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Nominal Anchor: *Economic and Statistical Aspects*

by

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ABSTRACT

In this paper we give a concrete meaning to the concept of a nominal anchor, based on the Israeli inflationary experience. The lack of a nominal anchor in the inflationary era is related to the aversion of the policymakers to *changes* in inflation, without regard to the absolute level of the inflation plateau. The incentive to use surprise-inflation tactics in this environment, led to an acceleration of inflation. This contrasts with the early period of the 1950's and 1960's where the Bretton-Woods system provided a nominal anchor and hence a concern with the *level* of inflation. A similar attitude prevailed after the 1985 stabilization, particularly since Israel adopted the inflation target regime. We test this hypothesis by relating the lack of a nominal anchor in the inflationary era to a random walk (unit root) process in the inflation series. This hypothesis should be rejected in the stable inflation periods. The results confirm these implications.

> "theory is to be judged by its predictive power for the class of phenomena which it is intended to 'explain'. Only factual evidence can show whether it is …tentatively 'accepted' as valid or 'rejected'… the only relevant test of the *validity* of a hypothesis is comparison of its predictions with experience".
>
> Milton Friedman in *Essays in Positive Economics.*

I. Introduction

In the fifties and sixties of the twentieth century the world enjoyed price stability, that was mainly associated with the exchange rate system of Bretton-Woods. In spite of its many imperfections this system provided a nominal anchor for the member countries of the IMF. After the collapse of this system in the early seventies inflation surged in the world, especially in the so-called chronic inflation countries, as a result of the loss of the nominal anchor and the energy crises. However, inflation was brought down drastically in the past twenty years and now the world is again in an environment of relative price stability, governed not by a fixed exchange rate, but rather by monetary rules. This paper tries to understand the inflation cycle using the Israeli experience.

It goes without saying that real factors played a role in the surge in inflation and its subsequent decline. In Israel the fiscal deficit reached 17% of GDP in the inflationary era of 1973-85 and the public debt doubled in this period from 80% of GDP in the beginning of the inflationary period the 160% at its end. There are some economists that hold the view that these "fundamentals" alone explain the inflationary process. Yet the Israeli stabilization program (Bruno (1993)) specified a nominal anchor (the exchange rate), indicating that the fiscal aspect, with all its significance, did not provide a sufficient anchor for inflation. Similarly, the Maasticht criterion did not confine itself to the fiscal indicators, but stated the monetary target of low inflation.

Substantial progress in the explanation of the role of monetary policy in the transition to low inflation has been provided by the studies of the change in central bank independence (see A. Cukierman (1992), (2005)). Additional progress was made by the recent literature on the change of monetary policy rules in the disinflation process; for example, the change in the Taylor rules for setting the interest rate by the central banks was important in securing price stability (Taylor (1993), Clarida et. al. (1999) and Woodford (2003)).

In this paper we propose an entirely different criterion for the differentiation between the monetary regime in the inflationary period and those of low inflation Our main point is that in the inflationary period the monetary regime lacked a nominal anchor, while in the other periods it had one We show that absence nominal anchor (inflation target) in the policymakers' objective function can provide an explanation to the inflationary episode, while the adherence to the Bretton-Woods rules or the introduction of an inflation targets regime were conducive to price stability.

Our starting point is that in the inflationary regime in Israel, the authorities adopted the approach that the negative effect of inflation on the real sector can be neutralized by indexation. Indexation can indeed cope with very high inflation levels, but it is relatively ineffective in handling acceleration of inflation. This motivates the concern of the authorities with the *change* in inflation. Accordingly we adopt a version of the Kydland and Prescott (1977) and Barro and Gordon (1983) models for the inflationary period with the important modification, that instead of being concerned with the inflation *level* the policymaker is concerned only with the *change* in that level. For the Bretton-Woods period and that after the 1985 stabilization we assume that the policymaker was concerned also with the level of inflation. We retain the assumption of these models that the policymaker has an incentive to employ surprise-inflation tactics because they may give him real advantages (such as the erosion of real wages or the reduction in the fiscal deficit by raising public sector prices), although these advantages were suppressed in the low inflation periods.

Indirect evidence for our basic approach is the fact that the Israeli inflationary process in 1973–85 got stuck a number of times on various levels (see Liviatan and Melnick (1998), and Bruno and Fischer (1985)), which the policymakers were content with, but were seemed to be troubled by a *change* in the inflation rate. The main concern was that such changes may upset the inflation indexation arrangements that existed at the time. We take this feature as a basis for developing the concept of a nominal anchor. Although the existence of the "inflation steps" is not essential for the explanation of the rising inflation trend, the model has to be consistent with them. We explain this connection when we deal with the statistical framework of out model.

Policy game models can provide some insight for the case of a policymaker who tries to preserve the current inflation plateau *without regard to its absolute level*. As noted, in our modification of the Barro-Gordon model, the policymaker is averse to *changes* in the inflation plateau, but benefits from a surprise change in inflation[1]. Subject to this policy setup, the discretionary policymaker will be unable to maintain the current inflation plateau in a policy game equilibrium[2]. The equilibrium outcome will include a persistent rise in inflation. *This is in sharp contrast to Barro-Gordon models, which explain only the level of inflation but not its acceleration over time*. This setup gives a concrete meaning to the absence of a nominal anchor. By contrast, the existence of a nominal anchor corresponds to the case where the policymaker cares not only about the *change* in inflation but also about its *absolute level*.

This approach also enables us to formulate a statistical test of the existence, or lack of, a nominal anchor. For the purpose of statistical testing of the latter hypothesis we make use of the fact that *absence of a nominal anchor* entails *a unit root* in the series of inflation

[1] Especially in the case of balance of payments crises which require big devaluations, as was the Israeli case.
[2] Thus, for example, to the situation during the crawling peg period 1975-77 in Israel where the government accepted the initial inflation plateau, of some 40% annually, which emerged from the 1973 crisis, but could not prevent the further acceleration of inflation in later years.

(since the policymaker's loss is related only to the *change* in inflation, which implies that the current inflation is related to past inflation with a coefficient of one). This corresponds to a random walk with a positive drift in the inflation series ("positive" - since the policymaker can benefit from surprise inflation). The drift may even increase over time both because the incentive to generate inflation surprises is greater at higher inflation rates and because the economy learns more how to live with inflation, as expressed by moving to higher degrees of indexation. Indeed we find that the property of a random walk is not rejected in the quarterly series of inflation in 1973:1-84:4, and that the drift is increasing over time, which implies that inflation was accelerating quickly towards the hyperinflationary environment.

By contrast, the hypothesis of a unit root in the inflation series is rejected for the the Bretton-Woods period and the post stabilization period (after 1985). As noted, this is consistent with the policy game model when the policymaker is concerned not only with the *change* in inflation, but also with its *level*. From a statistical point of view, the latter periods imply a stationary (detrended) inflation series. We find that this has really been the case in the low inflation periods.

In the post-stabilization era we focus on the inflation target regime that was implemented since 1992. In these years the trend of inflation is significantly negative, reflecting the ongoing disinflation process. In this regime there is a loss associated with the deviation of inflation from target, which ensures that the economy reaches price stability down the road.

We find that other nominal variables are governed by the path of inflation throughout the entire period, and the deviations of these variables from inflation tend to be random.

The paper is organized as follows, in section II a simple model is presented, including some intuition regarding the conceptual framework, and the statistical characterization. The empirical analysis appears in section III followed by a discussion of the economic implications. In the end of the paper we present evidence to the fact that other nominal variables behave like the inflation.

II. The Model

a. Simple economic intuition

To bring out the basic economic intuition of the model, before specifying the statistical assumptions, we consider a simple policy game model, as follows. Suppose first that, in the inflationary period, the policy maker is concerned only with maintaining the existing inflation plateau, that is, his aversion to inflation takes the form of an aversion to a *change* in the inflation plateau, *without regard to its level*. In addition, the discretionary policymaker may benefit from a temporary surprise-inflation for reasons stated earlier. Thus, the policymaker has to weigh the loss associated with the change in inflation as against the benefits of surprise inflation.

The loss function of the policymaker can then be expressed as[3]:

(1) $$L = \frac{\beta_0}{2}(\pi_t - \pi_{t-1})^2 - \beta_1(\pi_t - \pi_t^e)$$

where π_t denotes inflation in period *t*, the superscript *e* denotes expectations formed a period earlier and the coefficients β_i are positive constants. The optimal inflation in discretion (which equals the inflation expectations in equilibrium) is given by

(2) $$\pi_t^{opt} = \pi_{t-1} + \frac{\beta_1}{\beta_0} = \pi_t^e$$

which means that inflation will be increasing at a constant rate (the optimal change in inflation is $\Delta\pi_t = \pi_t - \pi_{t-1} = \frac{\beta_1}{\beta_0}$). The fact that inflation goes to infinity with t, certainly

[3] The linear form of $(\pi_t - \pi_t^e)$ is chosen for convenience. The quadratic version would be $(\pi_t - \pi_t^e\text{-A})^2$, where A>0, representing the intention of the policymaker to raise output above the competitive level. Imposing $(\pi_t - \pi_t^e)$=0, we still obtain that the optimal $\Delta\pi_t$ is a positive constant.

precludes long term optimization, but it can be used as an approximation for the inflationary period with all its political and economic uncertainties. Whether the policymakers behaved in a way consistent with the model is a matter for statistical testing (see the quotation from Friedman at the head of the paper).

Statistically, the inflation series will exhibit a random walk (or unit root), since the policymaker is concerned only with the *change* in inflation. Thus although the policymaker is concerned with preserving the inflation plateau he will not succeed in attaining this objective in a policy game equilibrium if he has an incentive to employ surprise-inflation tactics. We consider this as a case of *lack of a nominal anchor*.

It is important to stress that the policymaker in this model is not *planning* a rising trend in inflation, it is rather an outcome of a Nash equilibrium, just as the positive *level* of inflation is a non-intentional outcome of the equilibrium in the Barro-Gordon model under discretion.

For this case we may substitute (2) in (1) to obtain

(1a) $$L(\Delta\pi_t) = \frac{\beta_0}{2}(\Delta\pi_t)^2 - \beta_1\Delta\pi_t + \frac{\beta_1^2}{\beta_0}$$

which is an alternative way of showing that the endogenous variable in this problem is actually $\Delta\pi_t \equiv \pi_t - \pi_{t-1}$ and not π_t. Since the $\Delta\pi_t$'s for different t are independent, it follows from (1a) that the solution (2) is valid regardless the (finite) number of periods considered simultaneously in the loss function[4].

