence of permanent shocks (such as unit roots and other nonstationary processes) may lead to spurious regression results (Granger and Newbold, 1974). We use standard parametric tests to assess the presence of nonstationarity. Since parametric tests can potentially be flawed by model misspecification, appendix B reports the results of two additional procedures: variance-ratio tests (used to reveal slow processes of reversion to the mean) and endogenous-breaks unit-root tests (used to test for the presence of structural breaks). Those results complement and support the conclusions reached in this section.

Table 2 presents the results of augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit-root tests on real asset prices. As expected for this type of data, the series exhibit strong autocorrelation, which is usually indicative of the presence of unit roots. In fact, ADF and PP tests cannot reject the null hypothesis of a unit root for equity and real estate prices. For land prices, however, both tests reject the null hypothesis, although for the ADF test only marginally. With regards to the first (log) difference of the series (asset returns), stationarity is easily rejected.

The results of the nonparametric tests in appendix B confirm these results but also suggest that persistence in land prices is marked and that considering the series as an integrated process could be a reasonable characterization. Recursive parametric tests conclude that no series should be characterized as trend-stationary with structural level and/or trend breaks.

Table 2. Unit-Root Tests on Real Asset Prices[a]

	Level			*First difference*		
Price series	*Autocorrelation*	*ADF test*	*PP test*	*Autocorrelation*	*ADF test*	*PP test*
Equities	0.958	–1.57	–1.99	0.267	–5.17	–6.75
Land	0.694	–2.60	–3.91	–0.380	–8.62	–14.66
Real estate	0.880	–2.09	–2.04	–0.060	–9.62	–9.67

Source: Authors' calculations.
a. Results are based on quarterly data, 1978:1-1998:2. Critical values of the tests at the 10 percent and the 5 percent level are –2.58 and –2.90, respectively.

2.3 Characterizing Long-Term Relationships among Asset Prices

The results of stationarity tests sustain the notion that it is convenient to treat asset prices as integrated processes of order 1 without significant structural breaks. Once it is established that the variables behave as integrated processes of the same order, a natural question arises regarding the possibility of observing a long-run relationship among them. If there exists such a relationship (dubbed the cointegrating vector), it would imply that, although the series are subject to both permanent and transitory shocks, they present a stochastic common trend. The natural interpretation of this cointegrating vector would be as the long-run equilibrium that emerges from the arbitrage process between asset prices and returns.

There are several procedures for estimating cointegrating vectors.[4] Johansen's (1988) procedure, used in this paper, sets up a simultaneous-equation, error correction model, which corresponds to a quasi-VAR model and is parameterized by maximum likelihood methods. An advantage of this procedure over other methods is that it allows for direct testing of the number of cointegrating vectors. In addition, one can control for exogenous forcing variables in the estimation. However, two problems of Johansen's technique are its reliance on Gaussian processes for innovations and its tendency to obtain overparameterized results from VAR techniques.

Table 3 presents the results of applying Johansen's procedure to asset prices when controlling for government expenditure, capital inflows, terms-of-trade shocks, and other forcing variables described below. It can be seen that both informational criteria (Schwartz and Akaike) and the likelihood ratio test suggest the existence of one cointegrating vector. This result is important not only for our econometric purposes, but also for understanding the evolution of asset prices in Chile. It indicates that, despite the apparent randomness of their fluctuations, there is a tendency toward fulfilling arbitrage conditions.

Table 4 presents the results of the estimated cointegration model for asset prices, in particular the normalized representation for housing prices. Negative coefficients obtained for the companion asset

4. Two techniques for estimating cointegrating vectors are the two-stage estimation strategy suggested by Engle and Granger (1987) and the maximum likelihood procedure developed by Johansen (1988). Refinements of such techiques are the three-step algorithm of Park (1992) and the methods suggested by Phillips and Loretan (1991) and Boldin (1995).

Table 3. Johansen Cointegration Test on Asset Prices[a]

Rank or no. of cointegrating equations	*Data trend: None*		*Data trend: Linear*		*Data trend: Quadratic*
	No intercept, trend	*Intercept, no trend*	*Intercept, no trend*	*Intercept, trend*	*Intercept, trend*
	Akaike information criteria by model and rank				
0	−14.08431	−14.08431	−14.22726	−14.22726	−14.16908
1	−14.27120	−14.46594	−14.49107	−14.51170	−14.47655
2	−14.23298	−14.41194	−14.42528	−14.59933	−14.58360
3	−14.09388	−14.29109	−14.29109	−14.45054	−14.45054
	Schwartz criteria by model and rank				
0	−13.80192	−13.80192	−13.85074	−13.85074	−13.69844
1	−13.80056	−13.96392	−13.92630	−13.91556	−13.81765
2	−13.57408	−13.69029	−13.67225	−13.78355	−13.73644
3	−13.24673	−13.34980	−13.34980	−13.41513	−13.41513
Likelihood ratio test	Rank = 1	Rank = 1	Rank = 1	Rank = 1	Rank = 2

Source: Authors' calculations.
a. Based on quarterly data, 1978:1-1998:2.

Table 4. Normalized Cointegrating Vector for Real Estate Prices Using Johansen's Procedure[a]

	Constant	*Equity prices*	*Land prices*	*Linear trend*
Coefficient	12.886	−0.676	−0.647	0.069
Standard error	...	0.255	0.220	0.021

Source: Authors' calculations.
a. Data are for 1978:1-1998:2. All variables are in logarithms.

prices reflect the role of arbitrage in the economy's portfolio. The linear trend component is remarkably high (reaching 6.9 percent), similar to the average annual growth rate of the economy during the period (7.1 percent). This estimated cointegration equation is used below to estimate the short-term dynamic models. However, before discussing the results of the error correction models, it is important to note that the cointegration estimations are remarkably stable across the eighty-six-quarter period. Standard stability tests (Cusum and Cusum of squares) as well as the recursive estimation of residuals and parameters show little evidence of instability (see figures B3 and B4 in appendix B).

The estimated models closely track the long-term evolution of all asset prices, in particular those of equities and land. This is a remarkable result, considering the high volatility of asset prices, and suggests that arbitrage, as discussed in section 1, indeed takes place. As expected for long-term equilibrium values, short-term fluctuations show persistence and are not of a negligible magnitude. This, in turn, implies that error correction models are necessary to account for the short-term dynamics.

2.4 Error Correction Models

The estimation of the short-run models of asset prices follows the suggestions of Phillips and Loretan (1991) to include leads of the right-hand-side variables to capture potential feedback effects from asset prices. Drawing on the results on causality presented in appendix B (table B2), we include leads of those variables for which there was evidence of two-way causality: housing rental prices and capital inflows.

