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ESTIMATING THE NAIRU FOR ISRAEL, 1992–2013

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Abstract

We use a state space model to estimate a time-varying Non-Accelerating Inflation Rate of Unemployment (NAIRU) for Israel for the period 1992–2013. The NAIRU is an unobserved variable and can be estimated by the Kalman filter, exploiting the assumed relationship between the NAIRU and several observed variables such as the inflation and unemployment rates. We assume a forward looking Phillips curve, according to which actual inflation is influenced by expected inflation (which is also unobserved), by the unemployment gap (the gap between the actual unemployment rate and the NAIRU) and by real exchange rate changes. As a proxy for expected inflation, we use data on inflation expectations derived from the bond market in Israel ("breakeven inflation"). This enables us to avoid making an assumption regarding the formation of expectations, especially avoiding the usual practice of assuming adaptive expectations and using lags of inflation as proxies for inflation expectations. We find that the estimated NAIRU is fairly variable and explains a great deal of the low frequency dynamics of the actual unemployment rate. For example, from 2004 to 2013, the actual unemployment rate declined by 7.7 percentage points. Our estimates suggest that 6.2 percentage points (most of the reduction) were due to a decline in the NAIRU. This is probably a result of policy steps to encourage employment that were implemented by the government beginning in 2002. We also found that the estimated NAIRU fits very well in a Beveridge curve, and thus helps to identify the negative relationship between job vacancies and unemployment, as one would expect from a measure of structural unemployment.

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1. INTRODUCTION

Between 2004 and 2013, the unemployment rate in Israel declined from 13.5 percent to 5.8 percent without any noticeable upward pressure on inflation. One clear reason for this is the prolonged decline in the relative price of imports, which moderated imported inflation. Another possible explanation for the lack of inflationary pressures is a decline in the Non-Accelerating Inflation Rate of Unemployment (the NAIRU—also frequently referred to as the natural rate of unemployment). The decline in the NAIRU is probably a result of government policy to encourage employment, which started in the years 2002–03 and was implemented in the years that followed. To assess this possibility we estimate a time-varying NAIRU for Israel over the period 1992:Q1–2013:Q4.

The time-varying NAIRU is, of course, an unobserved variable. We therefore specify it as a latent variable and use the Kalman filter method to evaluate it. As is usual practice in the literature[1], the NAIRU is inferred by exploiting the information available in an expectations-augmented Phillips curve—that is, by estimating the relationship between the inflation rate, its lags, supply shock variables and the gap between the actual and the natural rates of unemployment (henceforth, the unemployment gap). An additional important unobserved variable that appears in the Phillips curve relationship is the public's inflation expectations. Usually, inflation expectations are proxied by distributed lags of inflation. Here we instead use an estimate of one-year inflation expectations that is derived from the capital market. Specifically, we use the difference between the yields to maturity on two one-year bonds that differ from each other only in that one is unindexed while the other is indexed to the CPI. (This difference is also known as "breakeven inflation"). As a check for robustness we estimated the model under the assumption of adaptive expectations as well.

The estimates of the inflation equation show that the gap between the actual unemployment rate and the NAIRU (i.e., the unemployment gap) indeed has a meaningful negative effect on the inflation rate. An unemployment gap of one percentage point decreases the inflation rate, with a one year lag, by 0.18 percentage points. Such a change also temporarily reduces inflation, for one quarter after the change, by 0.29 percentage points. With respect to supply shocks, a change in the relative price of imported goods has a sizable and prolonged effect on inflation. An increase of one percent in the relative price of imported goods raises inflation by a cumulative effect of 0.14 percent over three quarters.

We also find that the estimated NAIRU is fairly variable and explains a great deal of the low frequency dynamics of the actual unemployment rate. For example, from 2004 to 2013, actual unemployment declined by 7.7 percentage points. Our estimates suggest that 6.2 percentage points (most of the reduction) were due to a decline in the NAIRU. This decline seems to be a result of a government policy to encourage employment that was implemented beginning in 2002–03, and included measures of prolonged reduction in

[1] See, for example, Gordon (1997), Laubach (2001) and Turner et al. (2001).

social security payments, in unemployment benefits and in the rate of income tax on wages.[2]

In order to assess the usefulness of the estimated NAIRU, we examined its ability to explain movements in the relationship between the unemployment rate and the vacancy rate, reflected in the Beveridge curve (as specified by Dickens (2009)). The results suggest that the estimated NAIRU (from the Phillips curve) makes a sizable and significant contribution to the identification of the relationship between vacancies and unemployment. That is, the estimated NAIRU behaves as one would expect from a measure of structural unemployment.