Suppose, alternatively, that the policymaker is concerned in addition with *the absolute level of inflation* (say in the framework of an inflation target regime), then we add

[4] In that respect, it is immediate that the solution (2) for t that satisfies $\frac{\partial L}{\partial \pi_t} = 0,$ guarantees $\frac{\partial L}{\partial \pi_{t-1}} = 0$ as well.

to the loss function in (1) a term[5] $\frac{\beta_2}{2}\pi_t^{\ 2}$, $\beta_2>0$, which yields an optimal inflation of the form[6]

(3) $$\pi_t^{opt} = \frac{\beta_1}{\beta_0+\beta_2} + \frac{\beta_0}{\beta_0+\beta_2}\pi_{t-1} \equiv a + c\pi_{t-1}$$

which implies that π_t converges to $\beta_1/\beta_2 = a/(1-c)$ and does not have a unit root[7], since $c<1$. If, in addition, the policymaker abandons the option of using surprise inflation tactics ($\beta_1=a=0$) then inflation converges to zero. The implied unit-root test equation will be

(3a) $$\pi_t - \pi_{t-1} = s\,\pi_{t-1} + a = -\frac{\beta_2}{\beta_0+\beta_2}\pi_{t-1} + a.$$

with H_0: $s\equiv-(1-c)=0$. If H_0 is not rejected then the data do not contradict the hypothesis of a random walk. The alternative hypothesis is $\beta_2>0$. We may thus consider the policymaker's concern with the absolute level of inflation as representing the existence of a nominal anchor which stabilizes the nominal system at some specified inflation rate.

In the inflationary period which is characterized by the *absence* of a nominal anchor ($\beta_2=0$), the existence of a positive constant drift ($a>0$) will generate a linear trend in the inflation series. However, under these circumstances it is even possible that the parameter *a* (in (3a)) should actually *increase* over time as a result of the reasons stated earlier. For example, an improvement in the indexation arrangements will reduce β_0 and cause an additional rise in inflation over time. It can be assumed that this change in indexation arrangements will be perceived as an externality. In this case we consider *a* as a function of time, $a(t)$, with $a'>0$. As a special case we have a drift with a positive *linear* trend $a(t) = \gamma_0 + \gamma t$, ($\gamma>0$). Combined with $\beta_2=0$ this case implies a quadratic trend in the π_t series (since the

[5] This formulation assumes implicitly an inflation target of zero.
[6] In this case the one period constraint in the loss function is binding.
[7] If we have and explicit inflation target π^T, we penalize deviations from it by inserting in the loss function a term $\beta_2(\pi_t-\pi^T)^2/2$, which is the formulation we use below.

time-slope in (3a) is increasing over time). By contrast, in the presence of a nominal anchor, the inflation series is stationary (there is no unit root).

b. The statistical framework

The statistical framework below of the economic model described earlier, is based on the phenomenon of the inflation steps that characterized the Israeli inflationary process. However, the basic economic considerations of a random walk in the inflation series in the inflationary era and the trend-stationarity in the low inflation periods, is independent of the inflation steps. Nevertheless any economic model of the Israeli inflation process has to be consistent with these steps. In order to adapt the model to the inflation steps we make two (realistic) assumptions: a lag in the formation of inflation expectations by the public (as in wage contracts), and an asymmetric policy reaction to shocks- reflecting the assumption that the policymaker reacts only to crises.

b.1 Absence of a nominal anchor (the inflationary period)

In the statistical framework, we assume that there is a random element in the economy that with probability 1-p induces the monetary policymaker to resort to surprise-inflation (say in order to deal with balance of payments crises[8]), and with probability p not to use surprises. We assume that policymakers observe the realization of period *t* (current) shock prior to making their policy decisions (the inflation rate), while the public determines its period *t* inflation expectations before the realization of current shock. We express this feature by considering β_1 in (1) as a binomial random variable getting the value 0 with probability *p* and the value b > 0 with probability (1-*p*). This means that in ordinary times the policymaker has no incentive (unlike the Barro-Gordon models) to use surprise inflation tactics. In our model, the policymaker uses surprise-inflation tactics only in the case of a negative shock. As noted, the above formulation involves an assumption of *asymmetric* treatment of shocks.

[8] Liviatan and Piterman (1986) show that the main jumps in inflation occurred in conjunction with balance of payments crises.

Specifically, the policymaker reacts only to *adverse* shocks, ignoring favorable ones[9]. This assumption seems reasonable in an inflationary environment. Since we deal with a random walk in the inflationary era, every shock has a permanent effect on the path of inflation. Thus a balance of payments crisis generates a shock that raises the level of inflation permanently[10]. It follows that the rising inflation trend is generated by a series of shocks to which the policymaker reacts in the above manner (this is contrary to the explanation offered by the long term models of the Sargent-Wallace variety, mentioned below).

Given this loss function, the policymaker chooses the optimal path of inflation (that minimizes (1)) satisfying[11]

(4) $$\pi_t - \pi_{t-1} = \begin{cases} 0 & \text{with probabilty} \quad p \\ \dfrac{b}{\beta_0} & \text{with probabilty} \ 1-p \end{cases}$$

In this regime the inflation expectations are derived from (4) as follows

(5) $$\pi_t^e = (1-p)\frac{b}{\beta_0} + \pi_{t-1}$$

It follows from (5) that $\pi^e_t > \pi_{t-1}$ for all t; however, current π_t will equal π_{t-1} (i.e. be below average) with probability p, and will equal $\pi_{t-1}+ b/\beta_0$ (i.e. be above average) with probability 1-p. On average π_{t-1} equals π^e_{t-1}, which implies, by (5), that there is an upward trend with a slope of $(1-p)\dfrac{b}{\beta_0}$ in the inflation process of this regime (we assume that the path of π^e_t represents the inflation trend). These specifications also define a random variable ε_t by

(6) $$\varepsilon_t \equiv (\pi_t - \pi_{t-1}) - E_t(\pi_t - \pi_{t-1}) = \pi_t - E_t\pi_t, \quad \text{with } E_t(\varepsilon_t) = 0.$$

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[9] The asymmetry can be made more explicit by an alternative model. Suppose that the loss function is $2L=(\pi_t-\pi_{t-1})^2+\lambda(y_t-ky^*)^2$, with $k\geq 1$,where y* is full employment output, and $y_t=y^*+\alpha(\pi_t+u_t-\pi_t^e)$. Assume that the shock u_t is observed currently only by the policymaker, but not by the public when forming its expectations. Assume first that k=1. Assume further that $\lambda>0$ for $y<y^*$ and $\lambda=0$ for $y\geq y^*$, reflecting reaction only to negative shocks. Suppose that the probability of $y<y^*$ is (1-p) and of the alternative is p. Then $E_t\Delta \pi_t^{opt}= -(1-p)\lambda \alpha E_t(y_t-y^*)$, where the expectations E_t on the RHS are taken over $y_t<y^*$. The term $-\lambda \alpha E_t(y_t-y^*)>0$ is the counterpart of our b/β_0. If k>1and $\lambda>0$ for all y, we obtain $E_t\Delta \pi_t^{opt}=\lambda\alpha(k-1)y^*>0$. This shows that asymmetric reaction to shocks has similar effect on $E_t\Delta \pi_t^{opt}$ as k>1.

[10] See Liviatan and Piterman (1986) for a detailed description of this process.

[11] In this case it is still true that the one-period optimization leads to an optimal solution for any horizon.

Substituting from (5) into (6) and solving for π_t yield (by iterating the solution)

$$(7) \qquad \pi_t = [(1-p)\frac{b}{\beta_0}]t + \pi_0 + \sum_{j=0}^{t-1} \varepsilon_{t-j}$$

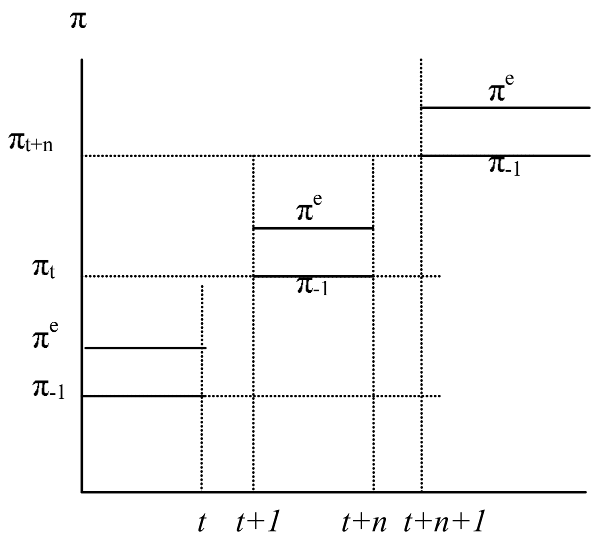
If $(1-p)\frac{b}{\beta_0} \equiv C$ is a positive constant, it means that the inflation series has a positive linear trend. But if, say, the drift is increasing linearly over time, say $(1-p)\frac{b}{\beta_0}$ $=\alpha_1 + \alpha_2 t,$ then we have in (7) a quadratic trend.

The characteristic of the inflation process can be displayed in Figure 1. Inflation in each period can either jump up to a new level or stay at the previous level. Suppose (as is displayed in Figure 1) that there are nominal shocks at *t* and *t*+*n* that drive the policymaker to create inflationary surprises. Accordingly, up to period *t-1* there lasts an *inflationary step* at a constant inflation level of π_{t-1}, while the inflation expectations are above it by the constant C (consistently with (5)). At time *t*, the shock is realized but is still unknown to the public, such that there is a jump in the inflation to the level π_t. However, the inflation expectations remain on their previous level and only in the next period, *t+1*, they will climb to the (higher) level (see Figure 1). That is, the inflation process has (in addition to the deterministic trend associated with the drift) a *stochastic* trend characterized by a non-stationary process.

Note that under our assumption regarding the stochastic process that governs the nominal shocks, the shock is distributed according to the exponential distribution with probability 1-*p*. Therefore, the implied average inflation step is of length of $\frac{1}{1-p}$. If, for example, the probability of the shock to occur is 10% in each period, then the average length of the inflation step will be 10 periods (quarters).[12]

[12] This may explain the shortening of the length of the inflation steps across time in the inflation process in Israel, since as inflation increases later in the inflationary period, the probability of the policymaker to resort to inflation surprise tactics increases as well. By the same token, in the inflation target regime, when 1-*p* tends zero, the length of the inflation step tends (on average) to infinity (i.e. the process of inflation steps is terminated).

Figure 1:
Stochastic Trend and Inflation Steps



So far we have assumed that the only source of random variation is associated with the variation in β_1. We shall refer to this as the "simple case". More generally, there may be other sources of random variation in inflation, such as imperfect control of the money supply or unforeseen external shocks by the policymaker. So we have to interpret ε_t more broadly.