The results of the estimation are presented in table 5, and a simulation of the models is presented in figure 2. It can be seen that the error correction models are able to capture the dynamics of asset prices with accuracy. In fact, the pseudo R^2 of the regressions between the actual and simulated asset prices (derived from simulating the error correction models) is 99.1 percent for equity prices, 79.8 percent for land prices, and 94.9 percent for real estate prices.

These estimated error correction models also provide interesting results regarding the dynamics of asset prices and, in particular, the role of selected macroeconomic variables that influence the short-term evolution of returns. As is apparent from table 5, several variables have some impact on returns, and their dynamic relationships are rather complex. To organize the analysis, the variables are grouped into “exogenous” variables (such as the terms of trade), “policy” variables (such as public expenditure), and “idiosyncratic” variables (such as dividends). Although the first two potentially affect all asset returns, the latter are expected to have an impact only in their reference market. Nevertheless, to provide a consistent parameterization, the analysis proceeds from general to specific, including all variables and a reasonable number of lags in an initial model and proceeding to eliminate sequentially those that yield non-significant results, individually or in groups.

Policy variables have a characteristic effect on asset prices. Public expenditure has a negative and significant impact on equity prices, but a positive and very significant effect on land prices. The effect on real estate is of negligible magnitude. The negative effect on equi-

Table 5. Error Correction Models for Asset Returns[a]

Variable	*Equities*	*Land*	*Real estate*
Error correction term	−0.235	−0.757	−0.140
Change in land prices (1)	0.141	...	...
Change in equity prices (1)	0.343	0.447	−0.055
Change in equity prices (2)	...	...	0.115
Change in real estate price (4)	0.070	...	...
Public expenditure	−0.197	0.561	...
Public expenditure (1)	−0.222	0.294	...
Public expenditure (2)	...	0.358	...
Public expenditure (3)	...	0.311	...
Public expenditure (4)	...	...	−0.028
Real interest rate	−0.011	...	...
Real interest rate (1)	...	0.032	0.004
Real interest rate (2)	...	−0.021	...
Real interest rate (3)	...	...	−0.006
Capital inflows	...	−0.176	−0.077
Capital inflows (1)	0.020	...	...
Capital inflows (3)	...	−0.642	...
Capital inflows (4)	...	−0.835	...
Terms of trade (1)	...	−0.623	...
Terms of trade (3)	...	0.454	...
Terms of trade (4)	...	0.832	...
Credit	...	0.305	...
Credit (1)	−0.098	...	0.021
Credit (2)	0.017	...	...
Credit (4)	−0.007	...	−0.016
Dividends (1)	0.026	...	...
Dividends (3)	0.023	...	...
Housing rental price	...	...	−0.704
Housing rental price (1)	...	...	−0.779
Housing rental price (2)	...	...	−0.115
Shares in pension funds	0.457	...	...
Shares in pension funds (1)	0.351	...	...
Shares in pension funds (3)	...	−0.692	...
Shares in pension funds (4)	0.220	...	...
Summary statistic			
Pseudo R^{2} [b]	0.991	0.798	0.949

Source: Authors' calculations.

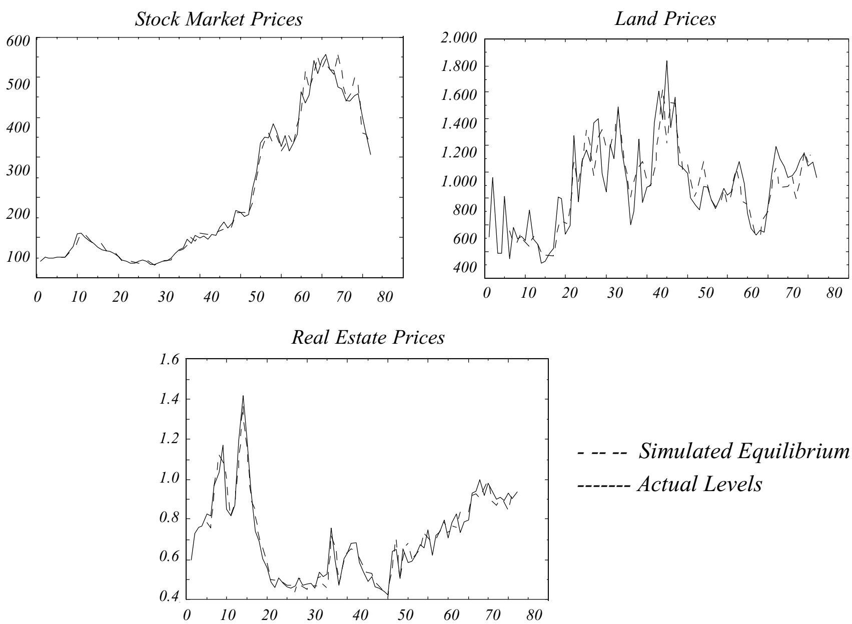
a. Numbers in parentheses indicate order of the lag.

b. Pseudo R^2 corresponds to the fit of a regression between the actual asset price and a simulated asset price obtained from the error correction models.

ties is consistent with Ricardian behavior on the part of investors, who anticipate that increased current public expenditure will require higher tax payments in the future. In contrast, landowners pay income taxes that largely depend on presumed rents (as opposed to actual profits), and they are less sensitive to income tax increases since presumed taxes are adjusted with significant lags. The size of the coefficients implies that an increase in public expenditure of 1 percent would induce a transitory decline in stock market returns of 0.4 percent and an increase in land prices of 1.6 percent. A change of this magnitude is not unusual in public expenditure in Chile; in the five years ending in 1998, for example, public spending fluctuated between 11.6 and 15.1 percent of GDP. On the other hand, a 1 percent increase in the real interest rate on risk-free government bonds reduces equity prices in the short run by 1.1 percent. A similar effect is seen for land prices, but there is an overshooting effect (of 3.2 percent), which is compensated the following quarter (a change of -2.1 percent). Finally, an equivalent change in the yield of government bonds also reduces real estate returns by 2 percent, but with a slightly longer lag.

We control these estimations for shocks to exogenous variables affecting the unobservable fundamentals underlying asset prices. Among these variables, a potential candidate is the real exchange rate, which could summarize effectively the competitiveness of traded

Figure 2. Asset Prices Simulated from Error Correction Models



Source: Authors' calculations.

goods for a small, open economy such as Chile. Nevertheless, the real exchange rate is in fact an endogenous variable, which depends on policies (such as the nominal exchange rate) and other fundamentals. Following Soto (1996), we use instead two key determinants of the real exchange rate: terms-of-trade shocks and capital inflows (a third variable, tariffs, proved nonsignificant). The effect of these variables on real estate returns is, as expected for a nontradable service, negligible. On the other hand, their impact on land prices, which reflect agricultural returns, is negative and significant. An improvement in the terms of trade or an increase in capital inflows leads to appreciation in the real exchange rate, thus reducing profits in the agricultural sector and therefore returns on land. Finally, the impact of both variables on stock market returns is, as expected for an index that combines both tradable and nontradable industries, of very small magnitude. The role of idiosyncratic variables is to some extent surprising, considering that these variables are intended to represent the effect of the flow of services derived from a stock of wealth. As is customarily found in the literature, dividends have a positive effect on equity prices both contemporaneously and after six months. One would have expected that rental prices for housing would behave in similar fashion; the results, however, show a zero overall effect, after some oscillatory behavior.

