INFLATION RISK PREMIUM IMPLIED BY OPTIONS

EDDY AZOULAY
Bank of Israel

MENACHEM BRENNER*
Stern School of Business
New York University

YORAM LANDSKRONER
College for Academic Studies Or Yehuda
And
School of Business Administration
The Hebrew University of Jerusalem

ROY STEIN
Bank of Israel

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* Corresponding author Menachem Brenner ; e-mail: mbrenner@stern.nyu.edu
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Abstract

One of the commonly used estimates of expected inflation is the yield differential between nominal bonds and inflation-indexed bonds (breakeven inflation). Breakeven inflation is however a biased estimate of expected inflation because it includes an inflation risk premium (IRP). The novelty of our approach is that we estimate the IRP using the volatility implied from foreign exchange (FX) option prices combined with a price of risk extracted from stock prices. Purchasing Power Parity theory provides the linkage between inflation and the foreign exchange rate. Using data from the Israeli government bond market, which has a long history of liquid markets in inflation-linked and nominal bonds as well as an active FX options market, we find a statistically and economically significant positive inflation risk premium.

JEL Classification: E31, E32, E51

Key words: Inflation expectations, inflation-indexed (linked) bonds, Inflation risk premium, foreign exchange options
1. Introduction

Inflation expectations are a key variable for investors in capital markets and also play an important role in determining monetary policy in many countries, especially in countries with strong and independent central banks. In this paper we derive a market-based measure of unbiased inflation expectations, net of inflation risk premium (IRP), using data on inflation indexed government bonds, nominal government bonds and options on foreign exchange (FX) in lieu of options on inflation which are not available.

A number of approaches are used to forecast inflation. Most models are econometric models, both structural and purely statistical. These models, however, rely on historic data and are not forward looking. Another source of inflation forecasts are surveys of professional analysts and economists.¹ Surveys are, however, based on samples that are usually small and therefore might not be representative of market expectations. In economies where inflation-indexed government bonds have been issued (e.g. TIPS in the U.S.) inflation expectations are derived from the yield differential between nominal bonds and inflation – indexed (real) government bonds. This estimate is referred to as breakeven inflation (BEI). ²

Inflation indexed bonds exist now in many countries.³ The BEI as a measure of inflation expectations is used by central banks in a number of countries (e.g., the Federal Reserve, the Bank of England, Bank of Canada and the Bank of Israel)⁴. The advantages of these estimates are that they are market based, forward looking, can be computed continuously and can provide the entire term structure of inflation expectations. Numerous papers have estimated inflation expectations in different countries from nominal and inflation indexed bonds.⁵

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¹ The most popular one is the Federal Reserve Bank of Philadelphia’s quarterly Survey of Professional Forecasters. This survey is based on a group of about 30 forecasters, professionals who work for Wall Street firms and other businesses, who forecast various economic variables.

² There is considerable research on forecasting inflation and economic activity using asset prices. For a review see Stock and Watson, (2003).

³ In the past they have been issued mainly in developing economies with high and uncertain inflation. However, since the 1980’s indexed bonds were mainly issued by developed countries. These include the U.K. (1981), Australia (1985), Canada (1991), Sweden (1994), U.S. (1997) and France (1998).

⁴ In his remarks Ben Bernake (2007), the chairman of the Federal Reserve Board, points out that the Fed is using all three approaches mentioned above, and tops them with expert judgment and out of model information.

⁵ For example, Sack (2000) derives inflation expectations from U.S. Treasury nominal bonds and the inflation indexed bonds (TIPS). Scholtes (2002) outlines the derivation and interpretation of breakeven
Breakeven inflation however is a biased estimate of expected inflation because it includes an inflation risk premium and possibly a liquidity premium. There is a growing body of literature, theoretical and empirical, on the IRP, providing estimates of this premium. These estimates, however, differ in size, maturity structure, volatility and even sign. As pointed out by Hördahl and Tristani (2007) the different results in the literature may be at least partly due to differences in samples (time) and/or country. Campbell and Shiller (1996) in a study that predates the issuing of TIPS in the U.S. use two methods to estimate the IRP from data on nominal bonds based on finance theory. They obtained estimates in the range of 50-150 basis points for a maturity of 5-year bonds. Foresi, Penati and Pennacchi (1996) use two factors to price bonds, expected inflation and the real interest rate. Accordingly they define two risk premia where each risk premium is the product of the market price of risk of the factor multiplied by the risk of that factor. Their estimated excess return IRP for the U.K. varies from 0 for the short run to 55 basis point for the long run; they found a much smaller IRP in Sweden.