The paper is organized as follows. In Section 2, we present the model and the estimates and check the robustness of the results with respect to the choice of the measure of inflation expectations, the exclusion of the housing component from the consumer price index, and the estimation period. In Section 3 we check the usefulness of the estimated NAIRU as a "driver" of the Beveridge curve, and Section 4 concludes.

2. THE PHILLIPS CURVE AND THE NAIRU

a. The model

Following the relevant literature[3] we start with a Phillips curve of the form:

$$(1) \qquad \pi_t = \pi^e_{t+1} - \sum_{i=1}^{ku} \beta_i (u_{t-i} - u^n_{t-i}) + \sum_{i=1}^{kz} \gamma z_{t-i} + e^{\pi}_t ,$$

where π_t is the inflation rate in quarter t, π^e_{t+1} is the expected rate of inflation in the following quarter, u_t is the unemployment rate, u^n_t is the NAIRU, z_t represents supply shocks and e^{π}_t is an i.i.d. shock with a mean zero and standard deviation σ_{π}. ku and kz represent the number of lags of the unemployment gap and the supply shocks, respectively, which are determined empirically.

The preceding specification contains two unobserved variables, π^e_{t+1} and u^n_t. It is commonly assumed in the related literature that expected inflation can be represented by several lags of inflation, that is:

[2] See Appendix 3 and the chapters on the labor market in the Bank of Israel Annual Reports for 2002 and 2003.

[3] See, for example, Laubach (2001) and the references there. Note that Laubach specifies expected inflation as inflation expected in $t-1$ to t, while we specify it as inflation expected from period t to the following period $t+1$, as in the spirit of the new Keynesian model.

$$\pi^{e}_{t+1} = \sum_{i=1}^{k} \alpha_i \pi_{t-i} \quad (2)$$

If we assume that the weights of the inflation lags sum to one (that is $\sum_{i=1}^{k} \alpha_i = 1$), then the Phillips equation in (1) can be written as:

$$\Delta\pi_t = (\alpha_1 - 1)\Delta\pi_{t-1} + (\alpha_1 + \alpha_2 - 1)\Delta\pi_{t-2} + \ldots + (\alpha_1 + \alpha_2 + \ldots + \alpha_{k-1} - 1)$$
$$\Delta\pi_{t-k+1} + \sum_{i=1}^{ku} \beta_i (u_{t-i} - u^{n}_{t-i}) + \sum_{i=1}^{kz} \gamma z_{t-i} + e^{\pi}_{t} \quad (3)$$

The above specification has been used in most existing studies, including Friedman and Suhoy (2005) who estimate a time-varying NAIRU for Israel for the period 1987:Q1 to 2001:Q4.[4]

Instead of assuming adaptive expectations, as in (2), we will use the data on one-year ahead "breakeven inflation" derived from the capital market as a proxy for inflation expectations.[5] That is, we estimate equations of the form:

$$\pi_t - \hat{\pi}(market)_{t,t+4} = \sum_{i=1}^{ku} \beta_i (u_{t-i} - u^{n}_{t-i}) + \sum_{i=1}^{kz} \gamma_i z_{t-i} + e^{\pi}_{t}, \quad (4)$$

where $\hat{\pi}(market)_{t,t+4}$ is one-year ahead breakeven inflation derived from government bonds traded in the bond market.[6]

We express the unemployment rate (u_t) as the sum of two unobserved components, the NAIRU (u^{n}_{t}) and the unemployment gap ($ugap_t$):

[4] This paper differs from Friedman and Suhoy (2005) in several respects: (1) Different periods of estimation; (2) As a proxy for inflation expectations we use data from the capital market; (3) We use different variables as proxies for supply shocks; (4) We include drift in the NAIRU dynamics which captures a potentially stochastic but stationary trend in the NAIRU; (5) We abstain from including data on output, which they use in order to improve the identification of the NAIRU (through Okun's law). This is because such a procedure requires estimation of potential output which is itself an unobserved variable and subject to uncertainty. It should be noted that the estimated NAIRU in our model is fairly variable, while the NAIRU estimated by them is nearly constant.

[5] The breakeven inflation rate is calculated as the difference between the nominal rate of return on a one-year unindexed bond and the real rate of return on a one-year indexed bond. This gap is an estimate of expected inflation plus a risk premium that we assume to be constant.

