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INFLATION COMPENSATION AND INFLATION EXPECTATIONS IN CHILE

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Resumen

Este trabajo estudia la relación entre la compensación inflacionaria y las expectativas de inflación en Chile. En primer lugar, usamos la metodología de valor presente descontado para descomponer la diferencia entre el retorno no anticipado de bonos nominales y reajustables en noticias acerca de inflación esperada y premios. En segundo lugar, usamos un modelo de valoración de activos de equilibrio general para estimar un premio por riesgo inflacionario que varía en el tiempo. Nuestros resultados muestran que movimientos en las expectativas de inflación explican alrededor de un 25% de los movimientos de los retornos relativos, indicando que los premios son una fuente importante de la variación de la compensación inflacionaria. También mostramos que el premio por riesgo inflacionario estimado varía a través del tiempo, pero que parece ser de tamaño despreciable, con media y volatilidad cercanas a cero.

Abstract

This paper studies the relationship between inflation compensation and inflation expectations in Chile. First, we use the present discounted value methodology to decompose the difference between the unanticipated return of nominal and inflation-linked bonds into news about expected inflation and premiums. Second, we use a general equilibrium asset-pricing model to estimate a time-varying inflation risk premium. Our results show that inflation-expectations movements account for about only 25% of the relative returns, indicating that premiums are a very important source of changes in inflation compensation. We also show that the estimated inflation risk premium is time-varying but seems to be of negligible size, with average size and volatility very close to zero.

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1 Introduction

Inflation expectations of market participants are of particular interest to central banks. Having an accurate measure of market inflation expectations helps the monetary authority assess its effectiveness in controlling inflation, as well as its credibility among market participants.

There is no single measure for evaluating inflation expectations. Rather, they are obtained using different methods, which typically provide a range of results. One commonly used measure of expected inflation is forecasts based on survey responses. However, surveys only cover a very small portion of the population, they are updated infrequently, and may not be completely reliable if respondents answer questions casually instead of giving their best efforts.

An alternative source of information about inflation expectations is available directly from financial-market data. In particular, the differential between the yields of nominal and inflation-linked bonds, known as inflation compensation, can be used to obtain information regarding expected inflation. The yield on a nominal bond must compensate its holder for the expected depreciation of the purchasing power of money during the life of the bond. In contrast, inflation-linked bond-holders demand no such compensation since the payoff of the bond is indexed to inflation. Hence, the difference between both yields should reveal inflation expected by market participants.

Inflation-compensation measures have the appeal of being based on investment decisions of a large number of agents who risk their own resources for such decisions. Furthermore, this information is available with very high frequency. The drawback is that the yield differential might contain additional information besides expected inflation. Specifically, the differential might also contain an inflation risk premium, since risk-averse investors dislike inflation uncertainty, and a liquidity premium, due to the liquidity difference between the two kinds of bonds. These premiums make extracting information about expected inflation from inflation compensation difficult.

Figure 1 depicts the evolution of inflation compensation for different maturities in Chile during the period 2002-2006.

[Insert figure 1]

There has been times when inflation compensation has coincided with the Central Bank's inflation target of 3%, while at other times there has been significant differences. Two questions arise from observing the figure. The first one refers to the level of inflation compensation. If we observe, for example, that in June 2005 the one-year inflation compensation equals 3.1%, does this mean that expected inflation is also 3.1% or is expected inflation anchored to the target and there are premiums of 10 basis points? The second questions regards the variation of inflation compensation. Between June and July 2005 inflation compensation increased by 20 basis points. Is this movement associated to a 20-basis-point increase in expected inflation or to a 20-basis-point increase in premiums? Identifying the level and volatility of the premium allows us to develop a much clearer view of how markets see expected inflation. If the premiums are different from zero, they will shift the level of inflation compensation from 'true' inflation expectations. But if they are relatively constant trough time, inflation compensation would still be a useful indicator since changes in inflation compensation can be attributed exclusively to changes in expected inflation.

In response to the questions described above, the literature has followed two strands of research. The first one is concerned in estimating the levels of the premiums. A first group of papers (Shen, 1998) estimates the premiums (inflation risk and liquidity) residually as the difference between inflation compensation and expected future inflation based on survey data. A second group of papers estimates directly time-varying the inflation risk premium within the context of general equilibrium consumption-based asset-pricing models. Some of these papers work with data for the US (Evans and Wachtel, 1992; Sarte, 1998; Ang and Bekaert, 2005; Buraschi and Jiltsov, 2005), while others work with UK data (Evans, 1998; Risa, 2001; Evans, 2003). With respect to the liquidity premium, to our knowledge there has not been direct estimation of it. The second strand of the literature studies the variation of the premiums. Based on the present discounted value model developed by Campbell and Ammer (1993), Barr and Pesaran (1997) assess whether unexpected inflation-compensation returns in the

UK are associated to news about future expected inflation or to future premiums.

In contrast to the number of studies for the US and the UK, only limited evidence exists for developing countries. Kandel et. al. (1996) and Balsam et. al. (1998) calibrate a consumption-based asset-pricing model and estimate the inflation risk premium for Israel. For the Chilean case, this line of research has remained largely unexplored. The only study related to these issues is the recent work by Jervis (2006). In the paper, the author follows the methodology of Shen (1998) and estimates the premiums residually using data of inflation compensation and survey data for expected inflation. The inconvenience of this approach is that it relies on survey data as a trustworthy measure of expected inflation, which, as explained above, may be questionable.

The purpose of this paper is to further study the relationship between inflation compensation and inflation expectations in Chile. In particular, we investigate whether inflation compensation provides a reliable source of information about inflation expectations in Chile. First, we use the present discounted value approach of Barr and Pesaran (1997) to decompose unanticipated inflation-compensation returns into news about expected inflation and premiums. Second, we use a simple general equilibrium asset pricing model, like in Evans and Wachtel (1992) and Sarte (1998), to estimate a time-varying inflation risk premium.

Our results indicate on the one hand that approximately from 40 to 65% of the variance of unexpected inflation-compensation returns are due to revisions in premiums and only from 21 to 25% to revisions in expected inflation. Thus it seems that the differential between nominal and indexed yields does not provide a reliable source of information about the way in which expectations of inflation change. On the other hand, the inflation risk premium estimated under the general equilibrium context seems to be insignificant. It is time-varying but its average size and volatility are very close to zero. Even though we do not have a direct measure of the liquidity premium, our results suggest that liquidity premium-movements could be a very important force behind inflation-compensation variation.

The rest of the paper is organized as follows. Section 2 presents the general framework used in the study. Section 3 shows the variance decomposition of unexpected inflation-compensation returns based on the present discounted value methodology. Section 4 estimates the inflation risk premium using a general equilibrium asset-pricing model, and section 5 concludes. An appendix contains figures and tables.

2 The general framework

In this section we study the basic framework used to understand the relationship between nominal, inflation-linked yields, and expected inflation.