In a study of the Israeli bond market Kandel, Ofer and Sarig (1996) reported that the IRP in periods of high inflation was about 34 basis points a month and only 5 basis points in periods of low inflation. Evans (1998) using U.K. index-linked and nominal bonds and survey data on inflation expectations finds a positive and significant time-varying IRP that co-varies positively with the spread between nominal and real yields (BEI). In a later study Evans (2003), using a regime-switching model and index linked bonds data, finds a large and negative inflation risk premium (-1.8% to -3.5% for 10 years horizon). Buraschi and Jilstov (2005) are using a structural monetary version of a real business cycle model to estimate the IRP and find an average IRP of 70 basis points that is time varying, ranging from 20 to 140 basis points. Chen, Liu and Cheng (2005) using a two factor CIR model and data on TIPS and nominal bonds found an inflation risk premium that ranges from -1 to 132 basis points. D’Amico, Kim and Wei (2007) finding indicate the presence of an inflation risk premium as well as a liquidity premium, using TIPS and Survey data of inflation expectations. Hördahl and Tristani (2007) find that on average the inflation from inflation linked gilts in the U.K. Christensen, Dion and Reid (2004) have estimated the breakeven inflation in Canada and found it to be higher on average and more variable than survey measures of inflation expectations. The Bank of Israel (BOI) has derived inflation expectations from the bond market since 1988. This is based on research done by Yariv (1990, 2000).
risk premium in the Euro zone is not significantly different from zero over an EMU sample of nominal and inflation linked bonds. Ang, Bekaert and Wei (2008) using a regime switching affine model and nominal yields estimate the quarterly inflation risk premium to be 31-114 basis points.

In this study we estimate the inflation risk premium (IRP) over time and investigate its properties using Israeli data on nominal bonds, inflation linked bonds and foreign exchange options. The use of FX options in lieu of options on inflation is due to the lack of the latter. Our approach provides unbiased inflation expectations that can be used by investors and monetary policy decision makers. The novelty of the methodology described here is that in estimating the IRP we use the volatility implied in options as a measure of forward looking inflation risk. The second contribution of the paper is that our estimates are based on data from reliable and liquid markets, namely the government bond market in Israel. The Israeli economy is a good candidate for research on inflation expectations since it has a long history of high and volatile inflation and of government policies aimed at conquering the high inflation. It also has a well-functioning capital market with a long history of a liquid inflation linked and nominal government bonds market. Unlike other countries, the U.S. included, where inflation index bonds are a more recent phenomena and thus have low liquidity, inflation index bonds and non indexed bonds, in Israel, have been trading in a very liquid market. Ideally we would have liked to use prices of options on inflation to estimate the implied inflation volatility as a measure of inflation risk. However, in the absence of a liquid market for options on inflation we use foreign exchange (FX) options that are actively traded on an organized exchange. The use of FX options in lieu of inflation options is based on the strong linkage between inflation and foreign exchange rates in countries like Israel. In a small and open economy such as Israel there is a significant relationship (transmission mechanism) between the exchange rate and inflation. In our model the linkage between inflation and the changes in the price of foreign exchange is based on purchasing power

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6 The liquidity in these markets has substantially improved since 2000 following deregulation of the foreign exchange market and the increased investment by foreign investors; it was further enhanced as a result of the introduction of a primary dealer system and the initiation of market makers. The daily average volume of trading of government nominal and inflation indexed bonds was 413 and 246 million Shekels respectively in 2002 and 1877 and 728 million Shekels in 2007.

7 In Israel there is a thinly traded over-the-counter market for options on inflation that is non-transparent and dominated by the banks.
parity between Israel and the U.S. The empirical estimation of short-term inflation is performed using an Error Correction Model (ECM).

The inflation risk premium (IRP) is a function of two factors, inflation risk and the so-called “market price of risk” (MPR) that reflects investors risk aversion. The MPR should be the same across different financial assets. Using stock market returns from the Tel-Aviv stock exchange (TASE) we estimate the market price of risk (MPR).\(^8\) We subtract the IRP from the BEI, to get a pure estimate of the expected inflation (defined as the expected change in the consumer price index (CPI)).\(^9\)

Our main findings are that there exists a non-trivial positive inflation risk premium that varies over time. The inflation risk premium expected a year hence was, on the average, about 25 basis points during the years 2002-2007. This premium accounted for about 15% of the difference between the nominal and real rates with a one-year maturity (the standard deviation was about 5%). This falls in the lower part of the range of results obtained in other studies. We have also estimated the liquidity premium and found it to be of negligible size. Also, subtracting this premium from the ‘break-even’ expectations provides inflation expectations that are closer to the realized inflation in the 2002-2007 period. We could not obtain a forward looking IRP after 2007 since the relatively strong linkage between inflation and the foreign exchange rates ceased to exist after the Bank of Israel started to intervene in the foreign exchange market (March 2008) and the transition from quoting many transactions in dollars to quoting them in Shekels (e.g. housing prices).\(^{10}\)

Our findings lend support to the conjecture that the break-even inflation expectations derived from the bond market provide an upward biased estimate of expected inflation. Investors and policy-makers should therefore take into account the risk premium embedded in this estimate. It is important to implement such a procedure especially in a period of high and volatile inflation when the central bank needs reliable and precise inputs for monetary policy decisions.

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8 Chowdhry, Roll and Xia (2005) use stock returns to extract estimates of realized pure inflation. They purge stock returns from the risk premium of the different economic factors by using the Fama-French three-factor model.

9 To simplify the analysis we disregard a convexity term, which is negligible in size.

10 This, of course, affected the value of the FX options, which we cannot use in deriving our market based estimates.
The methodology presented in this paper can be used to derive an IRP and inflation expectations in countries that have both inflation-linked and nominal government bonds, especially in open economies that are prone to high inflation-uncertainty due also to the linkage between the exchange rate and inflation.