[6] Note that the data that we derive from the bond market represent expectations for four quarters ahead, while the model is specified and estimated with quarterly data. This is a limitation of the data on expected inflation. However, such specification may be preferable since in reality it is probable that the relevant horizon for the expectations that affect players' behavior is much longer than one quarter ahead.

(5) $$u_t = u_t^n + ugap_t .$$

In order to distinguish empirically between the NAIRU and the unemployment gap we need to model the data generating process of each component in an independent manner. It is common to model the data generating process of the NAIRU as a random walk with a stationary drift (g_t)—that is, assuming that the NAIRU is a stationary variable. In some cases the changes in the NAIRU are specified as white noise, namely φ is imposed to be zero in equation (7). However, such an assumption limits the potential low frequency variability of the estimated NAIRU and increases the high frequency variability.[7] We allowed an autoregressive specification for the changes in the NAIRU as follows:

(6) $$u_t^n = u_{t-1}^n + g_{t-1}$$

(7) $$g_t = \varphi g_{t-1} + e_t^g .$$

Here, $-1 \le \varphi \le 1$ measures the degree of persistence of the change in the NAIRU, and e_t^g is an i.i.d. shock with mean zero and standard deviation σ_g. Note that we do not impose inertia on the changes in the NAIRU, but rather, we allow it to be determined by the data. The inclusion of the drift captures a potential stochastic but stationary trend in the NAIRU. This assumption is consistent with the approach of Turner, et al. (2001) and implies that changes in the NAIRU are stationary and that in the long run the NAIRU is constant. Laubach (2001) also included a stochastic drift in the natural rate, but assumed that the drift follows a random walk ($\varphi = 1$)—implying that only the second difference of the NAIRU is stationary (an assumption that does not seem reasonable for Israel's unemployment rate).[8]

As for the unemployment gap, following Laubach (2001), we assume that it follows an autoregressive process of the form:

(8) $$ugap_t = \delta_1 ugap_{t-1} + \delta_2 ugap_{t-2} + e_t^{ugap} ,$$

where $-1 \prec \delta_1 + \delta_2 \prec 1$, and e_t^{ugap} is an i.i.d. shock with mean zero and standard deviation σ_{ugap}.

[7] Such an assumption usually leads to an estimated NAIRU that tends to vary considerably from quarter to quarter. Friedman and Suhoy (2005) assumed that $\varphi = 0$ which is a private case of the specification in equations (6) and (7). As we shall shortly see this assumption is rejected by the data.

[8] This specification was largely motivated by a sample of European countries in which unemployment was trending up over the sample period used by Laubach (2001).

We assume that all shocks are mutually uncorrelated and normally distributed. The system of five equations (4)–(8) can be estimated by means of the Kalman filter in order to obtain maximum likelihood estimates of the relevant parameters and time series estimates of the unobserved variables including the NAIRU.

b. Estimation of the model using "breakeven inflation" as a proxy for inflation expectations

To estimate the model, we need to first choose relevant variables that can proxy the effect of supply shocks in the inflation equation. The change in the relative price of imported goods is a common choice for that purpose. Let Δpim_t stand for the rate of change in the world's price (in dollar terms) of Israeli imports, and let Δe_t represent the rate of change in the shekel/dollar exchange rate, and let $\Delta epim_t = \Delta e_t + \Delta pim_t$ represent the rate of change of Israeli import prices in shekel terms.[9]

In the estimation of the system we specified a Phillips curve in the form of equation (4)[10] where on the right hand side we allowed four lags[11] of each of the following variables: $\{ugap_t, \Delta epim_t, \pi_t\}$. We also included the current value of $\Delta epim_t$ [12] and the acceleration in the current value of the world price of imports (see equation 4a). At first, we did not impose any restrictions on the lags. We then gradually imposed restrictions that

[9] All the variables in the estimated equations, except Δe_t, are seasonally adjusted.

[10] For inflation (π) we use the change in the CPI excluding housing services, fruit and vegetables. We exclude these components due to empirical considerations (better fit, order of magnitude and significance of the coefficients) but it can be justified by additional considerations as follows. The fruit and vegetables component is known for high volatility and irregularity. As for housing services, in 1999 the CBS changed the method of deriving this data (until 1999 house prices were used as a proxy, from 1999 data on rents began to be used). Furthermore, until 2007, the housing component was almost fully linked to the shekel-dollar exchange rate. From 2007 onward, we observe a notable disruption in this linkage. The fact that the proxy for inflation expectations that we use (from the capital market) includes housing, fruit and vegetables, but the inflation data doesn't, reveals that in fact the inflation expectations included in the model have measurement errors. This increases the variance of the residuals in the inflation equation but does not bias the parameter estimates (note that inflation expectations appear on the left hand side of the estimated equation). In any case, in subsection 2.3.1 we present the results when we use the CPI including housing and fruit and vegetables. As can be seen, the results concerning the evolution of the estimated NAIRU are quite similar.