Two other variables have been proposed as capable of fueling changes in asset prices to the point of creating bubbles: private credit and purchases of equities by pension funds. The results tend to support the role of the latter but not the former. The overall impact of changes in the supply of credit on the stock market and on real estate prices is negligible. However, these changes do have some transitory effect on land prices. Purchases of shares by pension funds have had a positive effect on equity prices and, surprisingly, a negative effect on land prices. The latter result is additional evidence of a substitution effect between land and equities.

Finally, the error correction term shows that there are significant differences in the speed with which asset markets converge to equilibrium. To some extent, it is surprising to find that land prices converge to equilibrium twice as fast as the stock market (0.757 versus an implicit estimate of 0.357), whereas real estate prices converge very slowly (0.140). One would have expected stock markets to adjust faster than land markets, since transactions costs and entry fees are smaller.

2.5 Impulse-Response Functions

As is customary in time-series analysis, the error correction-cointegration model can be simulated to obtain the response of the dynamic system when subjected to a shock with impulse functions. A typical drawback of the methodology is that the ordering of shocks can have a significant impact on the decomposition of responses (Hamilton, 1994). Ordering should be based on the instant causality (that is, requiring intraperiod information), but in its absence econometricians usually simulate the models using different orders in alternation. In our case, however, we have the advantage of having monthly data available from which to obtain a reasonable (sample limited) approximation to intraquarter causality. Table 6 presents the results of Granger causality tests, which indicate that a sensible ordering would be to consider equity prices as precedent in time with respect to land and real estate prices, and real estate price shocks as preceding land price shocks.

Figure 3 presents impulse-response functions for the system. It is apparent that shocks have permanent effects, as expected in a cointegration model, and that these permanent effects tend to be rather idiosyncratic. Asset price shocks tend to be very persistent, with long-term responses above 50 percent of the initial impulse when considering "own" shocks (that is, shocks and responses within an asset category), which is consistent with the notion that changes in the order of the decomposition would not induce significant changes in the interpretation of the dynamics of the model. Cross-shocks are also of interest. Consistent with our intuition, stock market responses to shocks suggest that equities and land are substitute assets; positive land price shocks tend to lower equity prices as investors switch their portfolio structures toward more profitable alternatives. A similar effect is observed in land prices, which respond negatively to positive shocks in equity prices.

The response of real estate prices to shocks to other asset prices, on the other hand, suggests a relative insensitivity beyond an initial negative outcome. The negative response can easily be understood as a substitution effect. The long-term zero effect is interesting because, coupled with the response of equity and land prices, it provides an interesting asymmetry to the dynamics of asset markets. As is apparent, land prices show no long-run response to real estate shocks, but stock markets respond positively. This is consistent with the notion that higher real estate prices raise the value of the portfolios of firms (such as banks, which hold a large share of real estate), thus raising their valuation in the stock market. However, an increase in equity prices does not necessarily induce changes in real estate prices. Similar effects have recently been found by Ito (1995) for the Japanese case.

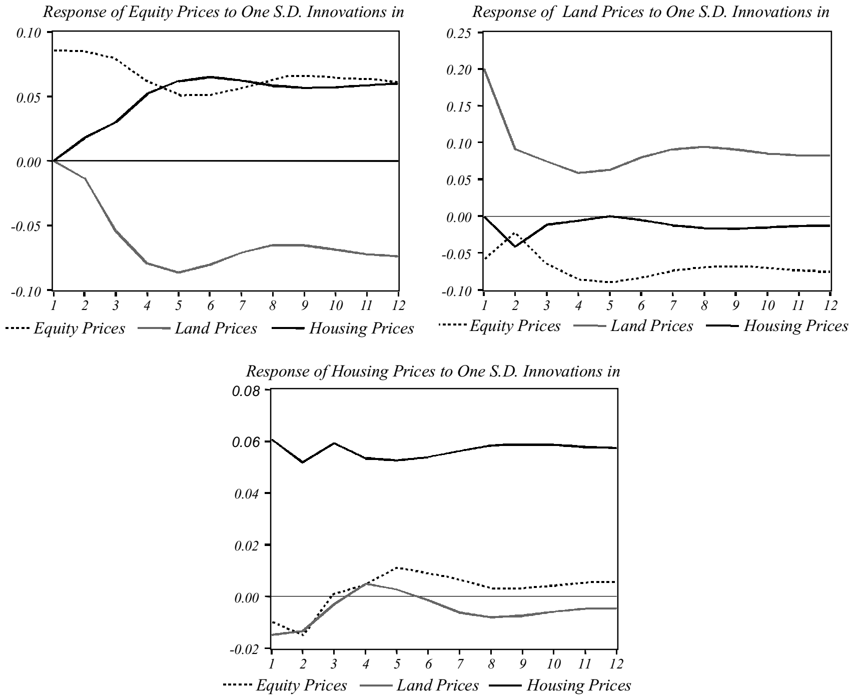
Table 6. Intraquarter Granger Causality Tests[a]

Direction of causality tested	*1 lag*	*2 lags*	*3 lags*
Land prices ⟶ real estate prices	0.95	0.59	1.29
Real estate prices ⟶ land prices	4.15**	2.04	0.84
Land prices ⟶ equity prices	0.67	0.65	0.72
Equity prices ⟶ land prices	5.40**	2.11	1.62
Real estate prices ⟶ equity prices	0.01	0.25	0.42
Equity prices ⟶ real estate prices	20.22**	7.08**	2.97**

Source: Authors' calculations.

a. The null hypothesis is that variable X does not Granger-cause variable Y. Data are monthly, 1975:1-1998:4. ** denotes significance at the 5 percent level.

Figure 3. Impulse-Response Functions



Source: Authors' calculations.

2.6 Non fundamental Movements and Bubbles in Asset Prices

As discussed in section 1, the definition of a bubble is to some extent arbitrary, since it is generally defined as a movement in asset prices not accounted for by a set of fundamentals, which are more or less arbitrarily defined by the researcher. Two alternative procedures can be derived from the above models.