2. Purchasing Power Parity; Methodology and Estimation

A basic premise of the analysis is that in a small open economy, like Israel, there is a strong link between exchange rates and inflation. In such economies changes in the exchange rate translate into changes in the general price level, including prices of locally produced goods and services which are affected by changes in input prices (oil, for example). To establish the transmission mechanism between foreign exchange rates and inflation we start with an analysis of purchasing power parity in the long run. We then estimate the short-run inflation relationship using an Error Correction Model. This enables us to arrive at an ex-ante estimate of inflation risk and the inflation risk premium. In our analysis we use implied volatility as an estimate of expected inflation volatility. In the absence of options on inflation we use implied volatility from currency options as an estimate of the expected inflation volatility. Also, the MPR, in real terms, is obtained from the Israeli stock market. Using an arbitrage argument, the law of one price, we assume that the MPR is the same for all sources of uncertainty including inflation uncertainty.

To establish the transmission mechanism between foreign exchange rates and inflation we start with an analysis of purchasing power parity in the long run. We then estimate the short-run inflation relationship using an Error Correction Model. This enables us to arrive at an ex-ante estimate of inflation risk and the inflation risk premium.

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11 The effect of the exchange rate on consumer prices in Israel was investigated and reported in many studies: Bruno and Sussman (1979), Azoulay and Elkayam (2001), to mention just a few. Azoulay and Elkayam (2001) examined the effect of monetary policy on inflation in Israel and found that devaluation of the currency coupled with worldwide inflation has a significant effect on domestic inflation.
2a. Purchasing Power Parity (PPP) in the Long Run

PPP simply says that the exchange rate reflects the relative price levels, of goods and services, of two countries. By and large, empirical studies have rejected PPP in the short run. However, several researchers have found that it holds in the long run. Rogoff (1996) states that there is a consensus that, in the long run, the real exchange rate approaches PPP. Cheung and Lai (1993) use co-integration tests to show that there is a stable long-run relationship between exchange rates and consumer prices. In this study we also use co-integration tests to examine the long-run relationship in Israel after it moved to a fully floating exchange rate regime in May 1997.

We start with the simple model of absolute PPP, between the New Israeli Shekel/U.S. dollar (ILS/USD) exchange rate and relative consumer prices in Israel and the U.S.:

Price level (ISR) = Exchange rate (ILS/USD) * Price level (U.S.)

We transform this equation into logarithmic form and obtain the following estimation equation

\[ P_t(\text{ISR}) = \beta_0 + \beta_1 S_t + \beta_2 P_t(\text{US}) + v_t \]  

(1)

Where \( P_t(\text{ISR}) \) is log of the price level (CPI) in Israel, \( P_t(\text{US}) \) is the log of the consumer price index (CPI) in the U.S., \( S_t \) is log of the ILS/USD exchange rate and \( v_t \) is the error term.

If absolute PPP holds local prices change with a change in either the exchange rate or as a result of a change in the foreign country’s prices. Thus the absolute PPP hypothesis can be stated as: \( H_0: \beta_1=1 \) and \( \beta_2=1 \). Using monthly observations, for the period May 1997 – June 2007 we obtained the following regression results.

\[ P_t(\text{ISR}) = 1.37 + 0.36 S_t + 0.5 P_t(\text{US}) + v_t \]  

(1a)

\( t \) statistics: \( (20.5) \) \( (26.7) \) \( (33.4) \)

\( R^2 = 0.98 \), \( DW = 0.32 \), \( N = 122 \)

(\( t \) statistics are in parentheses).
The coefficients $\beta_1$ and $\beta_2$ are significantly different from 1 (using the Johansen test). Thus, the null hypothesis $H_0$, is rejected. This result, however, comes at no surprise and is consistent with most studies, which have tested PPP in other countries.

Three main reasons are given for empirical results that reject the existence of absolute PPP, these reasons are also relevant in the context of Israel. First, the CPI includes non-tradable assets, housing prices for example, which adjust infrequently. Second, about 20 percent of the tradable items included in the CPI in Israel are affected by changes in the European currency, the Euro, and not the U.S. dollar. Third, the sampling period, which started immediately after the move from a band-controlled FX regime to a free floating one, is not long enough to test such a relationship as the one tested above. Moreover, within the sample period there was a recession (2002 to 2003), when producers could not afford to adjust prices upwards.

We have also tested (1a) in the extended period, February 1997- August 2010. We effectively divided the entire period into two sub-periods (the second one is from July 2007 to August 2010) using dummy variables in the estimation. As was expected, in the second period the link between inflation and foreign exchange rates has evaporated. The main reason is the change in the nonintervention policy of the central bank. Starting in March 2008 until August 2010 the Bank of Israel intervened almost daily in the foreign exchange market, effectively reducing the volatility of FX. In addition, since the end of 2007 there have been structural changes in the Israeli economy that affected the relationship between inflation in Israel and the exchange rate$^{12}$.