[11] We choose to limit the number of lags to four to save degrees of freedom, and also based on experience abroad of estimating inflation equations. However, note that when inflation lags are used to proxy expectations, up to twelve lags are sometimes needed (see, for example, Gordon (1997)).

[12] By including this variable we assume that it is exogenous to domestic inflation. For a small open economy, the exogeneity of the world price of imports is not questionable. As for the exogeneity of the exchange rate changes, we performed the exogeneity test of Durbin-Wu-Hausman. For the test we added several instrumental variables: the current value and four lags of the LIBOR rate and four lags of the Bank of Israel interest rate. The result of the test is that the null hypothesis (that the change in the exchange rate is exogenous) is not rejected for any significance level up to 19%.

seemed to characterize the data[13], and settled on the following empirical specification of the Phillips equation:[14]

$$\pi_t - \hat{\pi}(market)_{t,t+4} = \beta_1 \Delta ugap_{t-1} + \beta_2 ugap_{t-4} + \gamma_1 \Delta rq_t + \gamma_2 (\Delta rq_{t-1} - \Delta rq_{t-2})/2 + \gamma_3 \Delta\Delta pim_t + \gamma_4 \pi_{t-1} + e_t^{\pi} \quad (4a)$$

where $\Delta rq_t = \Delta epim_t - \pi_{t-1}$ represents the change in the import price in real terms. The inclusion of inflation lag in the specification of Δrq_t is used in order to keep the exogenous nature of Δrq_t in equation (4a).

Note that the unemployment gap affects inflation with a lag of four quarters. The first lag of the change in the unemployment gap also affects inflation. The significant effect of the change in the unemployment gap is a common result in such kind of equations, and implies that the effect of the unemployment gap on inflation is not necessarily linear[15] (the sign of both β_1 and β_2 is expected to be negative). The term $\Delta\Delta pim$ represents the acceleration (or deceleration) in the world price of imports, and is included in the equation due to its statistical significance. Its inclusion also improves the significance of other variables. The inclusion of $\Delta\Delta pim$ is quite common in the relevant literature and serves as an additional way to reflect the supply shock effect[16] (the sign of γ_3 is expected to be positive). For similar reasons (significance of the estimate and improved significance of other estimates) we included the change in the first lag of inflation, $\Delta\pi_{t-1}$, which might capture an effect of inflation persistence beyond the effect of inflation expectations. Such persistence is possibly due to staggered price and wage adjustments (the sign of γ_4 is expected to be positive).

The system which is represented by equations (4a) and (5)–(8) (henceforth: Model 1) was estimated by the Maximum Likelihood method, for the period 1992:Q1–2013:Q4. This yielded significant estimates[17], with the correct signs for all the parameters apart from the standard deviation σ_g. Insignificance of the standard deviation of g_t is a well-known problem with the application of the Kalman filter to such a system: the estimate of σ_g

[13] We were also guided by the concept that in the long run (that is, when inflation is constant) the inflation rate should be orthogonal to real variables such as the level of change in the real price of imports. Note that in the long run, all the variables on the right hand side of equation (4) should be zero, and this restriction should be imposed on the parameters. In particular, in the long run we should have: $\pi = \Delta e + \Delta pim$.

[14] In Appendix 2 we present the estimates of the unrestricted version of the inflation equation.

[15] This effect is termed the "speed limit effect"—see, for example, Turner (1995). The implication is that even when the unemployment rate is above the NAIRU, a large decline in the unemployment rate can increase inflation.

[16] For example, Laubach (2001) adds lags in the acceleration of the exchange rate and of commodities prices to reflect the effect of supply shocks.

[17] Here and in what follows "significant" means significance level of 5%.

might be insignificant and tends to zero even when its "true" value is positive.[18] To identify the variance of σ_g the following procedure was implemented: we estimated the system for various values of the variance ratio σ_g^2/σ_π^2 in the range {0.01 to 0.20} (in the unconstrained estimation the estimated value was 0.07 but insignificant). Inspection of a graph of the likelihood function reveals that in the above range, the likelihood function is rather flat, which makes the identification of σ_g^2/σ_π^2 very difficult. So we added another criterion: the maximum t-value of the unemployment gap in the Phillips curve (the parameter β_2). The "preferred" ratio under the above criterion is 0.03.