Let $i_{n,t}$ denote the annualized gross yield of a n-period nominal discount bond observed in period t. Such asset pays $i_{n,t}$ units of money n periods ahead. Let p_t denote the price level in period t. The average annualized gross inflation rate between periods t and $t+n$ is defined as π_{t+n}, where $\pi_{t+n} = (p_{t+n}/p_t)^{1/n}$. The real payoff of this bond is $i_{n,t}/\pi_{t+n}$. Since inflation is stochastic, the real payoff will also be stochastic and varies inversely with inflation.

Inflation-linked bonds, on the other hand, compensate their holders against inflation. Let $r_{n,t}$ denote the annualized gross yield of a n-period inflation-linked discount bond. This bond pays the nominal amount of $r_{n,t} \times \pi_{t+n}$ at maturity, which in real terms equals $r_{n,t}$. The real payoff of the bond is not stochastic and therefore the bond provides full compensation for price changes through the period that the bond is held.

The Fisher hypothesis relates the variables described above. Since investors will always purchase the bond with higher expected real yield, bond prices should adjust such that both nominal and indexed bonds end up with the same expected real yield.

The Fisher hypothesis only holds under restrictive conditions. Specifically, it assumes that market participants are risk neutral and that markets are complete. When either of these conditions fails to hold, the Fisher equation must be adjusted.

The first adjustment arises when agents are risk averse. As described above, the real payoff of a nominal bond declines when inflation increases because the nominal value of the payoff is fixed when the bond is issued. Thus, real returns on nominal bonds moves inversely with the actual rate of inflation during the life of the security. Since risk averse investors dislike uncertainty, they will require compensation for holding the bond.[1] As a result, the nominal bond will have to carry a higher expected real return than the indexed bond in order to be equally attractive to investors. The extra expected real return is called inflation risk premium.

The second adjustment appears under market incompleteness. When markets are not complete, demand and supply of bonds affect security prices. Investors will demand compensation for holding the bonds that are less traded, since they might not be able to sell them quickly or will have to sell them at unfavorable prices. In Chile, in contrast to the US, the market for indexed bonds is more liquid than the market for nominal bonds.[2] As a result, part of the nominal yield should include an extra return, which is called liquidity premium.

In the more realistic world, inflation compensation is no longer an accurate measure of expected inflation. The yield spread now equals expected inflation plus the sum of the inflation risk and the liquidity premium. That is:

$$\ln i_{n,t} - \ln r_{n,t} = \ln \mathrm{E}_t[\pi_{t+n}] + (\Theta_{n,t} + \Phi_{n,t}), \tag{1}$$

where $\Theta_{n,t}$ stands for the inflation risk premium and $\Phi_{n,t}$ for the liquidity premium. In this world, inflation compensation can be higher or lower than the level of expected inflation, depending on the sign of the premiums. However, if both premiums are roughly constant, changes in the yield spread might still be an accurate measure of changes in expected inflation.

Finally, it is important to note that indexed bonds in real life are not exactly equivalent to true real bonds. The difference arises from indexation lags. The period over which indexation is calculated usually ends before the payment of the bond is made. Thus any inflation after the end of the indexation period and before the payment reduces the real value of the payoff. This gives origin to a third premium, known as the indexation-lag risk premium. The indexation

[1] Strictly speaking, investor dislike asset uncertainty that increases the volatility of their consumption path.

[2] Nominal debt consists in approximately only 25% of the total outstanding debt of the Central Bank of Chile at the beginning of 2006.

lag in some countries is large, like for example 8 months in UK, which traduces in a considerable indexation-lag premium (Evans, 1998). In Chile, however, the indexation lag is only 1 month. Chumacero (2002) proofs that the difference in the yield of an inflation-linked and a true real bond in Chile vanishes as the maturity of the instrument increases. Jervis (2006) suggests that the difference is negligible for maturities over one year. Thus in this paper we abstract from the indexation-lag premium.

3 Variance decomposition

In this section, we measure the extent to which variations in unexpected inflation-compensation returns are due to changing expectations of inflation, as opposed to changes in inflation risk and liquidity premiums. In particular, we follow Barr and Pesaran (1997) and use the present discounted value methodology of Campbell and Ammer (1993) to decompose inflation-compensation returns into news about expected inflation rates and news about premiums.

3.1 Present discounted value approach

The analysis uses the present discounted-value model to provide a structure within which news about expected inflation compete with news about premiums to explain unexpected inflation-compensation returns. These two sources of variation are treated separately by using a vector autoregression (VAR) to generate the forecast revisions that constitute news. We briefly summarize the methodology below.

We start by defining holding-period returns. Let the log of the one-period return on a n-period nominal bond be defined as:

$$\ln h^{i}_{n,t+1} = n \ln i_{n,t} - (n-1) \ln i_{n-1,t+1} \tag{2}$$

Equation (2) can be solved forward to the maturity date of the bond, using the fact that at this date the gross yield equals 1 so its log yield is 0. After taking expectations conditional on time t information, we obtain:

$$\ln i_{n,t} = \frac{1}{n}\mathrm{E}_t \sum_{\tau=0}^{n-1} \ln h^i_{n-\tau,t+1+\tau} \tag{3}$$

If we insert equation (3) into equation (2), we can express the innovation of the one-period return as a function of news of future bond returns:

$$\ln h^i_{n,t+1} - \mathrm{E}_t[\ln h^i_{n,t+1}] = -(\mathrm{E}_{t+1} - \mathrm{E}_t) \sum_{\tau=1}^{n-1} \ln h^i_{n-\tau,t+1+\tau} \tag{4}$$

This equation expresses the fact that bond returns are certain at maturity date, so unexpected negative returns today will necessarily be offset by increases in returns tomorrow. By defining the one-period real excess return of the nominal bond as $\ln x^i_{n,t+1} = \ln h^i_{n,t+1} - \ln \pi_{t+1} - \ln r_{1,t}$, we can rewrite equation (4) in terms of excess returns:

$$\ln x^i_{n,t+1} - \mathrm{E}_t[\ln x^i_{n,t+1}] = -(\mathrm{E}_{t+1} - \mathrm{E}) \left\{ \sum_{\tau=1}^{n-1} \ln \pi_{t+1+\tau} + \sum_{\tau=1}^{n-1} \ln r_{1,t+\tau} + \sum_{\tau=1}^{n-1} \ln x^i_{n,t+1+\tau} \right\} \tag{5}$$

To simplify notation, we define $\widehat{x}^i_{n,t+1}$ as the unexpected component of the nominal excess return, $\widehat{x}_{n,t+1}(\pi)$ as the term that represents news about inflation rates, $\widehat{x}_{n,t+1}(r)$ as the term that represents news about interest rates, and $\widehat{x}^i_{n,t+1}(x)$ as the term representing news about future excess returns. Then equation (5) can be rewritten as:

$$\widehat{x}^i_{n,t+1} = -\widehat{x}_{n,t+1}(\pi) - \widehat{x}_{n,t+1}(r) - \widehat{x}^i_{n,t+1}(x) \tag{6}$$

Equation (6) stands that unexpected nominal bond returns must be associated either with decreases in expected inflation rates over the life of the bond, with decreases in future real interest rates or with decreases in future excess bond returns.