Though absolute PPP was rejected, we turned to tests of non-stationarity and co-integration, as was done for other countries, to see if a long-run relationship between consumer prices and exchange rates does exist. For the three variables, $P_t$ (ISR), $S$, and $P_t$(US)), to be co-integrated we need the following two conditions to hold: (a) at least two variables exhibit non-stationarity of the same order, and (b) the three variables exhibit at least one co-integration relationship.

We first use the Augmented Dickey-Fuller Test (ADF) to test for non-stationarity of the variables; we use a constant term and two lags.

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$^{12}$ For example, the common practice of quoting prices of real estate in dollars has changed to quoting prices in Shekels (the Israeli currency).
Table 1a: A Unit Root Test*

<table>
<thead>
<tr>
<th>VARIABLES</th>
<th>ADF</th>
<th>Significance Level 1%</th>
<th>Significance Level 5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>Levels:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$S_t$</td>
<td>2.31</td>
<td>3.49</td>
<td>2.89</td>
</tr>
<tr>
<td>$P_t$(US)</td>
<td>0.86</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$P_t$(ISR)</td>
<td>1.98</td>
<td>&quot;</td>
<td>&quot;</td>
</tr>
<tr>
<td>First differences: $\Delta (S_t)$</td>
<td>8.13</td>
<td>3.49</td>
<td>2.89</td>
</tr>
<tr>
<td>$\Delta P_t$(US)</td>
<td>-8.80</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\Delta P_t$(ISR)</td>
<td>7.06</td>
<td>&quot;</td>
<td>&quot;</td>
</tr>
</tbody>
</table>

* The unit root test is applied to the log exchange rate ILS/USD plus the log of the U.S. CPI (level) and to the log of the Israeli consumer price index (level). It is also applied to the first differences of the log price levels.

The results in Table 1a indicate that we cannot reject the hypothesis of a unit root in the level (in log form) variables, as the variables are non-stationary. When the test is applied to first differences, the rate of change of the exchange rate ($\Delta S$) and the inflation rates ($\Delta P$) in Israel and the U.S., we reject the existence of a unit root. In other words, the time series of first differences are stationary and integrated in the first order. These results are consistent with the findings in other developed countries (see, for example, Cheung and Lai (1993) and Corbae and Ouliaris (1988)). The next step is to test for co-integration using the approaches of Johansen and Engle and Granger. The purpose of the analysis is to see whether the results in equation (1a) represent a long-run relationship, which will assist us in understanding the short-run dynamics of inflation in Israel.

According to Engle and Granger (1991) a necessary condition for co-integration is that the error term is stationary. An ADF test of the error terms $\nu_t$ shows that the series is stationary and we can thus reject the hypothesis of a unit root at the 5% level and conclude that the variables are co-integrated. The Johansen co-integration test was applied to lags of 2, 4 and 7. In Table 1b we present the co-integration coefficients of the
long run, for each lag. The results show that there is at least one co-integration relationship, at the 1% level\textsuperscript{13}.

These results are consistent with the Engle and Granger test results that there is a long run relationship between the variables of equation (1a), consumer prices in Israel, consumer prices in the U.S. and the exchange rate.

Table 1b: Tests of Co-integration\textsuperscript{1}

<table>
<thead>
<tr>
<th>1) Engle and Granger test</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Error Term</td>
<td>ADF</td>
<td>Critical val.</td>
<td></td>
</tr>
<tr>
<td>(v_t)</td>
<td>3.9*</td>
<td>-3.48</td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>2) Johansen test</th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Max. Lags</td>
<td>(\beta_0)</td>
<td>(\beta_1)</td>
<td>(\beta_2)</td>
</tr>
<tr>
<td>2</td>
<td>2.31</td>
<td>0.40</td>
<td>0.28</td>
</tr>
<tr>
<td>4</td>
<td>1.46</td>
<td>0.39</td>
<td>1.30</td>
</tr>
<tr>
<td>7</td>
<td>2.01</td>
<td>0.34</td>
<td>0.34</td>
</tr>
</tbody>
</table>

* Significant at the 1% level
** Significant at the 5% level

1) The Engle and Granger test is applied to the error term \(v\) from equation 1b and it is significant at the 1% level. The Johansen test indicates that two co-integrating equations found at the 5% level and therefore it confirms the results of the Engle and Granger test of co-integration.

2b. PPP in the short run: The Error Correction Model (ECM)

The co-integration tests which indicate a long-term relationship enable us to examine the short-term behavior of these variables. Engle and Granger have suggested that for variables that are co-integrated it should be possible to find a process that is Error Correcting. This is a process that describes the convergence of the short-term deviations to the long-run relationship. Basically, the long run and short run come together with the inclusion of the lagged error term from equation (1a) in the short-term equation. In

\textsuperscript{13} There is a direct relationship between DW statistic and the stationarity of the residual. Despite the fact that the DW statistic is small, the residuals were found to be stationary due to the fact that the standard deviation of the error term is very small.
equation (2) we specify the short run behavior and the convergence process by an Error Correction Model (ECM).

\[
\Delta P_t(\text{ISR}) = \theta_1 \cdot \Delta S_t + \theta_2 \cdot \Delta P_t(\text{US}) + \theta_3 \cdot \Delta P_{t-1}(\text{ISR}) + \theta_4 \cdot EC_{t-1} + x_t \tag{2}
\]

where \(\Delta P(.)\) is the inflation rate, \(\Delta S\) is the change in the exchange rate, \(EC_{t-1}\) is the error correction component derived from equation (1a), and \(x_t\) is the error term. The basic idea here is that any short-term deviation from the long-run relationship would be reversed so there is a convergence in the long run. Thus, the coefficient \(\theta_2\), which is an estimate of the speed of convergence, should be negative and significantly different from zero.