First, one can use the cointegration models directly to determine the deviations of asset prices from equilibrium. The evidence for the Chilean case suggests the existence of episodes in which asset prices have drifted away from their long-term equilibrium level. As presented in table 7, equity prices have been as much as 25 percent above the level implied by the cointegrating relationship with other asset prices. Likewise, real estate prices have been "overestimated" by as much as 14 percent, in particular in the last five years. Both examples are consistent with frequently quoted opinions of market analysts that, as a result of different and often conflicting causes, bubbles have developed and burst.

As already discussed, however, a better representation of the dynamics of asset prices would be the error correction model presented in table 5. If economic agents are aware of the limitations to rapid adjustments in asset prices, it is natural to presume that they will internalize the existence of adjustment paths and act accordingly. Consequently, deviations from error correction forecasts and not cointegration models should be used as benchmarks for determining the presence of fads. Error correction models are able to properly accommodate both the long-run equilibrium relationship among

Table 7. Estimated Deviations of Asset Prices from "Equilibrium" Levels[a]

Percent

	Long-term equilibrium (full sample model)			*Full dynamic equilibrium (recursive model)*		
Period	*Equities*	*Land*	*Real estate*	*Equities*	*Land*	*Real estate*
1986:1-1989:4	0.7	10.5	–5.6	4.4	5.9	–0.8
1990:1-1993:4	24.7	–5.9	–17.5	2.9	1.0	1.2
1994:1-1998:2	–9.7	–2.0	13.9	5.0	3.3	3.0

Source: Authors' calculations.
a. Data are quarterly. A positive result implies that the actual value is above the estimated equilibrium level.

variables and their short-term fluctuations. The use of such a model estimated with the complete sample data, however, would yield unwarranted results, as it would amount to using more information than was available to agents at the time they made their portfolio decisions.

Consequently, a second procedure consists of estimating a recursive error correction model to estimate the "equilibrium" level of asset prices and determine the existence of bubbles. A recursive model uses only the information available to economic agents at each instant, and is thus more consistent with the situation faced by investors at the time decisions are made. In addition, recursive models are consistent with models in which agents learn the true nature of economic relationships and adjust to changes in those relationships (Sargent, 1993).

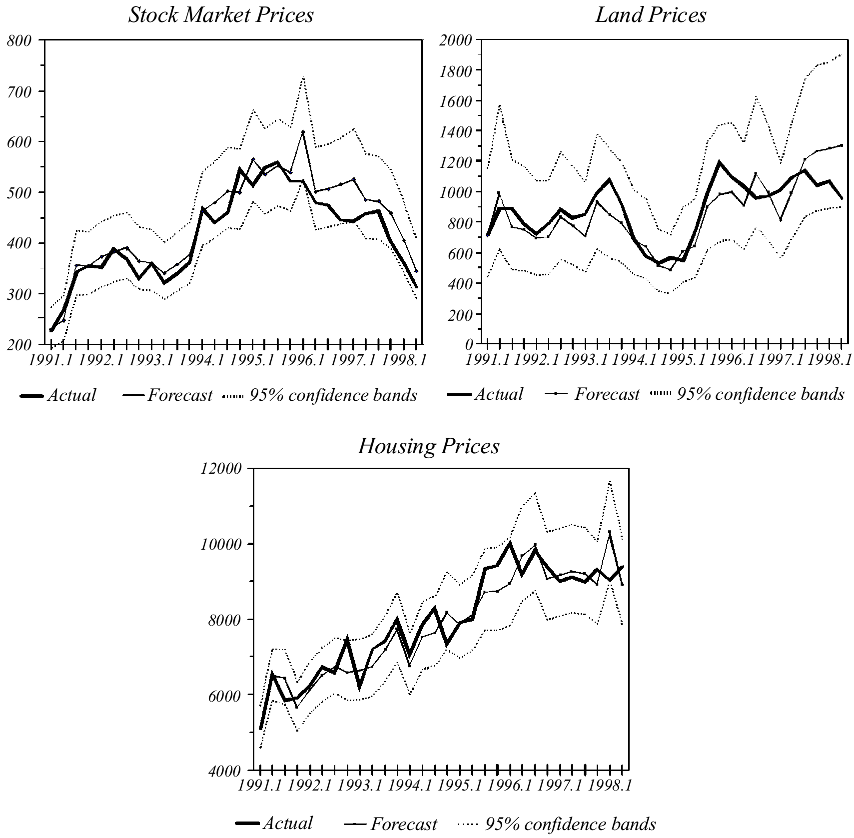
The recursive model is first estimated for the smallest sample size possible (say, T) and then used to compute the one-period-ahead forecast of asset prices and its standard error. Both the cointegration and the error correction models are estimated. In a second iteration, the sample size is increased to $T + 1$ and the model is reestimated, computing the forecast and its standard error. The procedure is repeated until all information is exhausted. Figure 4 presents the results of the recursive estimation of the error correction models and the forecast confidence.[5]

An interesting advantage of this procedure is that it allows us to obtain a 95 percent confidence interval for the forecast; thus it provides a statistical test for deviations of actual prices from "fundamental equilibrium" levels. It can be seen from figure 4 that actual asset prices do not cross the lines marking the 95 percent confidence interval; fluctuations are, consequently, comparable to random noise. The conclusion is that there is no statistical evidence of bubbles in asset prices in Chile during the period under study.

An alternative way of presenting these results is depicted in table 7. The recursive error correction models are used to determine equilibrium levels for each asset price and calculate the deviations of

5. The recursive estimation uses the Engle-Granger procedure, which is less data-demanding than Johansen's technique. This procedure is usually not optimal because of the presence of nuisance parameters (Campbell and Perron, 1991). For the data in this paper, however, two conditions ensure that the cointegration regression is asymptotically optimal: the right-hand-side variables are not Granger-caused by the left-hand-side variables (see appendix B, table B2), and the errors were not correlated.

Figure 4. Forecast Asset Prices Simulated Using Recursive Error Correction Models



Source: Authors' calculations.

actual prices from such a benchmark. It can be clearly seen that, on average, asset prices have not deviated substantially beyond 5 percent. In this regard, the deviations are again totally consistent with random noise.

2.7 Simulations of Alternative Macroeconomic Policies

The error correction models discussed above stress the role of macroeconomic policies in influencing asset prices. It is interesting to discuss what would have been the impact of alternative policy scenarios for public expenditure and interest rates. In addition, of special interest for the Chilean case is the analysis of a counterfactual scenario in which pension funds are forbidden to purchase equities.

Defining the alternative policies is quite difficult, as there is little guidance as to what would have been more appropriate (or even optimal) policies. Hence simple assumptions are made to model these alternative policies, which, nevertheless, help us understand the crucial role that policies played in a dynamic asset price context.

This section presents a series of dynamic simulations of the error correction models for the 1991-98 period. Simulations use the actual data for the 1978-90 period but simulate all asset prices thereafter. Considering that these are cointegrating models, they could accumulate large deviations from short-term equilibria.