Analogous results apply to inflation-linked bonds. By defining the log of the one-period return on an n-period indexed bond as $\ln h^r_{n,t+1}$, we can express the one-period return innovation as a function of news of future bond returns:

$$\ln h^r_{n,t+1} - \mathrm{E}_t[\ln h^r_{n,t+1}] = -(\mathrm{E}_{t+1} - \mathrm{E}_t) \sum_{\tau=1}^{n-1} \ln h^r_{n-\tau,t+1+\tau} \tag{7}$$

After defining the one-period excess return of the inflation-linked bond as $\ln x^r_{n,t+1} = \ln h^r_{n,t+1} - \ln r_{1,t}$, we can rewrite equation (7) as:

$$\ln x^r_{n,t+1} - \mathrm{E}_t[\ln x^r_{n,t+1}] = -(\mathrm{E}_{t+1} - \mathrm{E}) \left\{ \sum_{\tau=1}^{n-1} \ln r_{1,t+\tau} + \sum_{\tau=1}^{n-1} \ln x^r_{t+1+\tau} \right\} \tag{8}$$

Or in more compact form:

$$\widehat{x}^r_{n,t+1} = -\widehat{x}^r_{n,t+1}(r) - \widehat{x}^r_{n,t+1}(x) \tag{9}$$

The difference between equations (6) and (9) is that inflation news do not appear in the latter. Since the indexed bond protects against inflation, unexpected movements in inflation do not affect its return. If we subtract (9) from (6) we get:

$$\widehat{x}^c_{n,t+1} = -\widehat{x}_{n,t+1}(\pi) - \widehat{x}^c_{n,t+1}(x), \tag{10}$$

where $\widehat{x}^c_{n,t+1} = \widehat{x}^i_{n,t+1} - \widehat{x}^r_{n,t+1}$ indicates unexpected inflation-compensation returns (i.e. the unforeseen excess return of a long position on a nominal bond and a short position on a inflation-linked bond) and $\widehat{x}^c_{n,t+1}(x) = \widehat{x}^i_{n,t+1}(x) - \widehat{x}^r_{n,t+1}(x)$ indicates news about future relative premiums (i.e. inflation risk and liquidity premiums). News about future real interest rates do not appear in the equation since they cancel out. Therefore, unexpected inflation-compensation returns must be associated either with decreases in expected inflation rates or decreases in future premiums.

Based on equation (10), we are able to quantify the relative importance of the different components of inflation-compensation returns. In particular, (10) implies that the variance of the unexpected inflation-compensation returns can be written as:

$$\mathrm{var}(\widehat{x}^c_{n,t+1}) = \mathrm{var}(\widehat{x}_{n,t+1}(\pi)) + \mathrm{var}(\widehat{x}^c_{n,t+1}(x)) + 2\mathrm{cov}(\widehat{x}_{n,t+1}(\pi), \widehat{x}^c_{n,t+1}(x)) \tag{11}$$

The variance decomposition of (11) separates movements in surprise inflation compensation returns into two components: (i) news about expected inflation, and (ii) news about premiums.

Since revisions in expectations are not directly observable, they need to be estimated. Campbell and Ammer (1993) and Barr and Pesaran (1997) assume that the expectations can be proxied by forecasts based on a VAR that incorporates a range of financial variables, and that the agents' set of information is reflected in these variables. The VAR approach begins by defining a state vector $\mathbf{w}$ that contains the one-period inflation-compensation return, the one-period inflation rate and other variables (which will be described below) that help forecasting excess returns. Next, the state vector is assumed to follow a p-order VAR process:

$$\mathbf{w}_{t+1} = \boldsymbol{A}(\boldsymbol{L})\mathbf{w}_{t+1} + \boldsymbol{\varepsilon}_{t+1}, \tag{12}$$

where $A(L)$ is a lag polynomial of order p and ε_{t+1} is a vector of white noise errors. To obtain estimates of revisions to expectations, we use the fact that:

$$(\mathrm{E}_{t+1} - \mathrm{E}_{t})\mathbf{w}_{t+1+\tau} = \boldsymbol{A}(\boldsymbol{L})^{\tau}\boldsymbol{\varepsilon}_{t+1} \tag{13}$$

We obtain revisions to expectations of future inflation rates and current inflation-compensation returns by direct forecasting, leaving the revision in future premiums as the residual of equation (10) after substitution of the VAR-generated series.[3]

3.2 Data

Our sample is monthly and runs from September, 2002 to March, 2006. The size of the sample is limited by the fact that the Central Bank of Chile began issuing nominal bonds only in late 2002. The market for nominal debt is less liquid than the market for inflation-linked debt. In the beginning of 2006 the stock of the Central Bank's nominal debt was 3,150 million dollars, corresponding to approximately 25% of the total outstanding debt.

Interest rates considered in this section include the yields of nominal and inflation-linked discount bonds with maturities from one through five years. The data for the nominal and inflation-linked yield curves was provided by `Riskamerica`, that computes the term structures by using a dynamic model estimated from incomplete panel data. The term structure of inflation compensation

[3]This choice of residual is forced since the sequence of excess returns on the bond as its maturity falls is no directly measurable.

is obtained by subtracting the log of the inflation-linked (gross) yield curve from the log of the nominal (gross) yield curve.

Figure 1, already shown in the introduction, depicts the evolution of inflation compensations with maturities from one to five years, during our sample period. According to the figure, the term structure of inflation compensation is upward-sloping in some occasions and downward-sloping in others. The figure also shows that inflation compensation of different maturities are highly correlated and move in tandem.

Table 1 reports some summary statistics for inflation compensations of different maturities.

[Insert table 1]

According to the table, average inflation compensation for different maturities is around 13 to 28 basis points below the Central Bank's inflation target of 3%. The table also shows that inflation compensation is very volatile, with standard deviation as a fraction of mean ranging from 18 to 30%. The high range also confirms this result. Finally, the autocorrelation of inflation compensation decreases with maturity. In fact, inflation compensations for maturities up to 3 years are positively serially autocorrelated, and beyond are negatively autocorrelated, indicating the presence of mean reversion.

In order to estimate the VAR in equation (12) the state vector must include at least the one-period inflation-compensation return and the one-period inflation rate. In addition we include a set of variables that have shown to forecast excess returns (Campbell and Ammer, 1993; Barr and Pesaran, 1997). These variables include the long-short nominal yield spread, the long-short indexed yield spread, the long nominal yield, and the long indexed yield. Thus, the state vector equals:

$$\mathbf{w}_t = [\widehat{x}^c_{n,t+1}; \ln \pi_{t+1}; \ln i_{n,t} - \ln i_{1,t}; \ln r_{n,t} - \ln r_{1,t}; \ln i_{n,t}; \ln r_{n,t}]' \tag{14}$$

Throughout the paper, n will correspond to years.

3.3 Results

We estimate a first-order VAR including the six variables described above to obtain revisions of expected future inflation rates, unexpected contemporaneous inflation-compensation returns, and future premiums. Since the number of variables in the VAR increases very rapidly with lag length, and our sample size is already small, we choose a parsimonious first-order VAR to preserve degrees of freedom.

Table 2 reports the matrix of estimated first-order VAR coefficients, their standard deviations, and the R^2 of each equation. The VAR is estimated separately for maturities of two to five years.