Equation (2) was estimated for the period July 1997 to June 2007 using Two Stage Least Squares (TSLS) procedure to account for a possible endogenous effect and with the inflation rates in Israel and the US adjusted for seasonality\(^{14}\). The results are given in equation (2a).

\[
\Delta P_t(\text{ISR}) = 0.17 \Delta S_t + 0.29 \Delta P_t(\text{US}) + 0.3 \Delta P_{t-1}(\text{ISR}) - 0.17 EC_{t-1} + x_t \tag{2a}
\]

\[
(9.4) \quad (3.3) \quad (5.1) \quad (-4.4)
\]

\(R^2 = 0.61\) \quad \text{DW} = 2.25 \quad N = 120

(t statistics are in parentheses).

Equation (2a) is well specified, as evidenced by the \(R^2\) and by the DW statistic. Foreign inflation (U.S.), the exchange rate and lagged domestic inflation explain most of the variation in current domestic inflation. The ECM seems to work well, the Error Correction component is negative and significantly different from zero; about 17 percent of the deviation from the long run relationship is corrected in the following month.

At this point we would like to elaborate on the transmission process from changes in the exchange rate to changes in consumer prices. The immediate transmission channel is the prices of imported goods, which also affect the prices of domestic substitutes. The

\(^{14}\) The instrumental variables were: the one month Dollar LIBOR interest rate, the Euro/ Dollar exchange rate, the US inflation rate and its lags, in addition to the one period lag of the error correction term.
other channel is the prices of imported raw materials and services used in the production of domestic goods. The effect of these price increases on the CPI will depend on their weight in the consumer’s basket.

The transmission coefficient found here is similar to findings in other countries. Gagnon and Ihrig (2002) examined a sample of industrialized countries and found that during the period 1972 to 2000 the one-year transmission coefficient was on the average about 20 percent. Canada, for example, had a 20 percent transmission coefficient. By the end of the above period this coefficient was only 5 percent despite the fact that world trade had increased markedly and there were more imported goods in every consumer’s basket. The increase in imported goods, however, came along with lower prices due to a reduction in import taxes, cheap goods from the emerging markets and credible monetary policies in the developed countries. Another study, Elkayam (2001), examined the 1992-2000 period in Israel and obtained an estimate of 0.19, which is virtually identical to ours.

3. Estimating Inflation Volatility and the IRP

3a. The Volatility of Inflation

We use equation (2b) to derive the relationship between the implied volatility (variance) of the exchange rate and the volatility of inflation. Rearranging terms and taking the variance of (2b) yields:

\[
\text{var} (\Delta P_t(\text{ISR}) - 0.3\Delta P_{t-1}(\text{ISR})) = \text{var}(0.17\Delta S_t + 0.29 \Delta P_t(\text{US}) + (-0.17)E_{C_{t-1}} + x_t) \tag{3}
\]

---

15 See Bailliu and Bouakez (2004) for a discussion on the link between the decline in exchange rate pass-through and the low inflation rate achieved in the last decade in most industrialized economies.

16 Azoulay and Elkayam (2001) examined the period 1988-1996 and obtained a higher coefficient, 0.29, which again points to the changes that occurred in the Israeli economy during the 1990s.

17 An alternative approach to estimating inflation volatility is using a GARCH model. We have not used it since it is not forward looking.
We can expand (3) as follows taking into account the correlations between the variables:

\[
(1 - 2 \cdot 0.3 \cdot \rho_{\Delta P} \cdot \sigma_{\Delta P}^2 + 0.3^2) \cdot \sigma_{\Delta P}^2 = 0.17^2 \cdot \sigma_{\Delta S}^2 + 0.29^2 \cdot \sigma_{\Delta P}^2 \\
+ (-0.17)^2 \cdot \sigma_{\Delta P}^2 + \sigma_{\Delta S}^2 + 2 \cdot 0.17 \cdot 0.29 \cdot \sigma_{\Delta S} \cdot \rho_{\Delta P} \cdot \rho_{\Delta P} \cdot \rho_{\Delta P} \\
+ 2 \cdot 0.29 \cdot (-0.17) \cdot \sigma_{\Delta S} \cdot \rho_{\Delta P} \cdot \rho_{\Delta P} + 2 \cdot 0.17 \cdot \sigma_{\Delta S} \cdot \sigma_{\Delta S} \cdot \rho_{\Delta S} \cdot \rho_{\Delta S} \\
+ 2 \cdot (-0.17) \cdot \sigma_{\Delta S} \cdot \sigma_{\Delta S} \cdot \rho_{\Delta S} \cdot \rho_{\Delta S} \\
+ 2 \cdot (0.17) \cdot \sigma_{\Delta S} \cdot \sigma_{\Delta S} \cdot \rho_{\Delta S} \cdot \rho_{\Delta S} \cdot \rho_{\Delta S}
\] (3a)

Where \( \sigma^2 \) and \( \rho \) denote the variances and correlation terms respectively.