Public Expenditure

As already mentioned, public expenditure is characterized by high volatility and a declining trend for most of the 1980s and early 1990s. This trend reversed significantly in the late 1990s, however. We considered that the following estimation would suffice for the purposes of obtaining an approximation to its data generating process:

$$Log\ Pub.Invest_t = -1.105 - \underset{(-6.27)}{} 0.0095t + 0.0007t^2 + 0.378\ LogPub.Invest_{t-1} \quad (4)$$

$$(-6.27)\ \ (-3.57)\qquad (2.39)\qquad (2.65)$$
$$R^2 = 0.631, \qquad DW = 2.00.$$

This trend-stationary model presents a reasonable fit for a simple structure and is consistent with an adaptive expectations model for public expenditure. Note that the presence of a quadratic trend term allows public expenditure to converge to a long-run level different from zero. The simulation of the model using the counterfactual level of public expenditure shows that, as a consequence of actual policies, stock market and real estate prices were depressed in the 1991-98 period by 5 percent on average. However, when analyzing the subperiods as is done in table 8, one notices important differences among them. In the 1991-94 period, equity prices were above what the counterfactual suggests by 5 percent, whereas real estate prices were depressed by 7 percent. In the 1995-98 period, in contrast, equity prices were markedly below what the counterfactual suggests (14 percent), because of excess public expenditure of 1.5 percent of GDP. On the other hand, public expenditure increased land prices by approximately 5 percent. Note that although public expenditure

Table 8. Estimated Effect of Alternative Macroeconomic Policies on Asset Prices[a]

Percent

	Public expenditure			*Monetary policy*			*Pension funds restricted*		
Period	*Equities*	*Land*	*Real estate*	*Equities*	*Land*	*Real estate*	*Equities*	*Land*	*Real estate*
1991:1-1994:4	–5.1	–0.4	7.1	–19.1	17.8	–11.1	–11.7	6.7	4.8
1995:1–1998:2	14.3	–5.0	4.9	0.9	5.3	–7.3	2.7	–1.1	–1.4
1991:1–1998:2	4.9	–2.8	5.8	–8.8	11.4	–9.2	–4.2	2.7	3.1

Source: Authors' calculations.
a. Data are quarterly. A positive number implies that asset prices were actually below those that would have prevailed under the counterfactual scenario.

does not have a significant short-term impact (as apparent in error correction models), the arbitrage condition of the cointegrating equations induces the portfolio shift.

Interest Rate Policies

Designing a counterfactual scenario for interest rates involves the use of a larger model, which would consider elements such as monetary policies, financial structures, and integration with international markets. Some of the variables involved may have unknown effects: for example, reserve requirements may lead to increased interest rates in the short run, but may induce more confidence in the economy and hence lower interest rates in the long run. An alternative methodology is to assume that arbitrage between domestic and international financial markets guarantees interest rate parity, once expected devaluations and country risk premiums are properly included. Since a measure for the latter is not available, we use as a proxy the premium paid on the dollar in the parallel exchange market. This parallel market is totally legal in Chile, so that the exchange rate does not include any premium for the risk of undertaking illegal activities. The counterfactual scenario then sets the domestic interest rate at the six-month London interbank offer rate (LIBOR) plus the exchange rate premium.

As expected, the results in table 8 show the key importance of the interest rate in influencing asset prices. When analyzing the impact on equity prices, it is clear that there are different effects in the two subperiods. The 1991-94 period is characterized by a much larger premium (1.1 percent on average) than the 1995-98 period

(0.25 percent), as the country consolidated its transition to democratic rule and progressively integrated with international markets. Although stock market prices were heavily affected by interest rate policies in the former period, the distortion (an overvaluation of 19 percent) has almost disappeared in the second (0.9 percent). A similar asymmetric behavior is observed in the other asset prices. It is interesting to note that high interest rates depressed real estate prices by almost 10 percent in the 1991-98 period.

Pension Fund Policies

In early 1990 the Chilean government allowed pension funds to acquire increasing volumes of equities as a means of diversifying their portfolios. Massive investments were made in the 1991-94 period, raising equity prices. Nevertheless, in the same period the value of firms whose shares are traded on the stock market was also growing at a fast pace, as the economy expanded at an annual rate of over 7 percent and diversified markedly. Consequently, equity prices were also expected to increase for this reason. The proposed counterfactual scenario simulates the model restricting pension funds to hold no equities, and calculates the effect of such a restriction on asset prices.

Table 8 suggests that the “excess” value of equities amounted to almost 12 percent on average in the 1991-94 period. This was the period in which the pension funds actively demanded equities in the market. In the 1995-98 period, in contrast, the pension funds did not have any significant effect on equity prices; in fact there was an undervaluation of around 2.7 percent on average. As a side effect of the purchase of equities by pension funds, asset substitution led to depressed values for land prices of as much as 6.7 percent on average in the early 1990s. Real estate markets, however, proved to be less affected by these effects on the stock market.

3. Final Remarks

For over a decade, Chile enjoyed rapid and sustained economic growth. As expected, asset prices rose, reflecting the increase in wealth brought about by such growth. Nevertheless, real asset prices also displayed high volatility, and there have been periods in which returns were systematically negative or positive. Casual inspection, then, suggests the presence of sustained deviations from the fundamental value of assets, that is, the presence of fads or bubbles. The fact that some of these fads coincided with periods in which important changes in macroeconomic policies were implemented suggests that policies may be able to affect asset prices, at least in the short to the medium run.

The empirical analysis in this paper tends to refute this notion of asset bubbles in the Chilean case, when a dynamic error correction framework is used to model asset returns. The presence of a cointegration equation provides an anchor for asset price deviations in the long run, whereas the dynamic model is able to capture a substantial portion of short-run fluctuations. Arguably, any deviation from the long-run forecast of asset prices could be considered a bubble; in such a case, there have been episodes in which asset prices were as much as 25 percent higher than what the fundamentals would call for (an example is that of equities in the 1990-93 period). If we allow for market friction or less than perfect information, as is likely in the Chilean economy, deviations of actual asset prices should be evaluated against the short-term model (which includes the long-run restrictions on asset price arbitrage). In such cases, fads have been nonexistent, as the deviations do not amount to more than 5 percent.

When considering the dynamics of asset prices, it is found that shocks tend to be very persistent, with long-term responses to own shocks above 50 percent of the initial impulse. This result is consistent with the notion of significant permanent components in asset price shocks and costly adjustments to long-run equilibrium. In addition, and consistent with intuition, stock market responses to shocks suggest that equities and land are substitute assets; positive land price shocks tend to lower equity prices as investors switch their portfolio structures toward more profitable alternatives. A similar effect is observed in land prices, which respond negatively to shocks to equity prices.