[Insert table 2]

The VAR produces quite reasonable forecasting power as measured by R^2. We included other variables in the VAR (such as survey-based inflation expectations) and the results remained practically unchanged.[4]

Using the coefficients of the estimated VAR, we decompose unforseen inflation-compensation returns into news about inflation rates and premiums. Results are reported in table 3. The variances and covariances of the different components of the relative return are normalized by the variance of the return innovation itself so the number reported are shares adding up to one.

[Insert table 3]

According to the table, revisions to future inflation is not the dominant factor in explaining inflation-compensation returns. Revisions to expected future inflation explain only from 22% to 25% of the variance of unexpected relative returns. The remaining 75% to 78% is explained by revisions to future premiums and to their positive correlation with inflation news. The substantial role played by premium news suggests that movements in inflation compensations cannot be used as accurate measures of movements in expected inflation.

[4]The results of these estimations are available upon request.

4 Inflation risk premium

In this section we estimate the level of the inflation risk premium. Specifically, we use a simple general equilibrium consumption-based asset-pricing model as in Evans and Wachtel (1992) and Sarte (1998) to estimate a time-varying inflation premium that will depend on the covariance between consumption growth and inflation. Even though this framework has been proved to perform poorly in the US economy, primarily due to the lack of variability of the stochastic discount factor (Cochrane and Hansen, 1992), we are more confident in using it considering the relatively high volatility of Chile's economy.

4.1 Methodology

Consider the standard intertemporal optimization problem facing a representative agent in an endowment economy with money (Lucas, 1980). We can write the maximization problem for the agent as:

$$\max \mathrm{E}_0 \sum_{t=0}^{\infty} \beta^t \left(\frac{c_t^{1-\gamma} - 1}{1-\gamma} \right), \tag{15}$$

where c_t denotes real consumption of the single good, $0 < \beta < 1$ is the subjective discount factor, and γ is the coefficient of relative risk aversion.

We denote $b_{n,t}^i$ the demand for the n-period discount nominal bond at period t and $b_{n,t}^r$ the demand for the n-period discount indexed bond. The optimization problem will be subject to the following budget constraint for each period:

$$y_t + \sum_{n=1}^{\infty} i_{n,t-n}^n \frac{b_{n,t-n}^i}{p_t} + \sum_{n=1}^{\infty} r_{n,t-n}^n b_{n,t-n}^r + \frac{m_{t-1}}{p_t} \geq c_t + \sum_{n=1}^{\infty} \frac{b_{n,t}^i}{p_t} + \sum_{n=1}^{\infty} b_{n,t}^r + \frac{m_t}{p_t} \tag{16}$$

Here y_t is the agent's endowment and m_t stands for the demand for money balances.[5] The first-order conditions of the problem are given by:

$$i_{n,t}^{-n} = \mathrm{E}_t[\beta^n \left(c_{t+n}/c_t\right)^{-\gamma} \left(p_{t+n}/p_t\right)^{-1}] \tag{17a}$$

$$r_{n,t}^{-n} = \mathrm{E}_t[\beta^n \left(c_{t+n}/c_t\right)^{-\gamma}] \tag{17b}$$

[5] We assume a cash-in-advance constraint induces agents to hold money.

The term $\beta^n(c_{t+n}/c_t)^{-\gamma}$ represents the real stochastic discount factor or pricing kernel, and will be denoted by s^n_{t+n}. If we decompose equation (17a) using the property that for any two random variables x and y, $\mathrm{E}[xy] = \mathrm{E}[x]\mathrm{E}[y] + \mathrm{cov}(x, y)$, and we insert equation (17b) into the resulting expression, we get the following equation:

$$i^{-n}_{n,t} = r^{-n}_{n,t}\mathrm{E}_t[\pi^{-n}_{t+n}] + \mathrm{cov}_t\left[s^n_{t+n}, \pi^{-n}_{t+n}\right] \tag{18}$$

If we disregard the Jensen's inequality term and assume that $\mathrm{E}_t[\pi^{-1}_{t+n}] \simeq \mathrm{E}_t[\pi_{t+n}]^{-1}$ and then factorize, we get:

$$i^{-n}_{n,t} = r^{-n}_{n,t}\mathrm{E}_t[\pi^{-n}_{t+n}]^{-1}\left\{1 + \frac{\mathrm{cov}_t\left[s^n_{t+n}, \pi^{-n}_{t+n}\right]}{\mathrm{E}_t[s^n_{t+n}]\mathrm{E}_t[\pi^{-n}_{t+n}]}\right\}, \tag{19}$$

After applying logs and rearranging terms, we finally get:

$$\ln i_{n,t} - \ln r_{n,t} = \ln \mathrm{E}_t[\pi_{t+n}] + \Theta_{n,t} \tag{20}$$

where the risk premium $\Theta_{n,t}$ is defined by:

$$\Theta_{n,t} = -\frac{1}{n}\ln\left\{1 + \frac{\mathrm{cov}_t\left[s^n_{t+n}, \pi^{-n}_{t+n}\right]}{\mathrm{E}_t[s^n_{t+n}]\mathrm{E}_t[\pi^{-n}_{t+n}]}\right\} \tag{21}$$

Equation (20) is a generalized version of the Fisher equation, adjusted to include an inflation risk premium. The premium depends on the covariance between consumption growth and inflation. The covariance term refers to the usefulness of the nominal bond in smoothing consumption over states of nature. Suppose the covariance is negative. Then inflation will be high when consumption growth is low, and the real payoff of the bond will be low precisely when consumption is most valued. Since the nominal bond serves as a poor hedge against inflation risk, it will have to offer a higher interest rate to induce agents to hold it.

In order to estimate the inflation risk premium in (21), we follow the methodology described in Balsam et. al. (1998). We assume that the rational expectations hypothesis holds, so realized values will differ from their conditional expectations by an error term that is unpredictable given the agent's information set. Therefore:

$$\begin{bmatrix}\varepsilon^s_{n,t}\\ \varepsilon^\pi_{n,t}\end{bmatrix} = \begin{bmatrix}s^n_{t+n} - \mathrm{E}_t[s^n_{t+n}]\\ \pi^{-n}_{t+n} - \mathrm{E}_t[\pi^{-n}_{t+n}]\end{bmatrix} \tag{22}$$

Using the definition of covariance, we can rewrite the inflation premium as:

$$\Theta_{n,t} = -\frac{1}{n} \ln \left\{ 1 + \frac{\mathrm{E}_t[\varepsilon^s_{n,t}\varepsilon^\pi_{n,t}]}{\mathrm{E}_t[s^n_{t+n}]\mathrm{E}_t[\pi^{-n}_{t+n}]} \right\} \tag{23}$$