We can now summarize the characteristics of the inflation process of this regime by the following general (nested) stochastic inflation process, which admits a linearly increasing drift:

(8) $$\pi_t = \pi_0 + \alpha_1 t + \alpha_2 t^2 + \sum_{s=1}^{t} \varepsilon_s$$

or

(9) $$\pi_t = \alpha_0 + \alpha_1 t + \alpha_2 t^2 + \theta_t, \qquad \text{with} \quad \theta_t = \theta_{t-1} + \varepsilon_t, \quad \alpha_0 = \pi_0.$$

where α_0, α_1 and α_2 are scalars. The series $\{\varepsilon_t\}$ is a white noise process with zero mean and fixed variance σ^2_e. Note that the stochastic process $\{\theta_t\}$ is non-stationary and has a unit root[13]. Finally, for the unit root test we have from (8) the following test equation (abstracting from lagged values of the dependent variable, π_t - π_{t-1}, to which we refer later)

[13] It reflects the stochastic trend we alluded to before and that past realizations of ε have permanent effects on inflation.

(10) $\pi_t - \pi_{t-1} = \rho\pi_{t-1} + \alpha_1 - \alpha_2 + 2\alpha_2 t + \varepsilon_t$,

with $H_0 : \rho=0$. If α_2 is positive, indicating a quadratic trend, then the sign of the constant term ($\alpha_1 - \alpha_2$) is ambiguous. The variance of the inflation in (8) is not constant over time and obeys the following expression

(11) $\text{var}(\pi_t) = t\sigma^2_e$

b.2 With a nominal anchor

Consider next a policy regime where the policymaker objective loss function (1) with the addition of a third term which reflects aversion to deviation from an informal (as in the case of Bretton-Woods) or formal inflation target (π^T). Let the loss function be

(12) $$L = \frac{\beta_0}{2}(\pi_t - \pi_{t-1})^2 - \beta_1(\pi_t - \pi_t^e) + \frac{\beta_2}{2}(\pi_t - \pi_t^T)^2$$

where, as before, β_1 is a binomial random variable getting the value 0 with probability p and the value b > 0 with probability (1-p).

Given this loss function, the policymaker chooses the optimal path of inflation (that minimizes (12)), which yields

(13)

$$\pi_t - \pi_{t-1} = \begin{cases} -\frac{\beta_2}{\beta_0+\beta_2}(\pi_{t-1} - \pi^T_{t-1}) + \frac{\beta_2}{\beta_0+\beta_2}(\pi^T_t - \pi^T_{t-1}) & \text{with probabilty } p \\ \frac{b}{\beta_0+\beta_2} - \frac{\beta_2}{\beta_0+\beta_2}(\pi_{t-1} - \pi^T_{t-1}) + \frac{\beta_2}{\beta_0+\beta_2}(\pi^T_t - \pi^T_{t-1}) & \text{with probabilty } 1-p \end{cases}$$

In this regime the public's inflation expectations are derived from (13) as follows

(14) $$\pi_t^e = (1-p)\frac{b}{\beta_0+\beta_2} + \frac{\beta_2}{\beta_0+\beta_2}\pi^T_t + \frac{\beta_0}{\beta_0+\beta_2}\pi_{t-1}$$

Given that $\pi_t - \pi_t^e = \varepsilon_t$ is white noise, we have

(15) $$\Delta\pi_t = (1-p)\frac{b}{\beta_0+\beta_2} + \frac{\beta_2}{\beta_0+\beta_2}\pi^T_t - \frac{\beta_2}{\beta_0+\beta_2}\pi_{t-1} + \varepsilon_t$$

If the inflation target regime is strictly implemented we may assume that the policymaker will abandon the use of surprise inflation tactics[14]. If this is internalized by the public, we may set p=1 in (15). In a disinflation process the inflation targets follow some predetermined declining time path. It is then immediate from the second term in (15) that the inflation process possesses a deterministic trend. Note also that $0 < \frac{\beta_2}{\beta_0 + \beta_2} < 1$, and therefore the inflation process is stationary (there is no unit root)[15].

Still assuming p=1, we have that π_{t-1} equals on average π^e_{t-1}, and it follows from (14) that inflation expectations converge to the inflation target over time. Since in the disinflation period the target is decreasing, the regime ensures price stability down the road. If, as an approximation, we let the inflation target follow a declining linear trend, we can rewrite (15) as follows

(16) $$\Delta\pi_t = k_0 - k_1 t - k_2\pi_{t-1} + \varepsilon_t, \qquad 0<k_2<1,\ k_i>0$$

which does not have a unit root and follows a negative deterministic trend. In the Bretton-Woods regime we may set k_1=0, so that $k_0=\frac{\beta_2}{\beta_0 + \beta_2}\pi^T_t$, where the informal inflation target was historically small but positive.

As a preliminary comment to the statistical computations we note that the feature of a rising or decreasing inflation series is a separate issue from the existence of a unit root. The former can be influenced by the existence of an exogenous trend while the latter refers to the *detrended* series. In addition, there are theories of a rising inflation trend which are completely independent of the stochastic structure of the model, [such as the long term

[14] This may seem as an extreme assumption since it postulates that the policymaker does not react by inflationary measures even to observed balance of payment crises, but it is not unrealistic in a disinflation environment.

[15] Why wouldn't the inflation expectations under these circumstances be equal to the inflation target? Indeed, this would be the case if β_0 = 0. But since the reduction in the inflation to the target (within the disinflation policy) involves a cost (see objective function), the reduction is implemented gradually. As long as the slope of the inflation target is constant over time - actual inflation will exceed the target.

models of Sargent-Wallace (1981), Liviatan (1984) and Drazen-Helpman (1990)], while in our model the statistical structure plays a dominant role. Since these models do not refer at all to the inflation steps, they are less relevant to the explanation of the Israeli inflation process.

III. Empirical Analysis

The fifty-year period under our investigation 1955-2005 can be divided into four sub-periods as far as the monetary policy is concerned: (1) The Bretton-Woods period 1955-69 with low budget deficits and a non-inflationary world, (2) the period from 1970.1 to 1985.2 (prior to the 1985 stabilization plan) in which the government budget deficit was high and the monetary policymakers were subservient to budgetary needs (fiscal dominance), (3) the period from 1985.3 to 1991.4 (after the 1985 stabilization but prior to the institution of the inflation target) in which the fiscal authority ran a balanced budget and the monetary policymakers were dominant but did not announce yet inflation targets. This period should be conceived as a period of transition from a financially unstable environment to a stationary one in which the policymakers gain credibility; (4) the period from 1992.1 to 2005.1 in which the monetary authority had the dominance and the announced monetary policy included inflation targets[16]. In most of this period the monetary policy included a disinflation policy as was reflected by a downward trend of the pre-announced inflation targets.

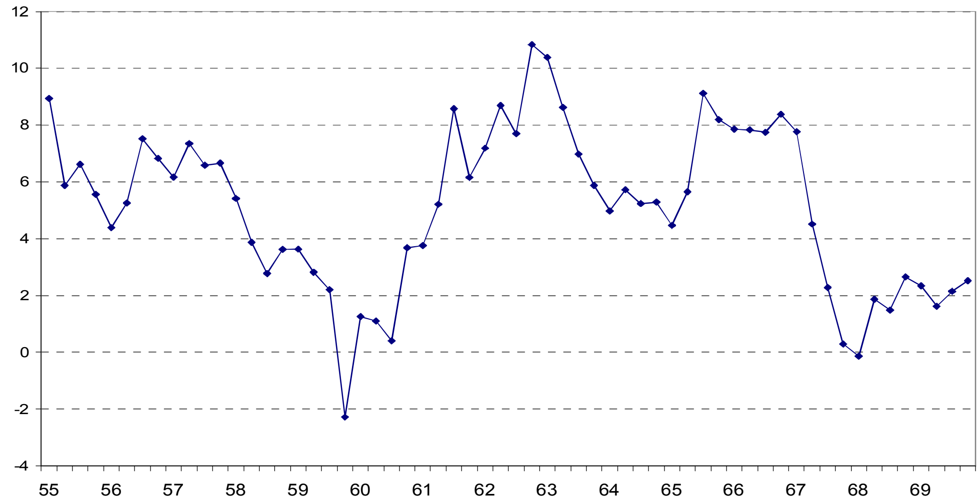
a. Data

To focus on the properties of the inflation processes, we use quarterly data on the Israeli inflation in the period 1955.1 up to 2005.1. Inflation is measured by the quarterly change in the CPI. In figure 2 we depict the (last 4 quarters)) inflation process of the first period, 1955-69. It can be seen that inflation is low and non-increasing in this period, with a tendency to

[16] The effective inflation targets were instituted only in 1994. However, the exchange rate policy that was implemented in 1992 required the announcement of the official inflation targets (See Djivre & Tziddon (1994) for an elaboration on the inflation targets in the 90's in Israel).

decline in the last years[17].

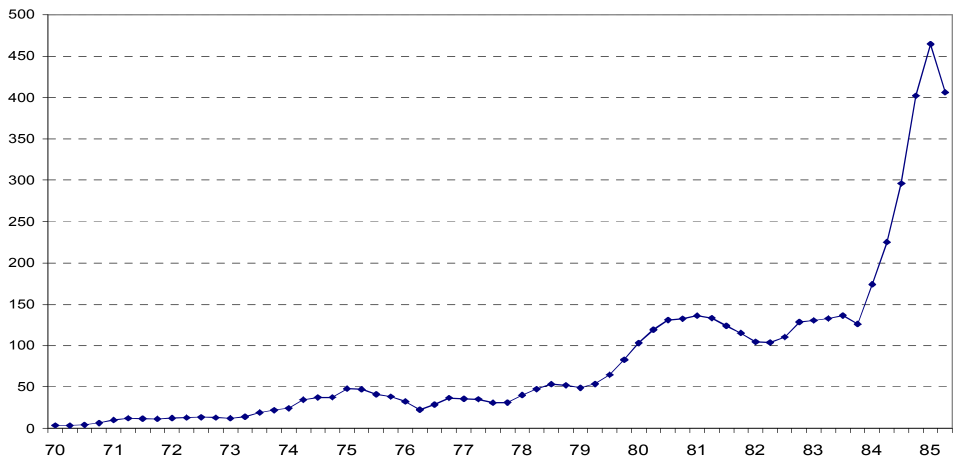
Figure 2:
The Israeli Inflation: 1955 – 69
(last 4 quarters)



Source: The CBS: *Statistical Abstract of Israel* 2005.