The response of equity and land prices to shocks provides an interesting asymmetry in the dynamics of asset markets. Land prices show no long-run response to real estate shocks, but stock markets respond positively. This is consistent with the notion that a rise in real estate prices induces a higher value of the portfolios of firms (such as banks, which hold a large share of real estate), thus raising their valuation in the stock market. However, an increase in equity prices does not necessarily induce changes in real estate prices.

An interesting feature of the model is that it allows for testing of the role of fiscal and monetary policies, as well as changes to regulations that, according to market observers, had a significant impact in fueling asset price bubbles. By simulating the model in a counterfactual scenario in which pension funds were forbidden to acquire equities, we determined that their impact has been sustained but reduced in magnitude. In the 1991-98 period, purchases of equities would have raised equity prices by an annual average of 4.2 percent.

Likewise, public expenditure has had a significant but asymmetric effect on assets. The estimated model finds a negative relationship between public expenditure and equity prices, but a positive link with regard to real estate prices. Consequently, the observed expansion in public expenditure during the 1990s led to a decline in equity prices on the order of 4.9 percent a year and an appreciation of about 6 percent in real estate prices. Moreover, the sudden increase in such expenditure during the 1995-98 period led to a drop in equity prices of 14.3 percent.

Monetary policies, on the other hand, have a significant effect on asset prices. When interest rates are allowed to fully reflect uncovered parities with international capital markets (that is, when restrictions on capital flows are lifted), we observe a reduction in equity prices of about 9 percent in the 1991-98 period and a higher predicted level of land prices of 11.4 percent.

Appendix A

Data Sources and Definitions

The macroconomic data were obtained mostly from the Central Bank of Chile; other series were obtained from specific sources discussed below or collected by the authors specially for this study. In particular, land prices, housing prices, and rental prices were estimated using Morandé and Soto's (1992) methodology. The definitions of the variables and their sources are described below:

– *Equity prices*: Indice General de Precios de Acciones, Santiago Stock Exchange.

– *Land prices*: Hedonic price of one hectare of agricultural land, for farms located within regions V through VIII. Farms smaller than 30 hectares or larger than 500 hectares are excluded, to eliminate land subject to subdivision for housing and land subject to forest exploitation, respectively.

– *Real estate prices*: Hedonic price of a house located in a homogeneous sector in the counties of Ñuñoa and Providencia, excluding houses devoted to commercial use and apartments.

– *Dividends*: Data from the Santiago Stock Exchange.

– *Exchange rates*: Bilateral rate of pesos to the U.S. dollar, from the Central Bank of Chile.

– *Capital flows*: Observed transactions in the form of foreign direct investment and long-term loans (long-run capital inflows), and portfolio investment and short-term lending (short-term inflows), all expressed as percentages of GDP using the official nominal exchange rate. Data from the Central Bank of Chile and Morandé (1992).

– *Public expenditure*: Expressed as a ratio to GDP, excluding investment undertaken by public enterprises and decentralized agencies. Data from the Central Bank of Chile.

– *Real interest rates*: Long-term yield on eight-year bonds, calculated ex post using nominal rates and observed consumer price inflation.

– *Rental price for housing*: Hedonic price of a house located in a homogeneous sector in the counties of Ñuñoa and Providencia, excluding houses devoted to commercial use and apartments. Data are consistent with real estate price data.

– *Pension fund purchases of equities*: Change in the value of equities held by pension funds. Data from the Superintendency of Pension Funds.

– *Credit*: Total credit by the financial sector (*colocaciones*). Data from the Central Bank of Chile.

– *Terms of trade*: Data from the Central Bank of Chile.

Appendix B

Data Properties and Estimated Model Results

B.1 Characterizing the Persistence of Shocks in Asset Prices and Fundamentals

Table B1 presents the results of augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests performed on all variables in the models. The first autocorrelation (ρ) is also presented. Tests optimized the number of lags to control for possible residual autocorrelation.

B.2 Nonparametric Tests of Nonstationarity in Asset Prices

Misspecification of the alternative hypothesis and persistence beyond the initial autocorrelations can lead to misjudgments regarding the existence of unit roots. This suggests complementing the standard tests with additional procedures as discussed below.[6] In addi-

Table B1. Unit-Root Tests[a]

	Level			*First difference*		
Variable		*ADF test*	*PP test*		*ADF test*	*PP test*
Equity prices	0.958	−1.57	−1.99	0.267	−5.17*	−6.75*
Land prices	0.694	−2.60	−3.91*	−0.380	−8.62*	−14.66*
Real estate prices	0.880	−2.09	−2.04	−0.060	−9.62*	−9.67*
Bond yield	0.848	−2.78	−2.86	−0.030	−9.49*	−9.21*
Fiscal expenditure[a]	0.707	−1.81	−3.59*	−0.38	−9.49*	−16.60*
Long–term capital inflows[b]	0.598	−3.09*	−4.43*	...	...	...
Real dividends	0.254	−2.26	−7.07*	...	...	...
Real loans	0.921	−3.55*	−3.32*	...	...	...
Terms of trade	0.913	−2.44	−2.41	0.221	−7.28*	−7.35*
Nominal exchange rate	0.966	−2.03	−2.60	0.560	−4.82*	−4.74*
Housing prices	0.791	−3.07*	−2.91*	...	...	...
Short–term capital inflows[b]	0.083	−7.80*	−7.96	...	...	...

Source: Authors' calculations.
a. Data are for the period 1978:1-1998:2. Critical values of the tests at the 10 percent and 5 percent significance levels were −2.58 and −2.90, respectively. * indicates rejection of the null hypothesis at the 5 percent level.
b. As a percentage of GDP.

6. Standard parametric procedures, such as those used above, are widely used by econometricians despite their lack of power in finite samples, when a root in the MA representation of the series is close to (but different from) one, or when the alternative hypothesis is a complex, nonstationary model other than a random walk, such as an ARIMA(1,1,1).

tion, in order to discuss asset price predictability, it is convenient to explore whether these prices follow a random walk.

A nonparametric alternative, suggested to cope with some of the limitations of parametric tests, is the variance-ratio (VR) test,[7] which has comparable power against stationary AR(1) alternatives in small samples but proves superior when the data are generated by more complex specifications (Chow and Denning, 1993). The test measures the importance of the nonstationary component of a series by comparing the variance of the permanent component to that of the innovations. Its main advantages are that it captures long-run reversions to the mean more accurately than do standard parametric techniques in small samples, and it is less sensitive to structural breaks (León and Soto, 1995).