The estimation of the premium is done in a two-stage procedure. First, we compute the pair of innovations $\varepsilon^s_{n,t}$ and $\varepsilon^\pi_{n,t}$ from an estimated VAR(p) process for s^n_{t+n} and π^{-n}_{t+n}:

$$\begin{bmatrix} s^n_{t+n} \\ \pi^{-n}_{t+n} \end{bmatrix} = \begin{bmatrix} A(L)s^n_{t+n} + B(L)\pi^{-n}_{t+n} \\ C(L)s^n_{t+n} + D(L)\pi^{-n}_{t+n} \end{bmatrix} + \begin{bmatrix} \varepsilon^s_{n,t} \\ \varepsilon^\pi_{n,t} \end{bmatrix}, \tag{24}$$

where $A(L)$, $B(L)$, $C(L)$ and $D(L)$ are lag polynomials of order p. In the second step, we generate the conditional covariation between the innovations from an estimated AR(q) process for the product of the innovations:

$$\varepsilon^s_{n,t}\varepsilon^\pi_{n,t} = E(L)\varepsilon^s_{n,t}\varepsilon^\pi_{n,t} + \mu^{s\pi}_{n,t} \tag{25}$$

where $E(L)$ is a lag polynomial of order q and $\mu^{s\pi}_{n,t}$ is a white noise error. Finally, by taking expectations of equations (24) and (25) we obtain the following expression for the risk premium:

$$\Theta_{n,t} = -\frac{1}{n} \ln \left\{ 1 + \frac{E(L)\varepsilon^s_{n,t}\varepsilon^\pi_{n,t}}{[A(L)s^n_{t+n} + B(L)\pi^{-n}_{t+n}][C(L)s^n_{t+n} + D(L)\pi^{-n}_{t+n}]} \right\} \tag{26}$$

4.2 Data

For this section we use quarterly data set running from the first quarter of 1986 to the first quarter of 2006. The estimation of the risk premium does not require data on nominal and inflation-linked yields, which allows us to use a larger sample than the one used in the previous section.

The price measure used consists in the the consumer price index. Real consumption is measured a per-capita consumption expenditures on non-durables deflated by the consumer price index. Summary statistics on inflation and growth in real consumption for different maturities are presented in table 4.

[Insert table 4]

Real consumption grew on average around 4.6% during the period 1986-2006. The consumer price index grew approximately 9.6% in the same period. Inflation has been more than twice volatile as consumption growth during our sample. Inflation has also presented more persistence.

We checked the stationary of our vector of stochastic process using a Dickey-Fuller test. We found no evidence of a unit root behavior in either of the series. However, inflation presented a deterministic trend and therefore we removed a linear trend from the series before using it in the VAR.

4.3 Results

We define a permissible domain for the preference parameters. Specifically, we calibrate the subjective discount factor throughout such that $\beta = 0.987$ (Cooley and Prescott, 1995). We let the CRRA coefficient vary between the values of 5, 10, 15, and 20. [6]

For the first stage of the procedure, we estimate a VAR with four lags (p=4).[7] The VAR is estimated separately for maturities from one to five years. Results are reported in table 5.

[Insert table 5]

The fit of both stochastic processes seems to be satisfactory. After computing the residuals of the VAR, we estimate a fourth-order autoregression process for the product of the residuals (q=4) and compute the inflation risk premium.

In figure 2 we depict the evolution of the inflation risk premium during our sample for maturities of one year trough five years, for a CRRA coefficient of 10.

[Insert figure 2]

From the figure we can see that the premium is not constant over time. Even though the premium has reached values ranging from 0.25% to -0.35%, the series

[6] The values of the CRRA coefficient used in the study are stringent. The highest value Balsam et al. (1998) use in their calibration for Israel is 10. Furthermore, Jervis (2006) estimates a value of 3 using a GMM procedure for Chile.

[7] This seemed to be the best lag-specification according to the Akaike information criteria.

has been closed to zero nearly all of the time. Table 6 reports summary statistics of the inflation risk premium for different maturities and different values of risk aversion.

[Insert table 6]

We find that the average premium for maturities of one and two years is positive, and for maturities beyond is negative. This arises from the fact that in our sample the correlation between inflation and consumption growth increases as the horizon increases. As a result, nominal bonds of longer horizons provide a better hedge against consumption fluctuations than nominal bonds of shorter horizons and therefore demand less risk compensation. Furthermore, as the CRRA coefficient increases, short-term premiums become more positive and long-term premiums become more negative. However, the absolute value of the average premium is extremely small in all cases. For example, the one-year premium ranges from 0.01% for $\gamma = 5$ to 0.09% for $\gamma = 20$. Similarly, the five-year premium ranges from -0.003% for $\gamma = 5$ to -0.015% for $\gamma = 20$. Volatility (measured by standard deviation) is also very small, ranging from 0.001% to 0.006% for the one-year premium, and from 0.000% to 0.070% for the five year premium.

Adding up, the inflation risk premium during our sample seems to have been negligible. This is true for relatively high values of risk aversion, and therefore will be even more so for the traditional values of risk aversion (below 5) used in the literature (Cooley and Prescott, 1995). The result is robust to using the same sample used in the previous section, consisting of monthly observations for the period 2002-2006.[8] The lack of strong covariation between consumption growth and inflation during the sample period suggests that inflation risk has been low and therefore its price has also been low.

Finally, recall from the previous section that we found that movements in premiums accounted for approximately 40 to 65% of movements in inflation-compensation returns. The premium-movements could be associated to both inflation risk or liquidity premium movements. However recall from table 1 that

[8]Results for this alternative sample are not reported for space reasons, although they are available upon request.

the standard deviation of one-year inflation compensation was 0.82%. This figure is 12 times larger than the highest standard deviation of the one-year inflation risk premium (for $\gamma = 20$). Hence it is unlikely that the inflation risk premium plays an important role in overall premium variation. Thus, even if we do not have a direct measure of the liquidity premium, these results suggest that liquidity-premium movements could have an important role in explaining inflation-compensation returns.

5 Concluding remarks

The difference between the yields of nominal and inflation-linked bonds, known as inflation compensation, contains useful information about market expectations of inflation. These expectations are key in the inflation-targeting framework that currently guides monetary policy in Chile, since they make it possible to evaluate the markets' perception of the Central Bank's commitment to the target and help the monetary authority assess its credibility among market participants.

However, the task of disentangling this information is complicated by the existence of other components in the yield differential, in particular the inflation risk premium and the liquidity premium. These premiums might shift the level of inflation compensation from true inflation expectations. However, if the premiums are relatively stable over time, inflation compensation might still be a useful indicator, since movements in the yield spread would be associated exclusively to movements in expected inflation.

In this paper we have studied the relationship between inflation compensation and inflation expectations in Chile. We first study the degree of time variation of inflation compensation. Using a sample of monthly observations of nominal and inflation-linked yields for the period 2002-2006, we decompose unanticipated inflation-compensation returns of different maturities into news about expected inflation and premiums. The results show that premiums are time-varying an have an important role in explaining yield-differential movements. In fact, premiums explain around 40% to 65% of inflation-compensation return variance. Hence variations in inflation compensation are not necessarily indicative of variations in inflation expectations.

We then proceed to study the levels of the premiums. It might be the case that the premiums are very volatile but have a mean value close to zero. In this case, the unconditional mean of inflation compensation would equal the unconditional mean of inflation. Since we lack of a methodology for estimating the liquidity premium, we only estimate the inflation risk premium, using a simple general equilibrium model. The estimation procedure requires only data from consumption growth and inflation, which allows to use a larger sample of quarterly frequency running from 1986 to 2006. According to the results, the risk premium is time-varying but seems to be of negligible size. Its average mean and volatility across different maturities and degrees of risk aversion is very close to zero.