Then there was the inflationary period 1970.1–1985.2 (Figure 3). There are three features that seem to come up from this figure: i) The upwards trend of the inflation; ii) the inflation process may be characterized by a step function (1973-79, 1979-83, 1983-1985); iii) the relative steep trend of 1983-85.

Figure 3:
The Israeli Inflation: 1970 – 85
(last 4 quarters)

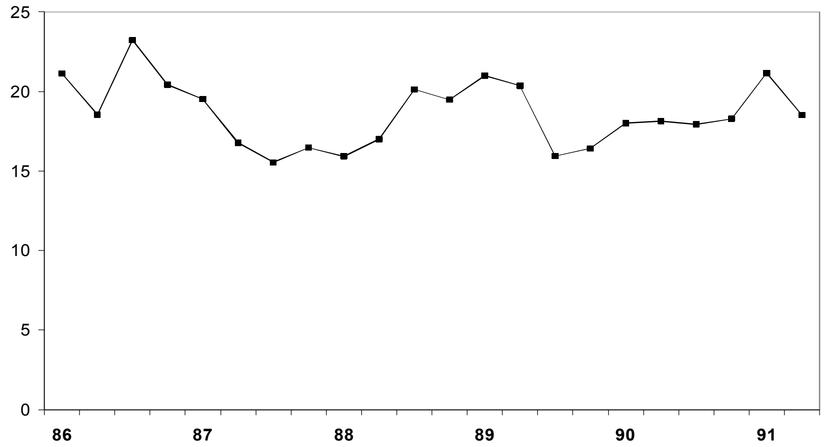


Source: The CBS: *Statistical Abstract of Israel* 2005.

[17] Barkai (2004) reports that the annual average change in the effective exchange rate on imports was 7.6% in 1955-59, 3.2% in 1960-66, and 5.5% in 1967-69. Annual (CPI) inflation in the same periods was 4.6%, 6.5% and 2.1%. The average four-quarter inflation in 1955-69 was 5.4%.

The inflation series for the second period 1985.3–91.4 is described in the next figure. For coherency we omitted the first four observations (1985.3-86.2) from the data displayed in figure 4 since we did not want to include any of the high inflationary quarters. It is apparent from the data that the inflation fluctuates around the level of 18-20 percent.

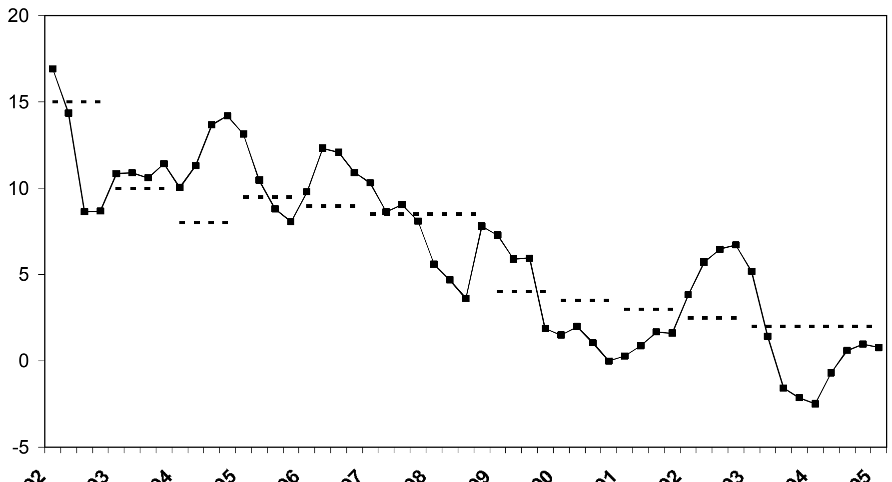
Figure 4:
The Israeli Inflation: 1986.3 – 1991.4
(last 4 quarters)



Source: The CBS: *Statistical Abstract of Israel* 2005.

Finally, we display both the inflation series and the inflation targets for the remaining period: 1992.1 – 2005.1 in figure 5 below. The deterministic trend of the inflation targets can be clearly identified from the data.

Figure 5:
The Israeli Inflation and inflation targets: 1992.1 – 2005.1
(last 4 quarters)



Source: The CBS: *Statistical Abstract of Israel* 2005.

Consistent with the analysis of section II, we put to test for each of the four periods, the following hypotheses regarding the inflation processes:

b. The Bretton Woods period 1955.1-1969.4

The economic policy in this period was dominated by the Bretton-Woods rules, which created a stable macroeconomic environment. Accordingly in order to match the characteristics of this period inflationary process, we set $p = 1$ in (15) and interpret π^T as an implicit inflation target dictated by the Bretton-Woods system. Our assumption is that in this system the policymaker's concern is with the level of inflation, and hence the unit root hypothesis should be rejected. Indeed the results in Table 1, in which we report the results of the unit root test for 1955:1-1969:4, reject the unit root hypothesis overwhelmingly. The trend is insignificant (as expected), and the constant term reflects the low positive implicit inflation target that was allowed by the Bretton-Woods regime. In this calculation we let the structure of the lags be determined automatically by the *E-views*5 software, which indicates that there were no lags in the inflation process.

Table 1:
The Results of the Unit Root Tests
1955.1 – 1969.4

Dependent variable*:	a. D(π)	
	t-Stat	Prob.
Augmented Dickey-Fuller Test statistic	**-8.58**	**0.00**
1% level	-4.12	
5% level	-3.49	
10% level	-3.17	
Independent variables:		
	Coef.	P-value
π_{t-1}	-1.12	0.00
C	-1.51	0.00
@trend(1973.1)	-0.01	0.62
No. of Observations	60	
Adj R^2	0.55	
D.W.	1.96	

* D is the difference operator.

We conducted a sensitivity analysis by performing the unit root test on various shortened samples in order to examine the robustness of our results with respect to the unit root and the existence of a deterministic trend in the data. The results of the analysis are reported in Table 2. We let *Eviews5* choose the lag-structure of the dependent variable in accordance with SIC criterion. Indeed the results indicate the robustness of these statistically confirmed characteristics. In all the samples, the test could not reject the existence of a unit root.

Table 2:
Sensitivity Analysis of the Existence of Unit Root and Deterministic Trend in the Israeli Inflation Data[a]

Sample: From 55.1 To	ADF *t*-stat.	Critical value 5%	*P*-value	No. of Lags (qtrs)	Constant *t*-value	Trend *t*-value	Adj. R^2	D.W.
1969.4	-8.58	-3.49	0.00	0	3.59	-0.51	0.55	1.96
1969.2	-8.36	-3.49	0.00	0	3.41	-0.29	0.55	1.96
1968.4	-8.20	-3.49	0.00	0	3.32	-0.29	0.54	1.96
1968.2	-8.23	-3.50	0.00	0	3.13	0.18	0.55	1.96
1967.4	-8.05	-3.50	0.00	0	3.01	0.17	0.55	1.95
1967.2	-8.23	-3.50	0.00	0	2.84	0.91	0.57	1.98

a. These Results are outcomes of ADF tests that were performed on Eviews5.

c. The period 1970.1 – 1985.2

This is the inflationary period which is the center of the paper. Here the statistical analysis is complicated because of inflation volatility. Our basic hypothesis for this period is that the detrended inflation follows a random walk. Econometrically there is the question what is the proper way of detrending the inflation series. It turns out that the various alternative ways have different implications on the results. We shall present accordingly two versions of dealing with this problem. The first one is to follow the standard procedure of including a linear trend in the test equation of a unit root in the Eviews5 program (this equation has the *change* in inflation as the dependent variable). This allows for a quadratic trend in the inflation series. The other is to compute first the inflation trend according to the Hodrik-Prescott filter, and then test for a unit root of the deviation of inflation from this trend. This dual treatment is more relevant for the inflationary period, because in the other periods the results are more clear-cut. We begin with the first method.

b.1 Including a linear trend in the test equation

According to this approach we make use of the representation (10), which describes a process that was characterized by a unit root and includes both a deterministic trend and a stochastic trend. For convenience we rewrite (10) as follows:

(17) $$\Delta\pi_t = \rho\pi_{t-1} + \alpha_3 + 2\alpha_2 t + e_t, \quad \text{where } \alpha_3 = \alpha_1 - \alpha_2 \text{ in (10)}$$

with H_0: $\rho=0$[18]. According to the model presented earlier H_0 should not be rejected for this period and α_2 is allowed to be positive .

This equation reflects our hypothesized inflation process for the high inflationary period 1973.1-1985.2. Although we define this period to last up to the date when the government officially adopted the price stabilization pact in 1985.2, Figure 1 indicates that there was a turning point in the inflation process two quarters earlier. It may indicate that the change in the reference of the policy makers to the level of inflation started already at the beginning of 1985. Therefore, in our statistical tests, we use the sample up to the end of 1984. In what follows we report the results of the unit root test we performed using quarterly inflation data of this period.

Accordingly we performed a number of Augmented Dickey-Fuller Unit Root tests. “Augmented” means that we allow the lagged values of $\Delta\pi_t$ to enter the test equation. It turns out that the treatment of lags has important consequences for our tests[19]. Generally, the arbitrary ruling out of lags tends to raise the ADF statistic and is thus biased against the unit root hypothesis. This concerns only the inflationary period, since in the post stabilization the automatic selection by Eviews5 results in a zero number of lags of the dependent variable in the test equation (see tables 4 and 6).

[18] The alternative hypothesis (H_1) is that the inflation series is trend-stationary.

To see the meaning of the inclusion of lagged values of $\Delta\pi_t$ in the test equation, we have to redefine our concept of inflation to

$$\pi_t^R \equiv \pi_t - \delta_1\pi_{t-1} - \delta_2\pi_{t-2}$$

(confining the example to two lags), and consequently our loss function will be

$$L = \frac{\beta_0}{2}(\pi_t^R - \pi_{t-1}^R)^2 - \beta_1(\pi_t^R - \pi_t^{\mathrm{Re}}) + \frac{\beta_2}{2}(\pi_t^R)^2,$$

where π_t^{Re} is the corresponding newly defined inflation expectations. Under the null hypothesis for the inflationary period (no nominal anchor) β_2=0, the optimal change in current inflation is

$$\Delta\pi_t = \frac{\beta_1}{\beta_0} + \rho\pi_{t-1} + \delta_1\Delta\pi_{t-1} + \delta_2\Delta\pi_{t-2}$$

with H_0: $\rho = 0$. Assuming a linear trend in the drift $\frac{\beta_1}{\beta_0} = \gamma_0 + \gamma t$ we obtain the augmented DF test equation.