In order to estimate the LHS of 3a we have estimated the serial correlation of Israel's inflation rate and found the first and second order serial correlation to be respectively \( \rho_{\Delta P} = 0.15 \) and \( \rho_{\Delta P} = 0.13 \), the higher order serial correlations were not significantly different from zero. For the RHS of (3a) we have estimated standard deviations and correlation coefficients of the exchange rate changes, the U.S. inflation and the error correction term for the period July 1997 - June 2007 (see Table 1c). These inputs were substituted in equation (3a) obtaining the following expression for the "model" variance of inflation in Israel:

\[
\sigma_{\Delta P}^2 = 0.0344 \sigma_{\Delta S}^2 + 1.23 \cdot 10^{-4} \cdot \sigma_{\Delta S} + 1.92 \cdot 10^{-5}
\] (3b)

where the implied exchange rate volatilities were converted from annual terms to monthly term to coincide with the monthly frequency of the inflation data. This equation links the volatility of inflation to the volatility of the exchange rate and therefore it enables us to use the implied volatility of the exchange rate obtained from options prices to estimate the forward looking ("implied") volatility of inflation. 18

We then converted the "model" inflation volatility that was estimated using monthly data to annual terms. Since inflation series are not a random-walk, having significant serial

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18 The implied FX volatilities (IV) were computed daily, from options with one month to maturity; however we are interested in obtaining forward looking one year estimates of inflation volatility. Since there are no one year options, which are traded on the exchange, we have converted the IV to annual terms.
correlation, we estimated the relationship between the monthly and the annual volatility empirically using two methods, as follows: (1) using an OLS regression with the variance of annual inflation as the dependent variable and the variance of monthly inflation as the independent variable, using monthly observations. The coefficient of the dependent variable was in the range of 20-30 depending on the sample. Thus the variance of annual inflation is 20-30 times the variance of the monthly inflation. In the following analysis we chose the mid-point 25. (2) using direct estimation, where the volatility of annual inflation equals the volatility of the product of 12 consecutive monthly inflation rates from the month in \((t-11)\) to month \((t)\). This method is described next.

\[
Var \left( \frac{P_t}{P_{t-12}} \right) = Var \left( \frac{P_t}{P_{t-1}} \times \frac{P_{t-1}}{P_{t-2}} \times \ldots \times \frac{P_{t-11}}{P_{t-12}} \right) \\
= 12 \left( \sigma_{\Delta P}^{mh} \right)^2 + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t,t-1} + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t,t-2} + \ldots + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t,t-12} \\
+ \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-1,t-1} + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-1,t-2} + \ldots + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-1,t-12} \\
+ \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-12,t-1} + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-12,t-2} + \ldots + \left( \sigma_{\Delta P}^{mh} \right)^2 \times \rho_{t-12,t-12} \\
\tag{3c}
\]

Where \(\left( \sigma_{\Delta P}^{mh} \right)^2\) and \(\rho_{t,t-i}\) are the historic variance of inflation in monthly terms and autocorrelation between monthly inflation at \(t\) and monthly inflation at \(t-i\), respectively. It turns out that only the first order serial correlation coefficients are significantly different from zero.

The estimated first order autocorrelation is 0.5. Thus, equation (3c) can be rewritten as:

\[
Var \left( \frac{P_t}{P_{t-12}} \right) = Var \left( \frac{P_t}{P_{t-1}} \times \frac{P_{t-1}}{P_{t-2}} \times \ldots \times \frac{P_{t-11}}{P_{t-12}} \right) = 12 \left( \sigma_{\Delta P}^{mh} \right)^2 + 24 \left( \sigma_{\Delta P}^{mh} \right)^2 \times 0.5 \\
= 24 \left( \sigma_{\Delta P}^{mh} \right)^2 \tag{3d}
\]

This result of the relationship between annual and monthly variance of inflation is similar to the one obtained by the OLS regression (method 1), 24 vs. 25. Therefore the conversion from the monthly model volatility in equation (3b) to annual terms according to the following relationship derived from historic volatility as outlined above:
\[ \sigma^A \nu = \sqrt{25} \sigma^M \nu \]  

where \( \sigma^A \nu \) and \( \sigma^M \nu \) are the "model" standard deviation of inflation in annual and monthly terms respectively. Since equation (3e) is using standard deviations the factor is the square root of 25.

Table 1c: Standard Deviations and Correlation of the Exchange Rate, the US inflation rate and the Error Correction Term*

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<tr>
<th></th>
<th>( \Delta S_t )</th>
<th>( \Delta P_{t,US} )</th>
<th>( EC_{t-1} )</th>
<th>( x_t )</th>
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\( \Delta S_t \) is the first difference of the log exchange rate LIS/USD, \( \Delta P_{t,US} \) is the US inflation rate, \( EC_{t-1} \) is the error correction term and \( x_t \) is the error term in equation (2) using monthly observations from July 1997 to June 2007.