Overall, since the risk premium is not considerable and premiums explain a large part of the variance of inflation-compensation returns, our results suggest that the liquidity premium might be an important driving force behind inflation-compensation movements. Since we do not have a direct estimate of this premium, we cannot determine its average size and therefore we do not know how far this premium shifts the level of inflation compensation from true expected inflation. However, recall that the liquidity premium arises when markets are incomplete. Since financial markets become more complete as a country becomes more economically developed (Jung, 1986), we could expect the importance of this premium to decrease with time, with the result that inflation compensation would become a more accurate measure of inflation expectations.

To conclude, we think our analysis can be extended in a number of useful directions. First, the general equilibrium model used to estimate the inflation risk premium is extremely simple and could be extended to include an open economy dimension (Bekaert et. al., 2002), more risk factors (Risa, 2001), different regimes (Evans, 2003) or consumption habits (Buraschi and Jiltsov, 2007). Second, efforts should be allocated to obtain a direct measure of the liquidity premium. Only after having estimated both premiums we will have a more complete picture of the relationship between inflation compensation and expected inflation in Chile.

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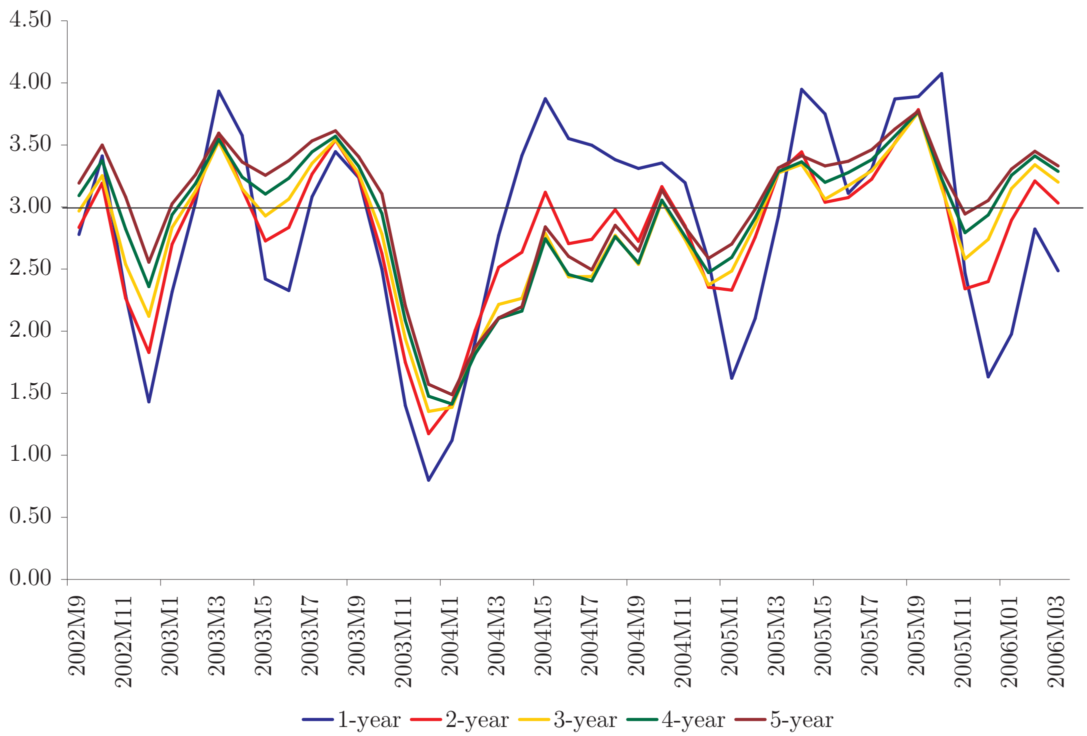
Figure 1: Inflation compensation for different maturities in Chile (in %)



The figure depicts the monthly evolution of inflation compensation (defined as the differential between the yield of a nominal bond and an indexed bond) for maturities from one to five years, during the period September 2002-March 2006.

Table 1: Descriptive statistics for inflation compensation of different maturities (in %)

Maturity in years (n)	Mean	Std dev	Max	Min	Persistence
1	2.78	0.82	3.95	0.77	0.54
2	2.72	0.54	3.66	1.14	0.32
3	2.73	0.53	3.64	1.30	0.00
4	2.79	0.53	3.64	1.36	-0.14
5	2.87	0.52	3.63	1.43	-0.18

This table reports descriptive statistics for monthly observations of inflation compensation for maturities from one to five years, for the period September, 2002 to March, 2006. All variables are expressed in percent per year. Persistence denotes the 12-month autocorrelation coefficient.