Note that π_t^R is the part of current inflation that is not "explained" by past inflation, as may be the case under indexation arrangements. Indeed, the number of lags increases after the surge in inflation to the 130% plateau in the early 80's (see table 4). The modified test equation suggests that it is only the inflation in excess of (partial) indexation that constitutes a "surprise". Since this seems reasonable, we are more inclined to accept the ADF test without restraining the number of lags. Our procedure in the main calculations was, accordingly, to allow the number of lags to be determined automatically by the *Eviews*5 software (by utilizing the Schwarz Info Criterion, SIC).

The results of our tests are summarized in the two panels of Table 3. In panel (a) we let the *Eviews*5 program determines the lag structure of the dependent variable that appears in the RHS of the estimated equation. In panel (b) we test the existence of a unit root in the

[19] Another problem which may potentially affect the results is seasonality. We performed an indirect test on this factor and it turned out that it does not affect the unit root test.

series of the quarterly change in the inflation, $\Delta\pi_t$, where we restricted the number of lagged dependent variable to zero.

Table 3:
The Results of the Unit Root Tests
1970.1 – 1984.4

Dependent variable*:	a. D(π)		b. D(D(π))	
	t-Stat	Prob.	t-Stat	Prob.
Augmented Dickey-Fuller Test statistic	**-0.49**	**0.98**	**-12.1**	**0.00**
1% level	-4.12		-3.54	
5% level	-3.49		-2.91	
10% level	-3.17		-2.59	
Independent variables:				
	Coef.	P-value	Coef.	P-value
π_{t-1}	-0.11	0.63		
$D(\pi_{t-1})$	-0.29	0.25	-1.49	0.00
$D(\pi_{t-2})$	0.03	0.91		
$D(\pi_{t-3})$	-0.05	0.84		
$D(\pi_{t-4})$	0.44	0.02		
C	-1.12	0.44	1.54	0.05
@trend(1973.1)	0.12	1.12		
No. of Observations	60		60	
Adj R^2	0.35		0.71	
D.W.	1.92		1.81	

* D is the difference operator.

The results indeed support our conjecture regarding the existence of a unit root in the inflation process in this period (see panels (a)). Also, the results in panel (b) reject the hypothesis of the existence of a unit root in the series of the quarterly *change* in inflation, so that the inflation series is indeed I(1).

We also conducted a sensitivity analysis by performing the unit root test on various shortened samples in order to examine the robustness of our results with respect to the unit root and the existence of a deterministic trend in the data. The results of the analysis are reported in Table 4. We let *Eviews*5 choose the lag-structure of the dependent variable in accordance with SIC criterion. Indeed the results indicate the robustness of these statistically confirmed characteristics. In most of the samples, the test could not reject the existence of a unit root. Also, regarding the hypothesized quadratic deterministic trend, we find in all samples the sign is positive, and that in most samples (in Table 4) the *t*-value indicates the existence of a significant trend. The constant term is insignificant, which according to (10) means that α_2 (which seems to be significantly positive) cancels out the positive α_1.

Table 4:
Sensitivity Analysis of the Existence of Unit Root and Deterministic Trend in the Israeli Inflation Data[a]

Sample: From 70.1 To	ADF *t*-stat.	Critical value 5%	*P*-value	No. of Lags (qtrs)	Constant *t*-value	Trend *t*-value	Adj. R^2	D.W.
1984.4	-0.49	-3.49	0.98	4	-0.77	1.12[c]	0.35	1.92
1984.2	-1.28	-3.49	0.88	1	-0.72	1.84	0.23	1.89
1983.4	-4.71	-3.49	0.00[b]	0	-0.86	4.69	0.28	1.67
1983.2	-2.79	-3.50	0.21	4	-0.23	2.73	0.51	1.81
1982.4	-2.72	-3.50	0.23	4	-0.21	2.66	0.50	1.80
1982.2	-2.70	-3.50	0.24	4	-0.18	2.58	0.53	1.79
1981.4	-3.29	-3.51	0.08	8	-0.23	2.99	0.64	2.08
1981.2	-2.80	-3.51	0.20	8	-0.21	2.83	0.61	2.09
1980.4	-0.22	-3.52	0.99	5	-0.14	0.94[c]	0.52	1.83
1980.2	-0.24	-3.52	0.99	5	-0.13	0.84[c]	0.47	1.84

a. These Results are outcomes of ADF tests that were performed on Eviews5. b. The unit root hypothesis is rejected at 5 percents. c. The trend hypothesis is rejected at 5 percents.

b.2 Detrending by the HP filter

According to this approach we compute first the inflation trend for the 1973-1984 period by the HP filter. It can be seen in figure 6 and 7 that not only there is a positive trend in the inflation process, but also that the slope of the trend itself is increasing. Moreover, figure 6 indicates that the increase in the slope accelerates in the eighties, which implies that the inflation trend accelerates even more than the quadratic function.

Figure 6:
The quarterly inflation and the Hodrick-Prescott Filter (lambda=1600) in Israel, 1970.1 1985.2

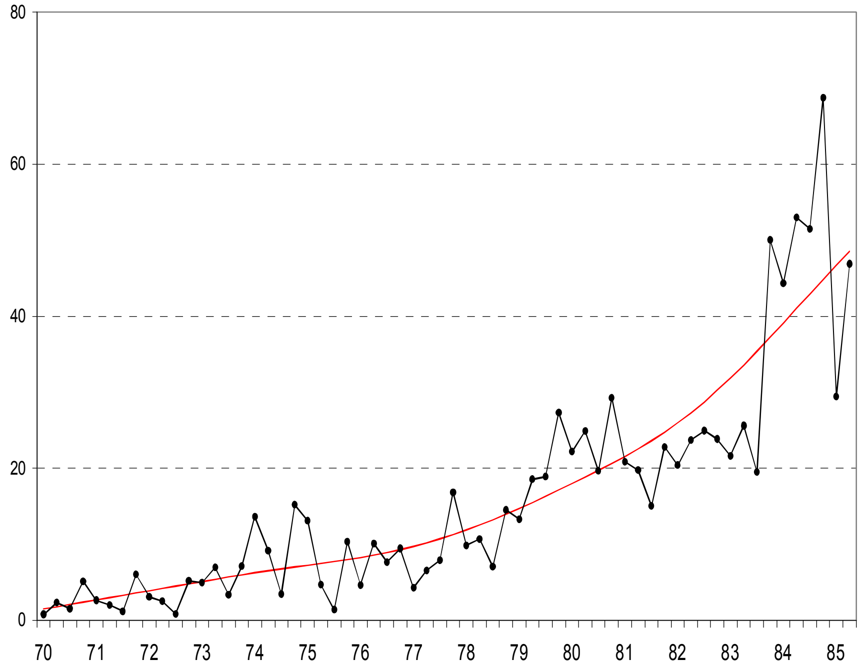


Figure 7:
The first difference of the Hodrick-Prescott filter of quarterly inflation and its Trend in Israel, 1970.1- 1985.2

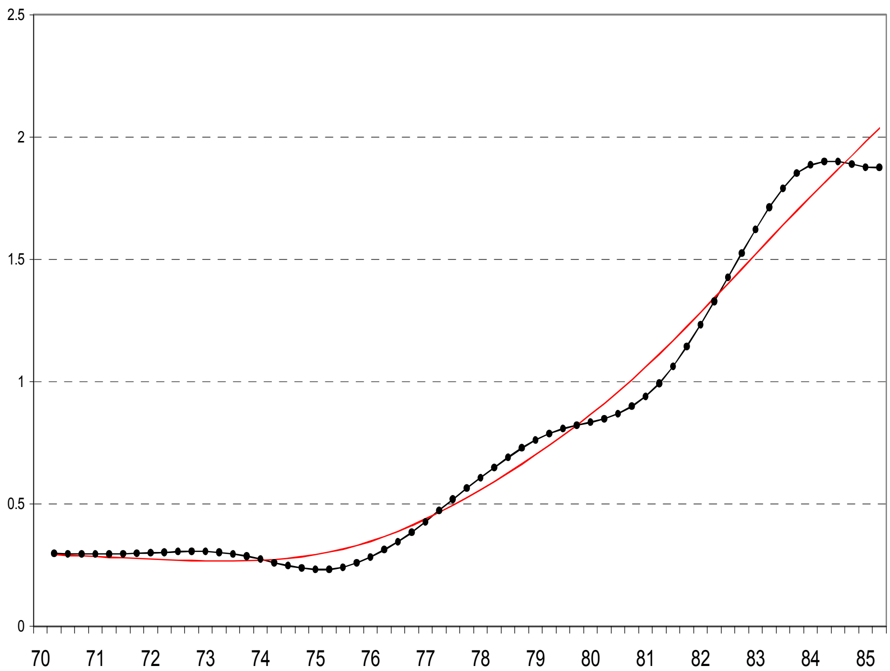


Table 3a shows the results of the unit root test for the residuals of quarterly inflation from the HP trend, on the pattern of table 3 above. The results suggest that the detrended series possess a unit root, as is hypothesized by the model.

Table 3a:
The Results of the Unit Root Tests
1973.1 – 1984.4

Dependent variable*:	a. D(π-HPπ)		b. D(D(π-HPπ))	
	t-Stat	Prob.	t-Stat	Prob.
Augmented Dickey-Fuller Test statistic	**-2.86**	**0.18**	**-11.4**	**0.00**
1% level	-4.13		-3.57	
5% level	-3.49		-2.92	
10% level	-3.18		-2.60	
Independent variables:				
	Coef.	P-value	Coef.	P-value
π_{t-1}-HPπ_{t-1}	-0.71	0.01		
D(π_{t-1}-HPπ_{t-1})	0.15	0.53	-1.52	0.00
D(π_{t-2}-HPπ_{t-2})	0.35	0.11		
D(π_{t-3}-HPπ_{t-3})	0.22	0.27		
D(π_{t-4}-HPπ_{t-4})	0.59	0.00		
C	-0.33	0.83	0.62	0.49
@ trend(1973.1)	0.01	0.78		
No. of Observations	55		48	
Adj R^2	0.44		0.73	
D.W.	1.99		1.89	

* D is the difference operator.

We also carried out a sensitivity analysis on the pattern of Table 4. The results are presented in Table 4a and indicate the existence of the unit root in the detrended data throughout the period. The lag pattern in 1984:2 and 1983:4 seem to be outliers to the general pattern of four lags. Since the lag is not supposed to change in the short run (reflecting indexation arrangements), we experimented with forcing a four quarter lag in the above two periods. This yielded ADF values for these two periods that are in line with the others, suggesting the existence of unit roots.