3b. The Market Price of Risk and the IRP

We now turn to the estimation of the second component of the IRP: the market price of risk (MPR).

In general equilibrium, the fact that all consumers hold risky assets in the same proportion implies that risk premia are determined by the standard CAPM. 19. The use of the one factor CAPM as an equilibrium model in the Israeli market is justified by

19 Though the consumption CAPM is suggested as an alternative to the traditional CAPM, there is extensive empirical evidence that the traditional CAPM outperforms the consumption-based CAPM in terms of predicting asset risk premia. The poor empirical performance of the consumption-beta model is an indication that, in practice, we cannot infer the marginal utility of nondurable consumption with sufficient accuracy (see Flavin and Nakagawa (2008)).
empirical evidence that validates the model. These results differ from those obtained in the US and other large markets. One of the explanations is that due to the small size of the market the crucial assumption that investors hold the market portfolio is closer to reality in Israel than in larger markets like the U.S. and others (see Levy (1980)). We define the MPR in real terms (see Siegel and Warner (1977)):

$$MPR = \left[ (\bar{R}_m - \bar{\Delta P}) - (\bar{R}_f - \bar{\Delta P}) \right] \sigma^2(\bar{R}_m - \bar{\Delta P})$$  \hspace{1cm} (4)

In our study we estimate the MPR using the Israeli stock market data, where $R_m$ is the average nominal return on the TA100, an index of the largest 100 companies on the Tel Aviv Stock Exchange; $R_f$ is the nominal risk free rate, using the yields to maturity on Israeli Treasury bills for one month. The sample period was May 1997- June 2007. Table 1d presents the parameter estimates used in the computation of the MPR.

### Table 1d: Parameter Estimates (percentage points, in monthly terms)*

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<th>$\bar{R}_m$</th>
<th>$\bar{R}_f$</th>
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* $R_m$ is the return on the market portfolio, $R_f$ is the nominal risk free rate, $\sigma(\bar{R}_m - \bar{\Delta P})$ is the standard deviation of the market portfolio in real terms, $\bar{\Delta P}$ is the average rate of inflation (CPI).

The IRP is the product of the MPR and the risk of inflation measured by the implied standard deviation of inflation, extracted from FX options (see 3b).

$$IRP_t = \sigma^2_{\Delta \bar{P}} \cdot MPR$$  \hspace{1cm} (4a)

where $\sigma^2_{\Delta \bar{P}}$ is the annual implied volatility of inflation.

---

20 For the market price of risk in nominal terms under uncertain inflation, see Friend, Landskroner and Losq (1976).
Estimation of the IRP enables us to extract the inflation expectations from the yield on nominal bonds minus the yield on the CPI linked bonds, the so-called real bonds.

\[ E(\Delta P_t) = (RN_t - RP_t) - IRP_t, \]  

(4b)

Where \( E(\Delta P_t) \) is the pure expected inflation at time \( t \), \( RN \) is the yield on nominal bonds and \( RP \) is the yield on the indexed ("real") bond. These are one-year forward looking expectations estimated on a daily basis. Currently, central banks (e.g., the BOI) use break-even inflation \( (RN_t - RP_t) \) as an estimate of expected inflation.

**INSERT TABLE 2**

In table 2 we present our results for the period October 2002 to June 2007. This is the period in which a liquid market in FX options existed in Israel. We present estimates of: Actual inflation, BEI, implied FX volatility, “implied” inflation volatility, the inflation risk premium, inflation expectations net of the IRP. We compare these estimates to the breakeven inflation. The results in these tables show that actual inflation was on the average slightly negative while the volatility was high, the implied volatility of the FX options exhibits a downward trend but the volatility of implied volatility has increased. The derived (“implied”) volatility of inflation exhibits a similar trend. This also shows up in the estimated IRP.

Our findings point to a sizable IRP, about 25 basis points on the average, with a standard deviation of 7 basis points. This is at the lower end of the range of IRP found in other countries. The IRP accounted, on average, for 15 percent of the break-even inflation. However, the proportion of IRP in break-even inflation is rather volatile. It has a standard deviation of 6.02%.

To complete the analysis we estimated also the liquidity premium. We have estimated the Amihud (2002) measure of liquidity, ILLIQ, for indexed and non-indexed bonds, their averages (in percent) were 0.0002 and 0.0001 respectively. Using Amihud (2002) factor loadings of 0.162 and 0.112 yields liquidity premia that are negligible in size.
In annual terms inflation expectations during that period were on average about 1.61 percent with a standard deviation of 0.62 percent. This compares to a small average actual inflation of 0.57% percent and a much higher standard deviation of 5.33 percent. It should be noted that the adjustment for the implied volatility term structure has a minor effect on the IRP.

**INSERT TABLE 3**

In Table 3 we present the inflation expectations from two sources, the CPI forecasts of analysts and the break-even inflation from the capital market. Interestingly the forecasts of the analysts are consistently higher than the break-even inflation despite the fact that they don’t include an inflation risk premium while the latter does. We also present the inflation risk premium we have estimated, which is a measure of inflation risk and compare these estimates to the dispersion of the analysts’ forecasts (range and standard deviation) which can also be considered a measure of inflation uncertainty. The positive correlation that we found in the monthly observations (of 12 months forecasts) for the period 2002-2007, was 0.4, between the range of the forecasts and IRP and 0.42 between the standard deviation of the forecasts and the IRP, supports the validity of our estimates of the IRP.