Table 2: VAR coefficients estimates for the variance decomposition

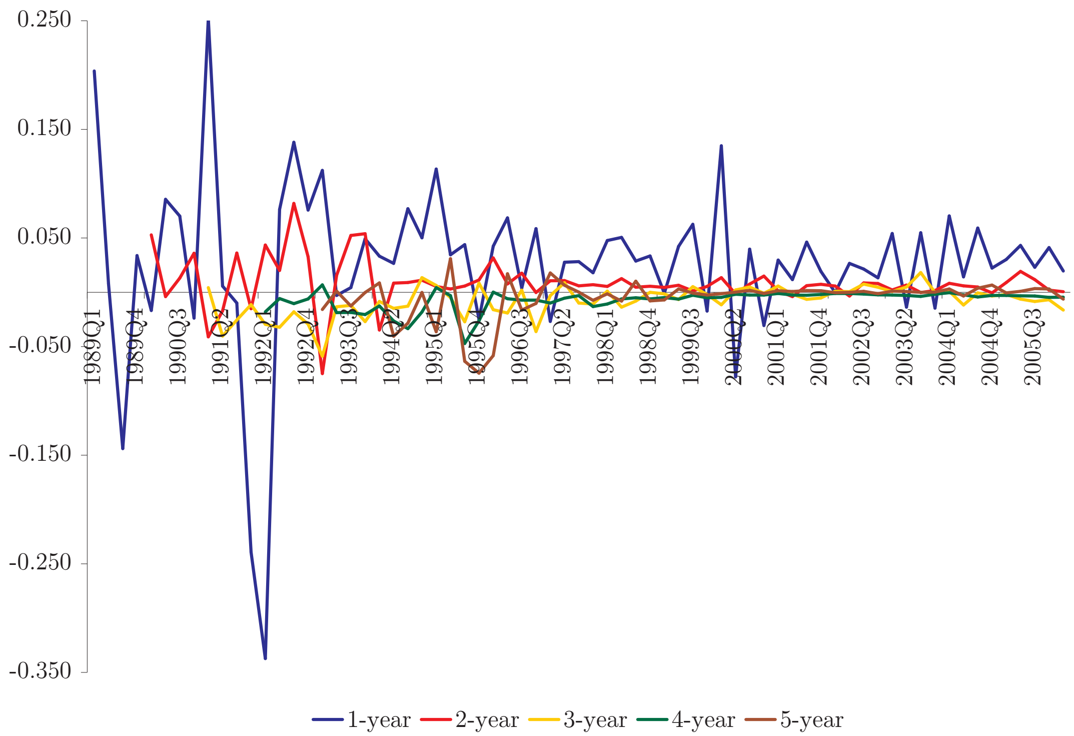
	VAR estimates for $n=2$						
	$x_{n,t}$	$\ln \pi_{n,t}$	$(\ln i_{n,t-1}-$ $\ln i_{1,t-1})$	$(\ln r_{n,t-1}-$ $\ln r_{1,t-1})$	$\ln i_{n,t-1}$	$\ln r_{n,t-1}$	R^2
$x_{n,t+1}$	0.83	0.45	4.02	0.52	-0.57	1.32	0.90
	(0.25)	(0.40)	(1.49)	(0.39)	(0.53)	(0.39)	-
$\ln \pi_{n,t+1}$	-0.12	0.58	-1.42	-0.05	0.30	-0.33	0.92
	(0.16)	(0.25)	(0.93)	(0.24)	(0.33)	(0.24)	-
$\ln i_{n,t} - \ln i_{1,t}$	0.00	-0.04	0.66	0.02	-0.02	0.00	0.95
	(0.02)	(0.03)	(0.11)	(0.03)	(0.04)	(0.03)	-
$\ln r_{n,t} - \ln r_{1,t}$	0.05	-0.11	-0.82	0.30	0.19	-0.70	0.86
	(0.10)	(0.15)	(0.57)	(0.15)	(0.21)	(0.15)	-
$\ln i_{n,t}$	0.05	0.15	-1.28	0.08	0.84	-0.27	0.98
	(0.05)	(0.09)	(0.33)	(0.08)	(0.12)	(0.09)	-
$\ln r_{n,t}$	0.00	0.11	0.53	0.28	0.37	0.74	0.87
	(0.12)	(0.20)	(0.74)	(0.19)	(0.27)	(0.19)	-
	VAR estimates for $n=3$						
$x_{n,t+1}$	0.92	0.45	2.30	-0.10	0.21	0.62	0.86
	(0.23)	(0.58)	(1.34)	(0.45)	(0.90)	(0.88)	-
$\ln \pi_{n,t+1}$	-0.03	0.67	-0.82	0.11	0.35	-0.41	0.92
	(0.09)	(0.21)	(0.50)	(0.17)	(0.33)	(0.33)	-
$\ln i_{n,t} - \ln i_{1,t}$	0.00	-0.08	0.68	0.04	-0.04	0.04	0.96
	(0.02)	(0.05)	(0.11)	(0.04)	(0.07)	(0.07)	-
$\ln r_{n,t} - \ln r_{1,t}$	0.08	-0.15	-0.20	0.37	0.10	-0.97	0.89
	(0.07)	(0.17)	(0.40)	(0.13)	(0.27)	(0.26)	-
$\ln i_{n,t}$	0.02	0.09	-0.64	0.10	0.81	-0.27	0.96
	(0.03)	(0.08)	(0.19)	(0.06)	(0.13)	(0.13)	-
$\ln r_{n,t}$	0.04	0.09	-0.04	0.14	0.36	0.59	0.75
	(0.06)	(0.15)	(0.35)	(0.12)	(0.24)	(0.23)	-

Table 2 continued

	VAR estimates for $n = 4$						
	$x_{n,t}$	$\ln \pi_{n,t}$	$(\ln i_{n,t-1} - \ln i_{1,t-1})$	$(\ln r_{n,t-1} - \ln r_{1,t-1})$	$\ln i_{n,t-1}$	$\ln r_{n,t-1}$	R^2
$x_{n,t+1}$	0.95	0.38	1.72	-0.31	0.67	0.33	0.87
	(0.22)	(0.76)	(1.43)	(0.62)	(1.22)	(1.36)	-
$\ln \pi_{n,t+1}$	0.00	0.69	-0.59	0.10	0.39	-0.53	0.92
	(0.06)	(0.20)	(0.39)	(0.17)	(0.33)	(0.37)	-
$\ln i_{n,t} - \ln i_{1,t}$	-0.01	-0.12	0.65	0.07	-0.07	0.13	0.96
	(0.02)	(0.06)	(0.12)	(0.05)	(0.10)	(0.11)	-
$\ln r_{n,t} - \ln r_{1,t}$	0.07	-0.22	0.14	0.40	-0.01	-0.99	0.91
	(0.05)	(0.19)	(0.35)	(0.15)	(0.30)	(0.34)	-
$\ln i_{n,t}$	0.01	0.04	-0.45	0.13	0.78	-0.19	0.91
	(0.03)	(0.09)	(0.17)	(0.07)	(0.14)	(0.16)	-
$\ln r_{n,t}$	0.03	0.05	-0.11	0.09	0.31	0.60	0.66
	(0.04)	(0.13)	(0.25)	(0.11)	(0.21)	(0.23)	-
	VAR estimates for $n = 5$						
$x_{n,t+1}$	0.95	0.30	1.36	-0.43	1.04	0.33	0.88
	(0.21)	(0.92)	(1.61)	(0.77)	(1.50)	(1.75)	-
$\ln \pi_{n,t+1}$	0.01	0.69	-0.44	0.08	0.42	-0.64	0.92
	(0.05)	(0.20)	(0.35)	(0.17)	(0.33)	(0.38)	-
$\ln i_{n,t} - \ln i_{1,t}$	-0.01	-0.15	0.60	0.11	-0.10	0.24	0.96
	(0.02)	(0.07)	(0.13)	(0.06)	(0.12)	(0.14)	-
$\ln r_{n,t} - \ln r_{1,t}$	0.05	-0.27	0.29	0.44	-0.11	-0.87	0.91
	(0.05)	(0.20)	(0.35)	(0.17)	(0.33)	(0.38)	-
$\ln i_{n,t}$	0.00	0.00	-0.40	0.17	0.74	-0.09	0.82
	(0.02)	(0.10)	(0.17)	(0.08)	(0.16)	(0.19)	-
$\ln r_{n,t}$	0.03	0.03	-0.15	0.08	0.26	0.65	0.70
	(0.03)	(0.12)	(0.21)	(0.10)	(0.19)	(0.22)	-

This table reports coefficient estimates and their standard deviations (in parenthesis) for a monthly one-lag VAR that includes the inflation-compensation return, one-year inflation rate, long-short nominal yield spread, long-short indexed yield spread, long nominal yield, and long indexed yield, for the period September, 2002 to March, 2006. The VAR is estimated separately for maturities from one to five years.

Figure 2: Inflation risk premium for different maturities in Chile (in %)



The figure depicts the evolution of the quarterly inflation risk premium (generated in accordance with equation (26)) for maturities from one to five years, during the period 1989:q1-2006:q3. The CRRA coefficient γ is assumed to take the value of 10.