Table 4a:
Sensitivity Analysis of the Existence of Unit Root
in the Detrended Israeli Inflation Data[a]

Sample: From 70.1 To	ADF t-stat.	Critical value 5%	P-value	No. of Lags (qtrs)	Constant t-value	Trend t-value	Adj. R^2	D.W.
1984.4	-2.86	-3.49	0.18	4	-0.21	0.29	0.44	1.99
1984.2	-5.18	-3.49	0.00[b]	0	0.19	-0.29	0.31	2.03
1983.4	-5.54	-3.49	0.00[b]	0	0.70	-1.11	0.35	1.83
1983.2	-2.32	-3.50	0.42	4	0.70	-1.10	0.48	1.82
1982.4	-2.67	-3.51	0.25	4	0.42	-0.63	0.50	1.80
1982.2	-3.19	-3.51	0.10	4	0.00	0.04	0.56	1.89
1981.4	-3.45	-3.53	0.06	8	0.20	-0.14	0.64	2.13

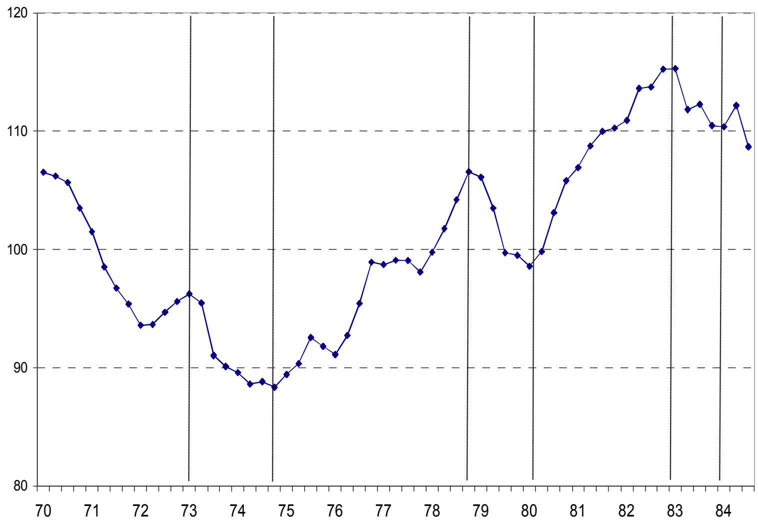
a. These Results are outcomes of ADF tests that were performed on Eviews5. b. The unit root hypothesis is rejected at 5 percents.

The phenomenon of the inflation steps.

The model described in Figure 1 suggests that the real wage should rise in the course of the inflation step and should decrease sharply the crisis. This is because the nominal wage rate is supposed to reflect inflation expectations which take into account the probability of the shock. There were three major crises in the inflationary period: In the end of 1973 (the first oil shock and the Yom-Kippur war), in 1979 (the second oil shock) and in the end of 1983 (the collapse of the stock market in the wake of the failure of Aridor's Tablita policy). The associated inflation steps as reported in Liviatan and Melnick (1998) were in 1973:4-1979:1 and in 1979:2-1983:3.

An examination of the unit labor cost (ULC) curve (a moving average of four quarters), shows that these crises were in fact associated with a decline in ULC and a subsequent rise along the inflation step (Figure 8). This confirms also the presumption that the crises came as a surprise to the labor market.

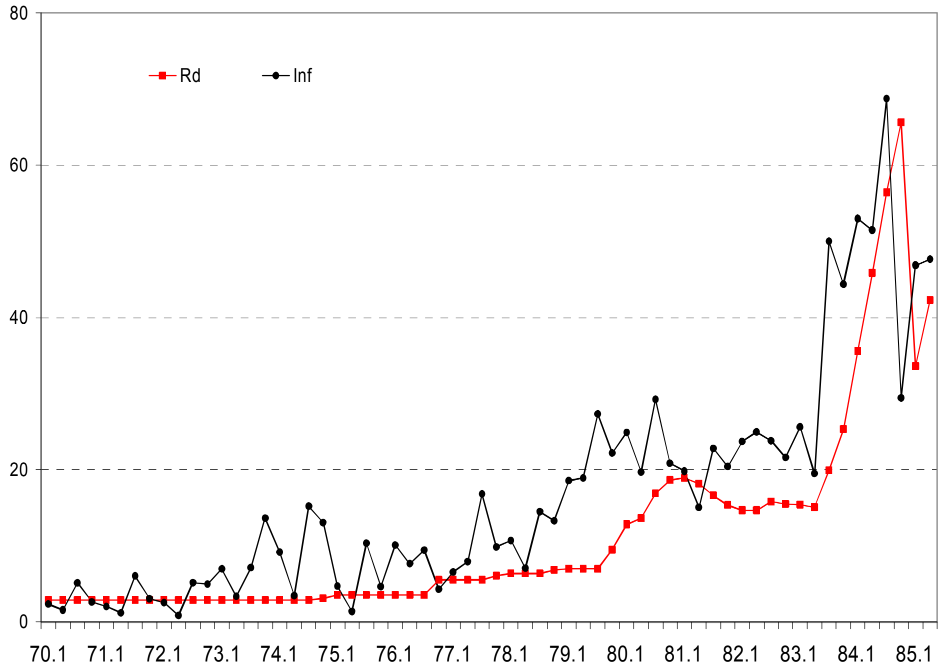
Figure 8:
Real Unit Labor Cost in Israel, 1970 – 1985
(moving average last 4 quarters)



Source: The Statistical Abstract of Israel, 2005, *Central Bureau of Statistics* and The Bank of Israel.

Another indicator of the surprise element is the real interest rate on short term deposits. That is, the comparison of the evolution of the nominal interest rate to that of the inflation. The problem with this indicator is that it is adjustable in a short time and that is much influenced by the developments in the capital markets. Nevertheless, in Figure 8, quarterly data of nominal interest rate on commercial banks' deposits are depicted along with one-quarter-ahead inflation rates. It does indicate a collapse of the real (ex-post) interest rate on deposits in 1973, 1979 and in 1983 crises. It can be seen that only as late as in the first quarter of 1975 the interest rate was first adjusted upward in response to 1973 inflation shock. Similar lagged adjustments of the nominal interest can be seen to characterize both 1979 and 1983 nominal shocks.

Figure 9:
Commercial banks' nominal interest rate on deposits and inflation*
in Israel, 1970.1 – 1985.2
(Quarterly data)



* The data on inflation are of one quarter ahead.
Source: Bank of Israel.

c. The period 1985.3 – 1991.4

This relatively short post-stabilization period is characterized by a stable rate of inflation of around 18-20 percent annually. In this period, that followed the 1985 stabilization, we can assume a concern about the level of inflation ($\beta_2 > 0$ in 10). The fact that inflation did not drop to international levels can be attributed to the possibility that the probability p (in (4)) has not risen to unity (the full credibility that the central bank has abandoned the use of surprise inflation), or to the belief that the policymakers were content with achieving a moderate inflation (i.e. they adopted an implicit inflation target of about 20 percent). Accordingly we postulate the following stationary inflation process, which is a slight modification of (15):

$$(18) \qquad \Delta\pi_t = [(1-p)\frac{b}{\beta_2} + \pi_t^T]\theta - \theta\pi_{t-1} + \varepsilon_t, \quad \text{where } \theta = \frac{\beta_2}{\beta_2 + \beta_0}$$

where π_t^T should be interpreted as the implicit target.

For this model it is reasonable to regard π_t^T as a constant π^T, and hence inflation converges to $[(1-p)\frac{b}{\beta_2}+\pi^T]$. Denote this value by Q and let $\phi_t = \pi_t - Q$. Then the model[20] can be described as

(18a) $\phi_t = \lambda\phi_{t-1} + \varepsilon_t$ where $-1 < \lambda < 1$, or

(19) $\phi_t = \sum_{i=0}^{t} \lambda^i \varepsilon_{t-i} + \lambda^t \phi_0$

and

(20) $\pi_t - \pi_{t-1} = -(1-\lambda)(\pi_{t-1} - Q) + \varepsilon_t \equiv \theta Q - \theta\pi_{t-1} + \varepsilon_t$

In terms of (18), $\lambda = \frac{\beta_0}{\beta_0 + \beta_2}$ and therefore $0 < \lambda < 1$. However, in the foregoing formulation the model converges under more general conditions, allowing $-1<\lambda<1$, i.e. the range $-1<\lambda<0$ is also acceptable.

From (19) we can see that the effect of past nominal shocks on current and future inflation is dissipating and inflation converges to Q. This value reflects the reduction of the fiscal deficit and the concern with the level of inflation, but does not succeed in delivering price stability because of the absence of an explicit and declining target. The hypothesized inflation process should have a stationary representation (no unit root) without a deterministic trend. This is essentially the result of the stabilization plan that was undertaken in 1985.

Our main concern with this period is whether there was a growing concern with the absolute level of inflation, so that the inflation series is convergent. We accordingly performed the Augmented Dickey-Fuller Unit Root tests for this period under the hypothesis (20) regarding the inflation processes. The results are summarized in Table 3.

[20] Note that this model with p=1 fits the Bretton-Woods model described earlier.

In panel (a) the test is restricted to include one lagged dependent variable, while in panel (b) of Table 3, we report the results of the unit root test for the inflation process in which the RHS explanatory variables include no lagged dependent variables, as is structured by the *E-views*5 in accordance with the SIC criterion. The results in both panels clearly indicate the rejection of a unit root for this period. However, the fact that the absolute value of the coefficient of π_{t-1}, ((1-λ) in (20)), is greater than unity in Table 5 (but less than two), is indeed in contradiction to (18), but in agreement with the more general formulation of (20). The model is still convergent because the absolute value of the estimate of λ is less than unity (Table 5), and therefore does not endanger the stability of the inflation process.

Table 5:
The Results of the Unit Root Tests
1986.1 – 1991.4

Dependent variable*: D(π)	a.		b.	
	t-Stat	Prob.	t-Stat	Prob.
Augmented Dickey-Fuller Test statistic	**-7.86**	**0.00**	**-8.35**	**0.00**
1% level	-4.39		-4.39	
5% level	-3.61		-3.61	
10% level	-3.24		-3.24	
Independent variables:				
	Coef.	P-value	Coef.	P-value
π_{t-1}	-1.38	0.00	-1.42	0.00
D(π_{t-1})	0.04	0.30		
C	6.22	0.00	6.14	0.00
@Trend(1986.3)	-0.01	0.79	0.007	0.88
Adj R^2	0.75		0.75	
D.W.	1.96		1.86	

* D is the difference operator.