The resulting parameter estimates are presented in the second column of Table 1. We note first that all parameters have the correct sign and are statistically significant. The estimates of equation (4a) imply that an unemployment gap of one percent decreases inflation, other things equal—including expected inflation—with a one year lag by 0.179%. Such a change also temporarily reduces inflation for one quarter after the change by 0.286%. As for the relative price of imports, a one percent increase acts toward a cumulative rise of 0.143% in inflation in three quarters. Other factors with significant effect on inflation are the acceleration in the world's imports and the lagged change of inflation.

The estimate of φ in equation (7) is high (0.932), which means that the changes in the estimated NAIRU are highly persistent. Therefore, the assumption of random walk by Friedman and Suhoy (2005) is strongly rejected.[19] The estimates of the parameters in the unemployment gap equation reveal that the unemployment rate is highly persistent and converges monotonically to the NAIRU.

[18] This is known as the "pile-up" problem. See Gordon (1997) and Laubach (2001) and the references there for discussions on that issue.

[19] A Dickey-Fuller test with respect to the changes on the NAIRU shows that the hypothesis that $\varphi = 1$ is rejected for any reasonable significance level. Specifically, the null hypothesis that the NAIRU is I(1) is not rejected for any reasonable significance level (a p-value of 0.899). In contrast, the hypothesis that the estimated NAIRU is I(2) (or that the changes of the NAIRU are I(1)) is rejected for any significance level (the statistics get the value of -3.268 and the p-value is 0.001). We should also note that the unemployment rate is random walk (according to Dickey-Fuller tests, the hypothesis that the unemployment rate is I(1) is not rejected, and that it is I(2) is rejected). Thus the result that φ is less than 1—that is, that the NAIRU is a random walk, means that the unemployment gap is stationary.

Table 1
The estimated model under three specifications, 1992:Q1–2013:Q4
(*t*-values in parentheses)

Parameters	Model 1: Inflation expectations derived from capital market, proxying expected inflation (equation 4a)	Model 1: Inflation includes housing and fruit and vegetables components	Model 2: Inflation lags proxying expected inflation (equation 3a)
	Phillips equation		
β_1	-0.286 (2.2)	-0.00 (0.0)	-0.271 (2.0)
β_2	-0.179 (2.8)	-0.281 (2.6)	-0.142 (2.4)
γ_1	0.079 (3.7)	0.147 (5.4)	0.135 (6.3)
γ_2	0.064 (2.2)	0.042 (1.2)	0.026 (1.8)
γ_3	0.047 (2.6)	0.002 (0.1)	0.084 (4.4)
γ_4	0.327 (4.8)	0.275 (4.5)	
α_i	$\alpha_1 = 1$	$\alpha_1 = 1$	$\alpha_1 = 0.792$ (11.2) $\alpha_2 = -0.129$ (1.9) $\alpha_3 = 0.080$ (1.0) $\alpha_4 = 0.259$
	NAIRU-variables equation (equation 7)		
φ	0.932 (19.9)	0.911 (16.3)	0.918 (16.6)
	Unemployment gap equation (equation 8)		
δ_1	1.122 (11.2)	1.158 (12.0)	1.185 (14.2)
δ_2	-0.266 (3.2)	-0.276 (3.4)	-0.277 (3.7)
	Variance		
σ_π^2	0.0044^2 (11.3)	0.0057^2	0.0048^2 (12.0)
$\frac{\sigma_g^2}{\sigma_\pi^2}$	0.03	0.023	0.019
σ_{ugap}^2	0.0044^2 (11.0)	0.0044^2 (11.2)	0.0045^2 (11.1)

Figure 1
Actual Unemployment and Estimated NAIRU of the Model of Equations (4a) and (5)–(8)