The VR test exploits the fact that the variance of a random-walk process [$X_t = \mu + X_{t-1} + \varepsilon_t$, with $\varepsilon_t \sim N(0, \sigma_\varepsilon^2)$] increases linearly in the sampling interval, and that the magnitude of the random-walk component of a series can be inferred from the $1/q$ ratio of the variance of q-differences to the variance of the first difference of the series. For example, if the series follows a random walk, the variance of quarterly increments must be three times as large as that of the monthly increments; dividing by $1/q$ is a convenient normalization.

For empirical purposes, consider a sample of T observations of the process $\{X_t\}$. The VR can be obtained as:

$$V(q) = \frac{1}{q}\,\frac{\hat{\sigma}^2(q)}{\hat{\sigma}^2(1)} \quad ,$$

$$\hat{\sigma}^2(1) = (T-1)^{-1}\sum_{t=1}^{T}\left(X_t - X_{t-1} - \overline{X}\right)^2 \quad ,$$

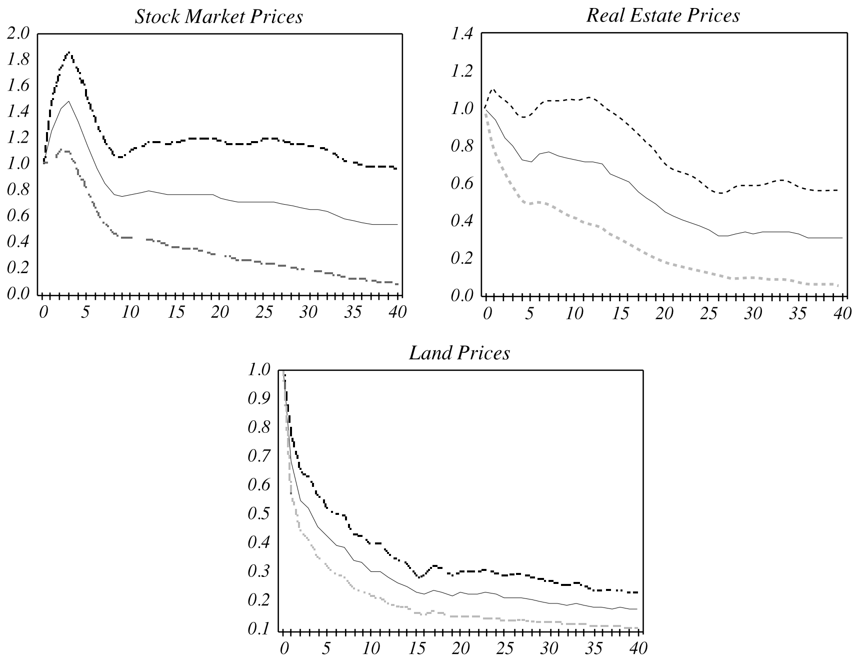
$$\hat{\sigma}^2(q) = \left[(T-q+1)\left(1-\frac{q}{t}\right)\right]^{-1}\sum_{t=q}^{T}\left(X_t - X_{t-q} - q\overline{X}\right) \quad ,$$

where $\hat{\sigma}^2(1)$ and $\hat{\sigma}^2(q)$ are the sample estimators of the variance of the first and the qth difference and X is the mean of the first differences of X_t.

7. This estimator of persistence was proposed by Cochrane (1988) and extended by Lo and McKinley (1989).

Figure B1 presents the variance ratios for each asset price, including a 95 percent interval of the null hypothesis of a random-walk component (represented by the dotted lines). As is apparent, asset prices do not follow random-walk processes, since all series exhibit important long-term mean reversion components and reject the null at 95 percent. Nevertheless, they also exhibit very strong persistence in the short run, which is indicative of nonstationary components. This type of evidence led Cochrane (1988) to suggest that, in the presence of strong autocorrelation and finite samples, it would be advisable to treat series as if they were integrated even if they revert to the mean in the very long run.

Figure B1. Variance Ratio Test for Asset Prices



Source: Authors' calculations.

B.3 Nonstationarity and Structural Breaks

Both parametric and variance ratio tests are sensitive to the presence of structural breaks (Hendry and Neale, 1991; León and Soto, 1995). As noted by Perron (1989), a stationary model with a level shift can easily be taken for a integrated process. In fact, Perron shows that several economic series usually regarded as containing a unit root (such as GDP and interest rates) actually could correspond to stationary models with shifts in the level, the trend, or both.

This observation is of some importance in the Chilean case, since a significant number of reforms were implemented during the period covered in this study. Although none of these reforms impinged directly on asset prices, they nevertheless affected fundamentals that guide asset price formation (examples are tax reforms, trade liberalization, and financial market deepening).

To assess the presence of structural breaks in asset prices, we use a test that endogenizes the dating of the breaks. This is done by estimating recursively a normalized version of Perron's (1989) additive-outlier model, starting from the beginning of the period and moving period by period toward the end of the sample (normalization is required to control for the changing size of the sample). Perron's (1989) test cannot be used directly, because it requires that the dating of the breaks be given by the econometrician with a priori information only (usually from visual inspection of the data). The test used in this paper allows the position of the break to be endogenously determined at the point at which the null hypothesis of a unit root with constant parameters is more easily rejected against the competing alternative TS representation with shifts in either the level or the trend of the series. The appropriate asymptotic distribution of the statistic, obtained by Zivot and Andrews (1992) using Monte Carlo simulations, provides critical values well below the full-sample DF critical values. When assessing the power of the test against stationary models, the recursive test has comparable power to that of the standard DF test in the absence of breaks.[8]

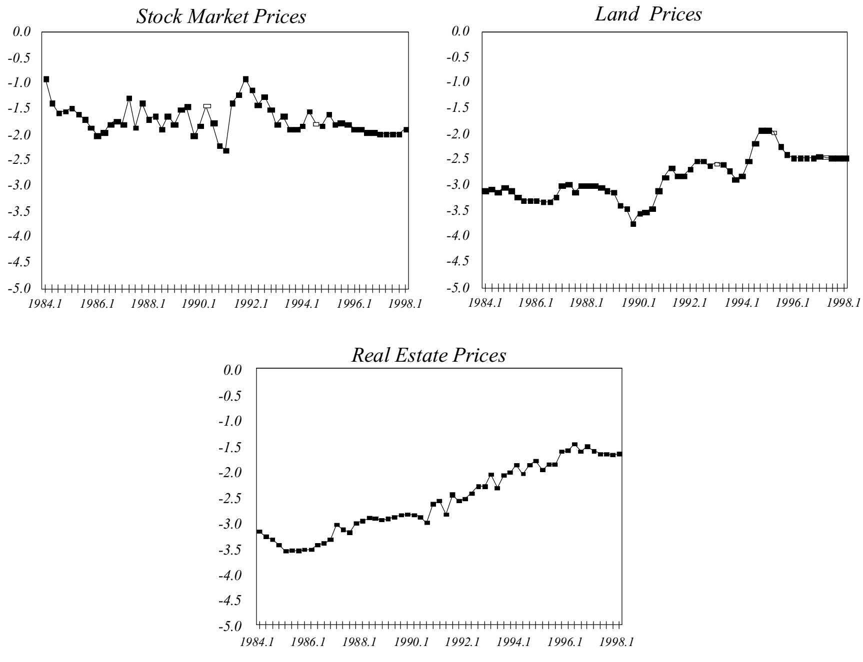
8. Its main limitation, however, is that it tests for only one break. It is, nevertheless, preferred over a similar test developed by Banerjee and others (1992), because the number of lags is endogenously determined at each recursion, avoiding potential misspecification problems.