Table 3: Variance decomposition for innovations in inflation-compensation returns of different maturities (in %)

Maturity in years (n)	Share in $\mathrm{var}(\widehat{x}^c_{n,t+1})$ of		
	$\mathrm{var}(\widehat{x}_{n,t+1}(\pi))$	$\mathrm{var}(\widehat{x}^c_{n,t+1}(x))$	2cov $(\widehat{x}_{n,t+1}(\pi), \widehat{x}^c_{n,t+1}(x))$
2	23.91	39.42	36.67
3	25.40	54.84	19.77
4	23.74	65.39	10.86
5	21.49	66.35	12.15

This table is based on the VAR presented in table 2. The VAR is used to calculate the components of the unexpected inflation-compensation return. The table reports the variances and covariances of these components, divided by the variance of the inflation-compensation return, so that the number reported sum up to 100%.

Table 4: Summary statistics for consumption growth and inflation of different maturities

Maturity in years (n)	Consumption growth				
	Mean	Std dev	Max	Min	Persistence
1	4.76	3.45	19.74	-3.16	0.13
2	4.63	2.45	12.43	-0.03	0.63
3	4.58	2.17	9.48	0.93	0.78
4	4.54	2.03	8.32	0.97	0.82
5	4.59	1.92	7.87	1.26	0.85

Maturity in years (n)	Inflation				
	Mean	Std dev	Max	Min	Persistence
1	9.71	7.47	29.15	0.02	0.85
2	9.58	7.10	25.23	1.15	0.93
3	9.58	6.87	22.80	2.02	0.97
4	9.58	6.67	20.94	2.12	0.98
5	9.46	6.32	20.36	2.50	0.99

This table reports descriptive statistics for quarterly observations of annualized consumption growth and inflation (defined as $\ln(c_{t+n}/c_t)^{1/n}$ and $\ln(p_{t+n}/p_t)^{1/n}$ respectively) for maturities from one to five years, for the period 1986:q1 to 2006:q1. All variables are expressed in percent per year. Persistence denotes the 4-quarter autocorrelation coefficient.

Table 5: VAR coefficient estimates for the estimation of the inflation risk premium (in %)

	VAR estimates for $n=1$								
	s^n_{t-1+n}	s^n_{t-2+n}	s^n_{t-3+n}	s^n_{t-4+n}	π^{-n}_{t-1+n}	π^{-n}_{t-2+n}	π^{-n}_{t-3+n}	π^{-n}_{t-4+n}	R^2
s^n_{t+n}	0.71	0.12	-0.12	-0.03	-2.20	4.61	-4.41	2.53	0.63
	(0.12)	(0.15)	(0.15)	(0.12)	(1.15)	(1.86)	(1.90)	(1.20)	-
π^{-n}_{t+n}	0.03	-0.03	0.03	-0.01	1.34	-0.30	-0.12	-0.05	0.90
	(0.01)	(0.02)	(0.02)	(0.01)	(0.12)	(0.20)	(0.21)	(0.13)	-
	VAR estimates for $n=2$								
s^n_{t+n}	0.68	0.22	-0.26	0.26	-2.37	3.77	-1.71	0.93	0.83
	(0.12)	(0.15)	(0.14)	(0.12)	(1.16)	(2.03)	(2.08)	(1.25)	-
π^{-n}_{t+n}	0.01	0.00	0.01	-0.02	1.39	-0.39	0.19	-0.29	0.96
	(0.01)	(0.02)	(0.02)	(0.01)	(0.12)	(0.21)	(0.22)	(0.13)	-
	VAR estimates for $n=3$								
s^n_{t+n}	0.61	0.18	-0.13	0.26	-2.56	3.93	-0.51	0.12	0.89
	(0.13)	(0.15)	(0.14)	(0.12)	(1.47)	(2.54)	(2.50)	(1.43)	-
π^{-n}_{t+n}	0.00	-0.01	0.00	0.01	1.30	-0.40	0.45	-0.39	0.97
	(0.01)	(0.01)	(0.01)	(0.01)	(0.12)	(0.21)	(0.21)	(0.12)	-
	VAR estimates for $n=4$								
s^n_{t+n}	0.70	0.19	-0.16	0.20	-4.02	4.77	0.22	0.01	0.93
	(0.14)	(0.17)	(0.16)	(0.13)	(1.53)	(2.46)	(2.45)	(1.48)	-
π^{-n}_{t+n}	0.00	0.00	-0.01	0.00	1.02	-0.12	0.32	-0.25	0.97
	(0.01)	(0.01)	(0.01)	(0.01)	(0.13)	(0.20)	(0.20)	(0.12)	-
	VAR estimates for $n=5$								
s^n_{t+n}	0.76	0.20	-0.33	0.32	-3.95	3.95	1.16	-0.05	0.95
	(0.13)	(0.17)	(0.16)	(0.13)	(1.96)	(3.45)	(3.45)	(1.96)	-
π^{-n}_{t+n}	-0.01	0.01	0.00	0.00	1.42	-0.70	0.49	-0.25	0.98
	(0.01)	(0.01)	(0.01)	(0.01)	(0.14)	(0.25)	(0.25)	(0.14)	-

This table reports coefficient estimates and their standard deviations (in parenthesis) for a quarterly four-lag VAR that includes the stochastic discount factor (define as $s^n_{t+n} = \beta^n(c_{t+n}/c_t)^{-\gamma}$) and the inverse of inflation, for the period 1986:q1 to 2006:q1. The VAR is estimated separately for maturities from one to five years.

Table 6: Estimates of inflation risk premium for different maturities in Chile (in %)

Maturity in years (n)	Mean	Std dev	Max	Min	Persistence
	Results for $\gamma = 5$				
1	0.010	0.001	0.138	-0.160	0.174
2	0.004	0.000	0.026	-0.017	-0.309
3	-0.005	0.000	0.012	-0.026	-0.067
4	-0.004	0.000	0.007	-0.021	0.282
5	-0.003	0.000	0.011	-0.019	0.058
	Results for $\gamma = 10$				
1	0.026	0.006	0.252	-0.337	0.138
2	0.009	0.000	0.082	-0.075	-0.061
3	-0.008	0.000	0.018	-0.059	0.332
4	-0.007	0.000	0.007	-0.047	0.427
5	-0.005	0.000	0.031	-0.074	0.225
	Results for $\gamma = 15$				
1	0.051	0.016	0.373	-0.516	0.137
2	0.018	0.003	0.200	-0.252	0.065
3	-0.010	0.001	0.054	-0.147	0.701
4	-0.023	0.015	0.049	-0.923	-0.069
5	-0.014	0.070	0.943	-1.581	-0.288
	Results for $\gamma = 20$				
1	0.095	0.067	1.581	-0.788	0.151
2	0.038	0.014	0.649	-0.485	0.055
3	-0.023	0.070	0.747	-1.880	-0.062
4	-0.041	0.162	0.269	-3.038	-0.070
5	-0.015	0.018	0.149	-0.851	0.205

The table reports quarterly summary statistics of the inflation risk premium for maturities from one to five years, during the period 1989:q1-2006:q3. The CRRA coefficient varies between the values of 5, 10, 15, and 20. Persistence denotes the 4-quarter autocorrelation coefficient.

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