Thus it appears that in the first period of the stabilization, the regression coefficient of π_t on π_{t-1} was not only less than one (as in (18)), but was in fact negative (but still less than one in absolute value, as required for stability). This implies that an increase in inflation of 1% was followed in the next quarter by *reduction* in inflation of (say) 0.3%, although an *increase* of less than 1% would had been sufficient (formally) for stability. This feature of disinflation, which presumably reflects a cautious stance in the wake of disinflation is not captured by the model. It stems apparently from the tight monetary policy in response to the historical heritage of high inflation. Although this feature is not captured by the model, it does provide valuable information which is uncovered by comparing the empirical results with the implications of the model.

We also conducted a sensitivity analysis on quarterly data letting the number of lags be determined by SIC. The results presented in Table 6 below confirm our conclusions about the absence of a unit root. They also point to a significant constant drift in this period which may indicate that *p* is still less than one. But it does not endanger the stability of inflation as long as the process is stationary.

Table 6:
Sensitivity Analysis of the Existence of Unit Root and
Deterministic Trend in the Israeli Inflation Data[a], 1986:1-1991:4

Sample: From 86.1 To	ADF t-stat.	Critical value 5%	P-value	No. of Lags (qtrs)	Constant t-value	Trend t-value	Adj. R^2	D.W.
1991.4	-8.35	-3.61	0.00	0	6.19	0.15	0.75	1.86
1990.4	-8.02	-3.66	0.00	0	5.79	-0.15	0.77	1.68
1989.4	-7.06	-3.73	0.00	0	4.88	-0.12	0.77	1.69

a. These Results are outcomes of ADF tests that were performed on Eviews5.

d. The period 1992.1 – 2005.1

The hypothesized inflation process has the representation (8), i.e. the process has a deterministic process (without a stochastic trend), and hence the detrended inflation series should be stationary and should have no unit root. The predetermined trend should be generated from the path of the inflation targets. Again for convenience we rewrite (16) as follows:

(16) $$\Delta\pi_t = k_0 - k_1 t - k_2 \pi_{t-1} + \varepsilon_t, \qquad 0<k_2<1,\ k_i>0$$

As for the previous periods we performed the Augmented Dickey-Fuller unit root tests, given the hypothesis (16) regarding the inflation processes. The results are reported in the 3 panels of Table 7 below. In panel (a) we report the result of the Unit Root test where we allow a deterministic trend in the inflation series and the inclusion of the lagged dependent variables is determined in accordance with the SIC. The results indicate clearly the existence of a stationary inflation process (there is no unit root) with a negative deterministic trend. In panel (b) we restricted the test to include 1 lagged dependent variable in the RHS and the results remain similar to those reported in panel (a). Finally, in panel (c) we explicitly detrended the inflation series by substracting the trend of the inflation target (generated by the HP filter). As in panel (a) we allow *Eviews* to choose the lagged dependent variables in according with the SIC, but we do not allow a deterministic trend in the detrended series. The results indicate the stationarity of the detrended inflation series as well, supporting our hypothesis.

Table 7:
The Results of the Unit Root Test
1992.1 – 2005.1

Dependent variable*:	a. $D(\pi)$		b. $D(\pi)$		c. $D(\pi-\pi^T)$	
	t-Stat	Prob.	t-Stat	Prob.	t-Stat	Prob.
Augmented Dickey-Fuller Test statistic	**-6.31**	**0.00**	**-5.15**	**0.00**	**-6.37**	**0.00**
1% level	-4.14		-4.14		-3.56	
5% level	-3.50		-3.50		-2.92	
10% level	-3.18		-3.18		-2.60	
Independent variables:						
	Coef.	P-value	Coef.	P-value	Coef.	P-value
π_{t-1}	-0.88	0.00	-0.95	0.00	-0.88**	0.00
$D(\pi_{t-1})$			0.07	0.59		
C	2.70	0.00	2.92	0.00	-0.11	0.48
@trend(1992.1)	-0.06	0.00	-0.06	0.00		
Adj R^2	0.42		0.41		0.44	
D.W.	2.00		2.04		1.99	

* D is the difference operator.
** The estimated coefficient of $\pi_{t-1} - \pi^T_{t-1}$.

As before we conducted a sensitivity analysis on quarterly data, letting the number of lags be determined by SIC. The results, reported in the Table 8 below, confirm our previous results.

Table 8:
Sensitivity Analysis of the Existence of Unit Root and Deterministic Trend in the Israeli Inflation Data[a]

Sample: From 92.1 To	ADF t-stat.	Critical value 5%	P-value	No. of Lags (qtrs)	Constant t-value	Trend t-value	Adj. R^2	D.W.
2004.4	-6.25	-3.50	0.00	0	5.03	-4.03	0.42	2.00
2003.4	-5.94	-3.51	0.00	0	4.81	-3.76	0.42	2.00
2002.4	-5.87	-3.52	0.00	0	4.65	-3.12	0.43	1.99
2001.4	-6.10	-3.53	0.00	0	5.04	-3.74	0.47	2.00
2000.4	-5.74	-3.54	0.00	0	4.74	-3.23	0.47	1.99

a. These Results are outcomes of ADF tests that were performed on Eviews5.

In all samples we obtain a significant negative trend, as required by the inflation target regime. We also obtain a significant positive constant term; this is explained, in view of (10), by the fact that here ($-\alpha_2$) is positive.

IV. Seasonally adjusted data.

Since there are indications of possible seasonal variation of inflation between quarters (inflation is on average lower in the third quarter and higher in the fourth) , we checked whether our results are robust to seasonal factors.

Table 9:
. Unit Root tests with seasonally adjusted inflation data*.

Period	ADF stat. (π-seasonally adjusted)	Critical value 5%	P-value	No. of lags	Const. t-value	Trend t-value	Adj. R^2	DW
1955:1-1969:4	-6.89	-3.49	0.00	0	3.43	-0.46	0.44	1.98
1970:1-1984:4	-0.46	-3.49	0.98	0	-0.72	1.34	0.02	2.11
1970:1-1983:4	-3.02	-3.49	0.14	0	-0.72	3.25	0.14	1.60
1970:1-1982:4	-3.26	-3.50	0.08	0	0.10	2.83	0.14	1.83
1970:1-1981:4	-2.83	-3.51	0.19	0	0.21	2.42	0.11	1.87
1970:1-1980:4	-2.42	-3.52	0.36	0	-0.15	2.43	0.09	1.93
1986:1-1991:4	-5.21	-3.51	0.00	0	4.23	-0.19	0.52	2.12
1992:1-2005:1	-5.20	-3.50	0.00	0	4.35	-3.51	0.33	1.94

* The dependent variable is $\Delta\pi_t$.

The results of these computations are reported in table 9, with extra detail for the inflationary period. It can be seen the basic result are unchanged- there still a presumption of a unit root in the inflationary period and trend stationarity in the other periods. The positive trend in the drift in the inflationary period comes out very clearly. It is noteworthy that in the seasonally adjusted data there are no lags in the inflationary era, which suggests that the lags we found earlier are probably due to seasonal factors.

IV. Economic Implications

What is the economic implication of these characterizations of the inflation processes?
There has to be a concern for the level of inflation and not only about its rate of change, as it seems to be the case in the inflationary period in Israel (1973-85). The objective of maintaining the existing inflation plateau, without regard to its absolute level, is bound to be self defeating in a policy game equilibrium, even if it is protected by widespread indexation. The goal of price stability requires the support of fiscal policy, which was lacking in the Israeli inflationary experience.

But fiscal policy alone, without the statement of a inflation target may leave inflation lingering on a moderate inflation step of some 15%-20% (annually), without a reduction to low inflation, as the first years of the post-1985 stabilization suggest. An inflation target regime directed at low inflation, can be an effective way of achieving price stability down the road, as has been the case in the inflation targeting period in Israel (1992-2005).

V. Nominal Anchor and the Other Nominal Variables

We divide the analysis of the relation of the nominal anchor in Israel with other nominal variables into two periods: The first is the inflationary period when there was a tremendous upward trend in inflation, which dominated everything else (on the nominal side), inflation shot up from around 20-30 percent in the early seventies to over four hundred towards the mid eighties. The second is the disinflation process that started with the implementation of the inflation target regime in the early nineties, when inflation decreased from the moderate level of around ten percent to price stability in the beginning of the new century. In this stage, the reduction in inflation was not dramatic and this is reflected in our data. Between these periods there is an in-between period when inflation was stable at around 18-20 percent. We shall confine our analysis to the other two periods when there was a clear trend in inflation.

a. The inflationary era

We have seen that this period was characterized by the absence of a nominal anchor which resulted in an upward trend in inflation. The natural hypothesis for this period, in a chronic inflation economy, is that the inflationary trend dictated the behavior of all other nominal variables. The idea is that the absence of a nominal anchor causes the inflation trend, which is then reflected in the path of all other nominal variables. Furthermore, in the model we postulate that the policymaker resorts to surprise inflation tactic only after the realization of nominal shocks. Accordingly absence of nominal shock in period *t* implies the π_t equals actual (known) π_{t-1} (see Figure 1). The realization of this policy requires the support of the other nominal variables, which implies that the latter should be characterized similarly to the inflation process. To empirically examine this hypothesis we look at the characteristics of the policy variables including the growth of various monetary aggregates, the nominal exchange rate, the nominal wage and various nominal interest rates relative to inflation.

In Tables 3 and 4 we presented empirical evidence to the fact that the inflation process possesses a deterministic trend as well as a non-stationary stochastic trend. Here our null hypothesis is that the growth rates of all nominal variables (including the levels of nominal interest rates) have similar characteristics as those of the inflation series. In order to test these hypotheses we subtract the actual inflation from each series of the nominal variables (so as to form real variables) and examine the characteristics of the constructed series using the unit root test and the cointegration test. Formally, let y denote any nominal variable defined above and let π denote the inflation, our null hypothesis is then that the series (y-π) are all trend stationary (the trend reflects the change in the real variable). Furthermore, each one of the y series is expected to be cointegrated with the inflation series.

In Table 10a we report the unit root test results of the series (y-π), where we impose no restrictions on the lag-structure and let E-views5 chooses it utilizing SIC. Accordingly we performed these tests on the quarterly deviations from the actual inflation of the growth of the money base MB; M_1; M_2; the nominal exchange rate, ER; the nominal wage, W; the interest rates on short term deposits, Rd, and loans, Rc.