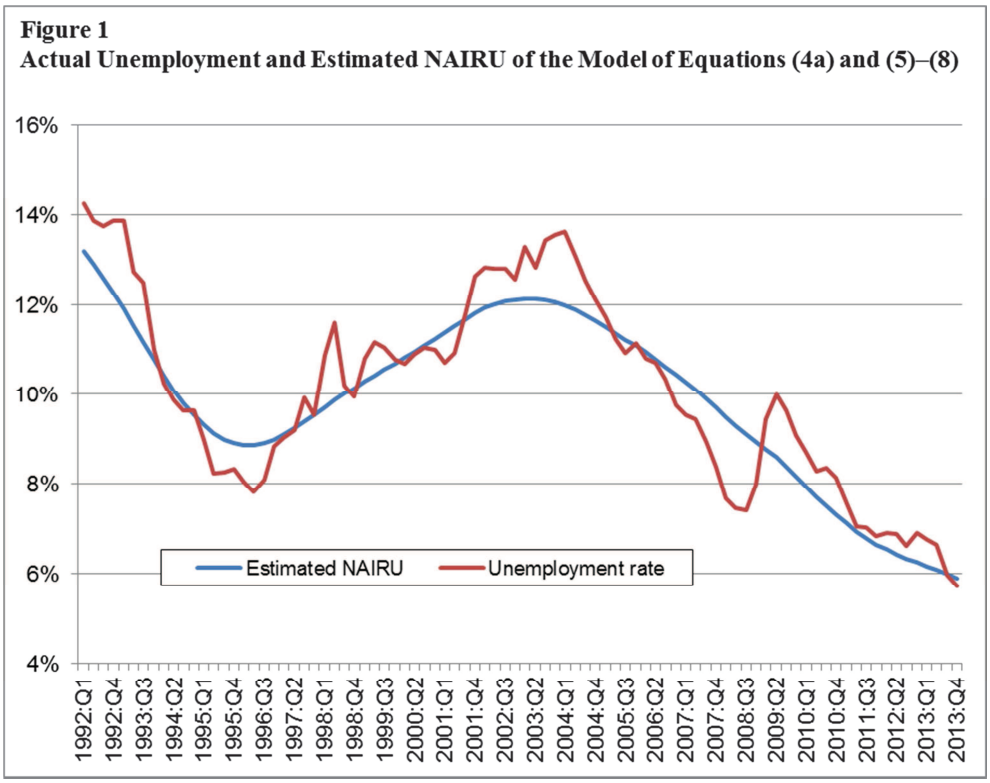


Figure 1 presents the unemployment rate and the NAIRU that is estimated by Model 1. We can see that the NAIRU is positively correlated with low-frequency dynamics of the unemployment rate.

At the beginning of the period (1992) the actual unemployment rate and the NAIRU were both very high (14% and 13%, respectively). This reflects the massive immigration from countries of the former Soviet Union which occurred during the years 1989 to 1992 (and continued afterward as well, though at a lower rate). Both rates then declined, to a local minimum of 9% for the unemployment rate and 8% for the NAIRU, in 1997. This seems to reflect, at least partly, absorption of the immigrants (of 1989 to 1992) in employment. From 1997, the unemployment rate and the NAIRU increased gradually to a sample peak of around 13.5% and 12%, respectively, in 2003. An important development that may explain the increase of the NAIRU is the massive contraction of the construction industry that took place in that period. The workers that were discharged from the construction industry found it difficult to find jobs in their profession, which may explain an increase in structural unemployment (mismatch). Another possible reason is an increase in the participation rate that occurred during the above period, and an increase in the labor

tax rate that started in 1994 and continued until 2002 (the rate was gradually increased from 26% to 31%).

From 2004 onward, the NAIRU has been on a course of persistent decline, continuing even in the crisis of 2008 and 2009, despite a dramatic increase in the actual unemployment rate at that time. An important factor that impacted on the NAIRU's decline beginning in 2004 is government policy adopted in 2002, and which continued afterward, with the goal of encouraging employment and reducing structural unemployment. This policy was reflected in, among other things, stricter criteria for receiving unemployment benefits, reduction of income support payments by the National Insurance Institute, steps taken to reduce the number of foreign workers, and a reduction in tax rates on wages.[20] A further indication of the decline of the NAIRU beginning in 2006 is the leftward shift of the Beveridge curve (see Section 3 below). In Appendix 3 we expand the discussion on the structural changes in the labor market in 2002–07, and their possible contribution to the decline of the NAIRU.

c. Robustness of the findings with respect to the choice of: the inflation variable, the inflation-expectations proxy and the starting point of the estimation period.

1. *Robustness of the findings with respect to choice of inflation variable*

For reasons noted above, the inflation variable in the model is based on the CPI excluding housing, fruit and vegetables—as opposed to the estimate for expected inflation derived from the capital market, which refers to the CPI including those components. To examine the robustness of the findings with respect to the choice of inflation variable, we estimated the model with inflation based on the CPI including these components as well. The estimates are presented in the third column of Table 1. As can be seen, there is some difference in some of the estimates, but the overall picture is similar, and the estimated NAIRU is similar in its development to the rate estimated using inflation excluding housing, fruit and vegetables.[21] The coefficient of changes in unemployment was near null, but in contrast the coefficient of the fourth lag of unemployment increased in its absolute value (from -0.179 to -0.281). It should also be noted that the contemporaneous effect of changes in the real price of imports strengthened markedly. This reflects the effect of the shekel's depreciation on the price of residential rents, a relation that was especially notable in the period prior to 2007. In contrast, the contemporaneous effect of the acceleration in dollar prices of imports contracted.