**4. Summary and Conclusions**

Central banks, financial institutions and other investors increasingly use forward looking financial market data to obtain unbiased expectations of future inflation. The standard approach has been to subtract the yield on a real bond, a CPI linked bond, from a nominal bond. This difference, termed break-even inflation (BEI), is an upward biased estimate since it includes an inflation risk premium (IRP).

The objective of this paper is to provide a methodology that derives estimates of inflation risk premiums and enables forecasters to extract unbiased inflation expectations from financial market data. We subtract an estimate of the IRP from the biased BEI estimate to obtain unbiased inflation expectations.

We found that the IRP for a year ahead was a sizeable 25 basis points during the estimation period, 2002-2007, and it accounted for 15% of the difference between nominal and real yields. Thus it should be taken into account when estimating inflation
expectations from capital markets data. Another empirical observation that supports our findings of a positive IRP is the positive gap between breakeven inflation and the actual inflation.

REFERENCES


Bernanke B. S. 2007," Inflation Expectations and Inflation Forecasting” Remarks at the Monetary Economic Workshop of the NBER. July 10


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<th>Month</th>
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<th>Implied S.D. of NIS/$ exchange rate</th>
<th>Annual inflation S.D.</th>
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**S.D.**

<table>
<thead>
<tr>
<th></th>
<th>2006</th>
<th>2007</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1.80</td>
<td>0.59</td>
</tr>
</tbody>
</table>

(1) Actual monthly inflation rate in annual terms.

(2) The difference between the nominal yield on one-year Treasury bills and the real yield on a CPI-indexed bond with a maturity of approximately one year. It is calculated using the methodology of the BOI Monetary Department. See Yariv (2000).

(3) Implied volatility from approximately one-month NIS/$ options traded on the Tel Aviv Stock Exchange, Calculation made by BOI Monetary Department.

(4) Calculated on the basis of Equation 3b from daily data. We substitute into Equation 3b the daily figure from Column (2) in monthly terms. To obtain the figure in annual terms we multiplied the result by the square root of 25 as in (3e).

(5) Based on Equation 3e. The IRP is the product of the market price of risk and the daily volatility of inflation, from Column 3. The market price of risk is constant and calculated using the definition in 4. The estimated market price of risk is 2.4, calculated from a sample of monthly averages from May 1997-June 2007, based on the yield of a market portfolio on the Tel Aviv Stock Exchange and the YTM on one
month Israeli's Treasury Bills, which represents the yield on a risk-free asset.

(6) Inflation expectations net of risk premium, calculated on the basis of daily data-as the difference between column 2 and column 5.

(7) Proportion of risk premium in break-even inflation, column 5 divided by column 2.

Table 3
Inflation Expectation, CPI Forecasts, and theirs Uncertainties
(percent, monthly observations in annual average)

<table>
<thead>
<tr>
<th>Year</th>
<th>CPI Forecasts Average</th>
<th>Inflation Expectations</th>
<th>STD</th>
<th>Range</th>
<th>Inflation risk premium</th>
</tr>
</thead>
<tbody>
<tr>
<td>2002</td>
<td>2.6</td>
<td>0.553</td>
<td>1.414</td>
<td>0.389</td>
<td></td>
</tr>
<tr>
<td>2003</td>
<td>2.0</td>
<td>1.6</td>
<td>0.461</td>
<td>1.296</td>
<td>0.338</td>
</tr>
<tr>
<td>2004</td>
<td>2.1</td>
<td>1.4</td>
<td>0.414</td>
<td>1.234</td>
<td>0.198</td>
</tr>
<tr>
<td>2005</td>
<td>2.1</td>
<td>1.8</td>
<td>0.308</td>
<td>0.837</td>
<td>0.201</td>
</tr>
<tr>
<td>2006</td>
<td>1.9</td>
<td>1.6</td>
<td>0.332</td>
<td>1.019</td>
<td>0.225</td>
</tr>
<tr>
<td>I/2007</td>
<td>1.8</td>
<td>0.9</td>
<td>0.335</td>
<td>1.002</td>
<td>0.243</td>
</tr>
</tbody>
</table>

(1) Average of the private sector forecasters. The number of forecasters increased, during the sample period, from 7 to 12.

(2) Inflation Expectations are computed as the difference between break-even inflation and the Inflation risk premium, as shown in Column 6 of Table 2.

(3) Standard deviation of the forecasts provided by the private sector forecasters.

(4) Range of the forecasts; Maximum-Minimum

(5) The Inflation Risk Premium (IRP) is computed using Equation 3e. The IRP is the product of the market price of risk and the daily volatility of inflation, from Column 3 in Table 2. The market price of risk is constant and calculated using the definition in 4. The estimated market price of risk is 2.4. It is calculated from a sample of monthly averages from May 1997-June 2007, based on the rate of return of a market portfolio on the Tel Aviv Stock Exchange and the YTM on one month Israeli's Treasury Bills as the proxy of the risk-free rate.