Table 10a:
Unit Root tests on deviations of growth rates of nominal variables (including interest rates) from the inflation [a], 1973.1 – 1984.4

Dependent Variable[b]:	ADF t-stat.	Critical value 5%	P-value	No. of Lags[c] (qtrs)	Constant t-value	Trend t-value	Adj. R^2	D.W.
MB-π	-7.53	-3.51	0.00	0	-1.04	0.69	0.55	2.06
M_1- π	-6.81	-3.51	0.00	0	-0.01	-2.38	0.49	1.93
M_2- π	-6.69	-3.51	0.00	0	-1.37	0.52	0.49	2.00
ER- π	-6.43	-3.51	0.00	0	-1.54	1.24	0.46	1.99
W- π	-8.54	-3.51	0.00	2	0.89	0.34	0.74	1.94
Rc- π	-4.93	-3.51	0.00	0	-1.74	2.80	0.32	2.07
Rd- π	-6.12	-3.51	0.00	0	-2.31	-1.16	0.43	2.02

a. These Results are outcomes of ADF tests that were performed on Eviews5. b. MB is the quarterly change in the money base; M_1 is the quarterly change in the M_1; M_2 is the quarterly change in M_2; ER is the quarterly change in the dollar exchange rate; W is the quarterly change in the nominal wage; Rc is the quarterly interest rate on credit; Rd is the quarterly interest rate on deposit. c. The number of lags is determined by the SIC (see E-views5 program).

Note that the first column on the left is the *change* in the dependent variable, so that if we have a negative value for the constant it implies a deterministic trend. The results indicate that the constructed series turn out to be stationary, i.e. the unit root is rejected in all cases. All the nominal variables follow the path of the inflation process, accept for stationary random deviations. Note also that the drift in the monetary aggregates contain negative elements which presumably indicate a lower growth rate than inflation.

The fact that the nominal wage follows the statistical properties of inflation requires an explanation. The former variable embodies inflation expectations and is determined prior to the shock. So according to our model its statistical properties should ordinarily be different from those of inflation, while in fact it is not. However, if the inflation steps are long, then most of the time the rate of change in the nominal wage rate will behave similarly to inflation.

An alternative explanation of the close connection between inflation and nominal wages is related to the distinction between the subjective and objective measures of p. If the public underestimated the actual (1-p) then the nominal wage should behave like inflation.

Returning to the statistics, in Table 10b we documented the results of two versions of the Johansen Cointegrated Trace tests allowing for both a linear (columns 1-3) and a quadratic (columns 4-6) trend in the examined series and an intercept (in the former columns) and a linear trend (in the latter) in the Cointegration equation. We also used a 1 period lag-structure in the first difference equation. The results regarding the cointegration of each of the nominal variable with the inflation indicate that all series but the deposit interest rate in the first version are cointegrated with inflation as is postulated. Note that the results also indicate that there exists only one cointegration relation (see the P-values of the third and sixth columns of Table 10b), i.e. the matrix in the test is less than full rank, which supports the hypothesis that the examined nominal series are non-stationary. The normalized cointegration coefficients should be approximately one, since after allowing for the deterministic trends, the real values of the variables are expected to be preserved. This is approximately so, except for the interest rates whose real value seems to have been eroded in the course of the inflation process.

Table 10b:
The results of the Johansen Cointegrated Trace Test[a] of selected nominal variables with the inflation in Israel, 1973.1 1984.4

Nominal variables:	H_0: No Cointegration Equation (CE) With linear deterministic trend		H_0:At Most 1 CE With linear deterministic trend	H_0: No Cointegration Equation (CE) With quadratic deterministic trend		H_0: At Most 1 CE With quadratic deterministic trend
	P-value	Normalized Cointegrating coefficient[b]	P-value	P-value	Normalized Cointegrating coefficient[b]	P-value
MB	0.0000	0.87 (0.11)	0.2293	0.0003	1.06 (0.17)	0.6178
M_1	0.0000	1.03 (0.10)	0.0761	0.0008	0.93 (0.16)	0.7815
M_2	0.0003	0.87 (0.08)	0.1408	0.0009	1.15 (0.16)	0.8087
ER	0.0017	0.93 (0.07)	0.0930	0.0148	1.00 (0.13)	0.9615
W	0.0000	0.90 (0.04)	0.1013	0.0000	0.91 (0.08)	0.9919
Rc	0.0007	0.71 (0.06)	0.5863	0.0000	0.69 (0.10)	0.0006
Rd	0.1004	0.83 (0.16)	0.4909	0.0149	0.40 (0.16)	0.3383

a. The tests were performed allowing for a linear trend and for a quadratic trend in the series and an intercept in the cointegration equation, and a lag-structure of (1 1). b. The normalized cointegrating coefficient of the nominal variable where the quarterly inflation is the dependent variable (standard error in parentheses).

b. The inflation target regime

In the period since 1992 Israel adopted the inflation target regime which was supposed to guide inflation to price stability. The reduction of inflation in this stage was not dramatic: from around 10% annually to 1%-3%. Again we took the declining actual inflation as representing the government's commitment of reaching price stability. Accordingly, our null hypothesis with respect to each the series $y - \pi$ is similar to that in the previous section, that is, that they are stationary and that the deviation of the nominal variable from the inflation is random. The results are presented in table 11.

It can be seen that in general the results indeed support our hypothesis regarding the stationarity and the similarity of the characteristics of the nominal variable series with those of the inflation process. The exception is the nominal wage, which turns out non-stationary.

These results confirm our basic hypothesis that in this period there existed an effective anchor for inflation.

Table 11:
Unit Root tests on deviations of growth rates of nominal variables (including interest rates) from inflation [a], 1992.1 – 2005.1

Dependent Variable:	ADF t-stat.	Critical value 5%	P-value	No. of Lags[b] (qtrs)	Constant t-value	Trend t-value	Adj. R^2	D.W.
MB-π	-9.53	-3.50	0.00	0	2.23	-0.67	0.65	2.13
M_1- π	-7.47	-3.50	0.00	0	1.45	0.38	0.51	1.83
M_2- π	-6.70	-3.50	0.00	0	4.77	-3.06	0.45	1.93
ER- π	-7.10	-3.50	0.00	0	-0.76	0.54	0.48	2.01
W- π	-2.89	-3.50	0.17	3	1.32	-0.78	0.65	1.98
Rc- π	-5.68	-3.50	0.00	0	3.56	0.92	0.37	2.07
Rd- π	-5.71	-3.50	0.00	0	0.38	1.60	0.37	2.02

a. These Results are outcomes of ADF tests that were performed on Eviews5.

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Appendix A: Inflation Targets

The examined period in Israel includes the implementation of a disinflation policy, where the policy maker (say, *other* than the monetary policy makers) had to decide how intensive the disinflation policy should be. The current inflation is, say, π^T_0 and eventually the goal is to maintain price stability, π^T_N, at some future date *N*. The policy maker has to determine the time path for the inflation targets $\{\pi_t^T\}_{t=1}^N$. This appendix deals with the choice of the path of the inflation targets in the disinflation period.

Given the inflation targets and the objective loss function (12), the *monetary* policy makers minimize this loss function by choosing the inflation series, $\{\pi_t\}$. Their choice of the inflation is given in (13), which can be rewritten as follows:

$$\text{(A:1)} \qquad \pi_t = (1-p)\frac{b}{\beta_0+\beta_2} + \frac{\beta_0}{\beta_0+\beta_2}\pi_{t-1} + \frac{\beta_2}{\beta_0+\beta_2}\pi_t^T + u_t$$

where $u_t \equiv \pi_t - E_t(\pi_t$ with $E_t(u_t) = 0$. One can verify that by recurrent substitution for lagged inflation in (A:1) and given the time path of the inflation targets, $\pi^T_0, \pi^T_1, \ldots, \pi^T_N$, we can express the current inflation in terms of the inflation targets as follows

$$\text{(A:2)} \qquad \pi_t = c_0 \sum_{j=0}^{t-1} (c_1)^j + (1-c_1) \sum_{j=0}^{t-1} (c_1)^j \pi^T_{t-j} + (c_1)^t \pi_0 + \sum_{j=0}^{t-1} (c_1)^j u_{t-j}$$

where

$c_0 = (1-p)\dfrac{b}{\beta_0 + \beta_2}$, $c_1 = \dfrac{\beta_0}{\beta_0 + \beta_2}$, and π_0 is the inflation at the start of the disinflation period.

We can think of the *other* policy maker, which in setting the path of the inflation targets actually chooses an *optimal* policy, as a solution to the following control problem. He/she chooses the series $\{\pi^T_t\}$ to minimize the quadratic objective function (12), subject to the choice of the monetary policy maker (A:2). In order to solve this control problem, we substitute from (A:2) into (12), and get the objective function as a function of π^T_t, $t = 1,2,\ldots,N$.

Note that in solving this problem we can apply the Envelope Theorem such that only the last term in (12) is relevant for the first-order conditions. Indeed, after the substitution of (A:2) into (12), the last term in (12) becomes

$$\text{(A:3)} \qquad \frac{\beta_2}{2}\left[c_0 \sum_{j=0}^{t-1} (c_1)^j + (c_1)^t \pi_0 + (1-c_1) \sum_{j=0}^{t-1} (c_1)^j \pi^T_{t-j} + \sum_{j=0}^{t-1} (c_1)^j u_{t-j} - \pi^T_t \right]^2$$

The first-order conditions for the optimization for $t = 1,2,\ldots,N$ are

$$\text{(A:4)} \qquad c_0 \sum_{j=0}^{t-1} (c_1)^j + (c_1)^t \pi_0 + (1-c_1) \sum_{j=0}^{t-1} (c_1)^j \pi^T_{t-j} + \sum_{j=0}^{t-1} (c_1)^j u_{t-j} - \pi^T_t = 0$$

Rewriting (A:4) yields

$$(\text{A:5})\ \pi_t^T - \pi_{t-1}^T = -c_1\pi_{t-1}^T + (1-c_1)\sum_{j=2}^{t-1}(c_1)^{j-1}\pi_{t-j}^T + c_0\sum_{j=2}^{t-1}(c_1)^{j-1} + (c_1)^{t-1}\pi_0 + u_t + \sum_{j=1}^{t-1}(c_1)^j u_{t-j}$$

There are few characteristics of the series of the inflation targets that come up from this solution (A:5):

i) If $c_1 > 0$, then there exists a constant in the difference equation, $\pi^T{}_t - \pi^T{}_{t-1}$, indicating a deterministic time trend in the inflation target series.

ii) If $c_1 > 0$, then the inflation target series is stationary.

iii) However, As β_0 gets smaller (c_1 gets smaller), the effect of $\pi^T{}_{t-1}$ on the difference $\pi^T{}_t - \pi^T{}_{t-1}$ gets weaker and apparently as β_0 approaches zero (c_1 approaches zero) this effect completely disappears. That is, when $\beta_0 = 0$, the inflation target series has a unit root. Indeed, $\beta_0 = 0$ means (see the objective function) that the policy maker does not care about how fast inflation is reduced in the period of the disinflation policy.