[20] Between 2002 and 2012, there was an extended decline in the average tax rate on wages, from 31 percent to 23 percent.

[21] Both estimates of the NAIRU are highly correlated with actual unemployment, and in both we note a prolonged decline since 2004. The estimated NAIRU in the above three alternatives is presented in Figure A1 in Appendix 1.

2. *Robustness of the findings with respect to the estimation of inflation expectations*

In the previous section we used bond market data on the yield spread as a proxy for inflation expectations. In order to examine the robustness of the findings relative to this choice, we also estimated the model using the generally accepted approach, in which inflation lags are used as a proxy for inflation expectations. That is, we repeated the model-estimation process with equation (3) replacing equation (4). With regard to lags of inflation in the inflation equation, in view of the experience in Israel and abroad, a relatively large number of lags may be required. Therefore, in the beginning we allowed eight lags (and four lags for the other variables) and repeated the process described for the estimation of equation (4). The empirical specification that was obtained for equation (3) is presented as equation (3a) below:

$$\begin{aligned}\Delta\pi_t &= (\alpha_1 - 1)\Delta\pi_{t-1} + (\alpha_1 + \alpha_2 - 1)\Delta\pi_{t-2}. + (\alpha_1 + \alpha_2 + \alpha_3 - 1)\Delta\pi_{t-3} \\ &+ \beta_1\Delta ugap_{t-1} + \beta_2 ugap_{t-4} + \gamma_1\Delta rq_t + \gamma_2(\Delta rq_{t-1} + \Delta rq_{t-2})/2 \\ &+ \gamma_3\Delta\Delta pim_t + e_t^{\pi}\end{aligned} \tag{3a}$$

The results of the estimation of the model (henceforth, Model 2) are presented in the last column of Table 1. As can be seen, four lags of inflation were found to be relevant[22] (that is, three lags of changes in inflation). Note that most of the effect is attributed to the first and fourth lags of inflation (coefficients of 0.79 and 0.26, respectively). The results in the remainder of the model's parameters are similar to those estimated for Model 1. The NAIRU estimation under this alternative is also similar to that yielded by Model 1 (see Figure A1 in Appendix 1). Note that the contemporaneous effect of supply factors (expressed by γ_1 and γ_3) is stronger here than in those of equation (4a). It is likely that this reflects the effect of these factors on inflation expectations derived from the capital market, which cannot be expressed in an estimate of expectations based solely on lags of inflation. This finding likely supports the preference of expectations derived from the capital market in estimating inflation expectations.

3. *Robustness of the results with regard to the starting point of the sample period*

At the beginning of the estimation period, in 1992, the unemployment rate was at its highest point, impacted on by a wave of immigration from the Soviet Union that began in 1989. Later, it declined to a local bottom level in 1996. A similar development characterizes the estimated NAIRU. In order to examine the robustness of the results with regard to the

[22] The third lag is not significant (its t value is slightly less than 1) but we retain four lags because when we allow five lags the fourth is significant. The estimated coefficient gets the value of 0.254 with t value of 3.7. Note also that the fifth lag is not significant and the sum of the coefficient of the first four lags is 0.996 (that is, we don't have to impose the homogeneity constraint).

starting point of the sample period, we conducted two alternative estimations of the model. Once beginning in 1988, before the effect of the immigration from the Soviet Union was seen, and once beginning in 1996, after the exhaustion of the effect of the immigration. Figure A2 in Appendix 1 presents the actual unemployment rate and the NAIRU estimated for the three alternatives in terms of starting date: the baseline alternative from 1992, the 1988 alternative and the 1996 alternative. As can be seen, the results for the overlapping period are similar. Likewise, in all the cases, beginning in 2004 the NAIRU is on a consistent trend of decline. That is, the results with regard to the development of the estimated NAIRU are not sensitive to the starting point of the estimation.