The results, presented in figure B2, imply that it is not possible to reject the null hypothesis that asset prices are characterized as a difference-stationary process at a 95 percent level of confidence. The alternative hypothesis of trend-stationary processes with level and/or trend shifts is not supported by the data.

B.4 Granger Causality Tests

Table B2 presents Granger causality tests for the variables used in the empirical estimation of error correction models.

Figure B2. Zivot-Andrews Unit-Root Tests (Structural Breaks Allowed)



Source: Authors' calculations.

Table B2. Granger Causality Tests[a]

	Price of		
Variable	*Stocks*	*Land*	*Real estate*
Terms of trade	1.32	1.31	0.16
Public expenditure	0.05	0.27	0.18
Real interest rate	1.86	1.49	0.03
Credit	1.83	0.08	1.42
Capital inflows	1.27	0.74	2.42*
Dividends	1.00	0.17	0.78
Housing rental price	0.07	0.17	5.35*

a. The null hypothesis is that innovations in asset prices do not Granger-cause changes in the indicated variable.
* indicates significance at the 10 percent level.

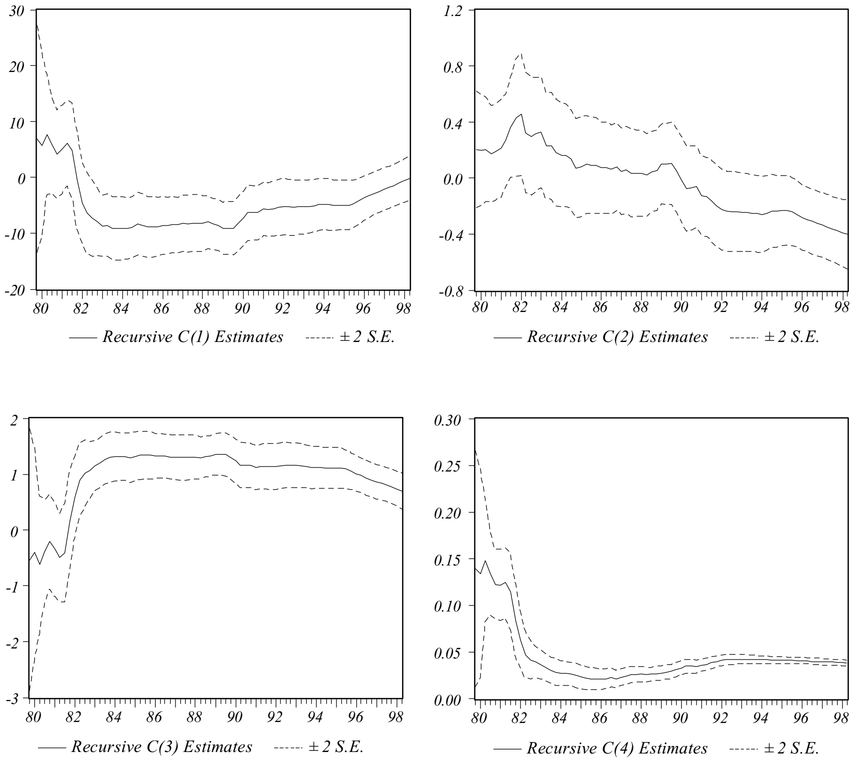
B.5 The Stability of the Estimated Model

Figure B3 displays estimated recursive coefficients of the long-run cointegration model used to assess the stability of the estimation. It can be seen that the estimated parameters are stable (recursively estimated coefficients for the error correction models are available upon request).

Figure B4 presents recursive estimates of the residuals of the models that show no signs of model mispecification.

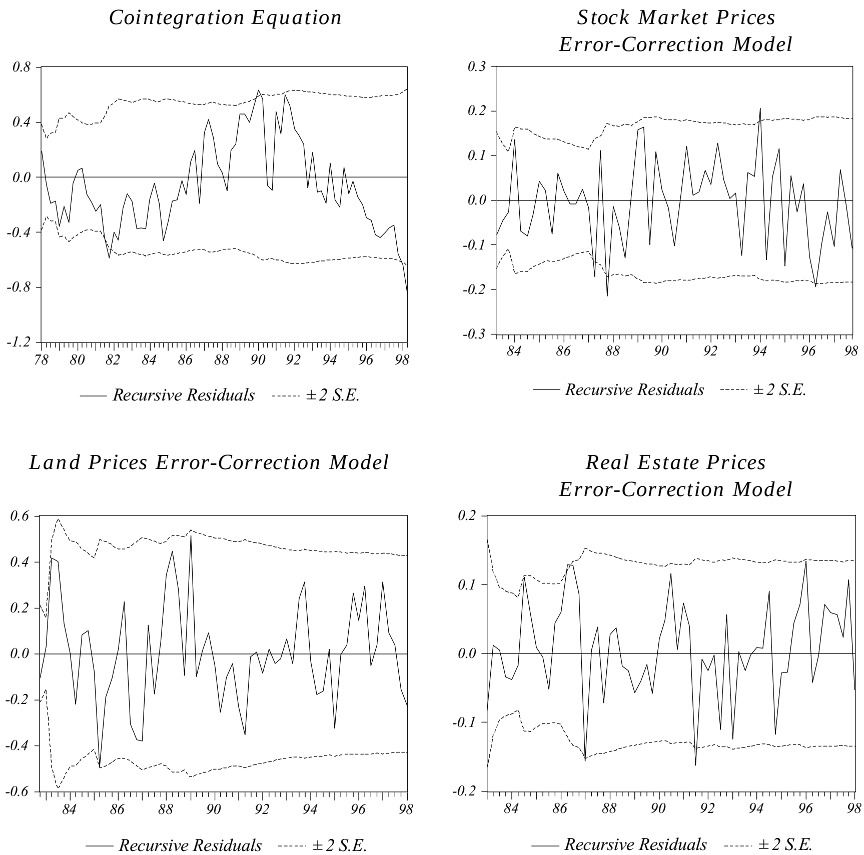
Figure B3. Recursive Estimation of Coefficients

Cointegration Equation



Source: Authors' calculations.

Figure B4. Recursive Residuals



Source: Authors' calculations.

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