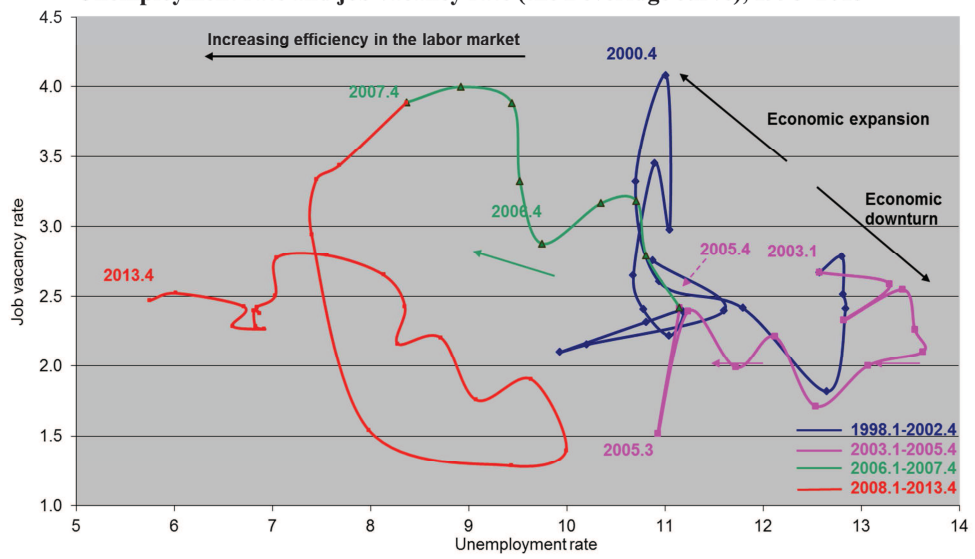
3. THE ESTIMATED NAIRU AND THE BEVERIDGE CURVE

The decline in the NAIRU, at least since 2006, might reflect, among other things, improved efficiency of matching between job seekers and job suppliers. This possibility was noted and discussed in the Bank of Israel Annual Report for 2011. Figure 2 presents a scatter plot of the unemployment rate and the job vacancies rate from 1998 to 2013.[23] As can be seen, in the period 1998 until 2006 the points show a curve that declines from left to the right. This curve is termed in the literature as the "Beveridge curve". It describes a negative correlation between the unemployment rate and the vacancy rate. When the economy is in a recession, we expect to see a low vacancy rate and high unemployment rate. The opposite in times of prosperity—when the unemployment rate is low we expect to see a high vacancy rate. As can be seen, since 2006, we can note a shift to the left of the Beveridge curve, which can be interpreted as an improvement in the matching process between vacancies and unemployed workers.[24]

[23] It is a similar diagram to the one which appears in the Bank of Israel 2011 Annual Report.

[24] This also means that the same correlation rate exists in lower rates of unemployment—that is, there is a decline in structural unemployment. For additional indicators of improved efficiency of the matching process, see Box 5.1 in the Bank of Israel 2011 Annual Report.

Figure 2
Unemployment rate and job vacancy rate (the Beveridge curve), 1998–2013



Another way to assess the usefulness of the estimated NAIRU is to check its ability to identify a Beveridge curve in the data for Israel—that is, to identify the negative correlation between the unemployment rate and the vacancy rate. To do so, we shall first describe in short the theoretical background to the Beveridge curve, and the role of the NAIRU in that context. Following Blanchard (2009), the relationship (in equilibrium) between the flow of workers out of employment and the flow into employment can be described as follows:

$$(9) \qquad s(1-u) = mF(u,v)\,.$$

The left hand side of equation (9) represents the flow of separations from employment (s is the separation rate), and the right hand side represents the flow of new hires, which is assumed to be captured by a matching function $F(u,v)$ of unemployed workers with vacancies. This function increases with u and v, and m is a scale variable that represents the efficiency of the matching process. Equation (9) is the Beveridge curve. Along the curve (holding s and m constant) it represents an inverse relationship between unemployment and vacancies. In recessions, unemployment is high and vacancies are low, and in boom periods, unemployment is low and vacancies are high. The factors that shift the curve are changes in the separation rate and changes in the efficiency of the matching process.

Based on Dickens (2009), we assume that the matching function has a Cobb-Douglas form:

(10) $$F(u,v) = ku^{1-b}v^{b}$$

Where $0 \le b \le 1$ represents the elasticity of the matching function with respect to the vacancy rate. $(1-b)$ is the elasticity with respect to the unemployment rate. Substituting (10) into (9) and arranging terms we get the following equation:

(11) $$(1-u)/u = (km/s)(v/u)^{b}$$

Following Blanchard (2009) and Dickens (2009), we assume that the NAIRU is a function of the separation rate (s) and the efficiency of the matching process (m) [25], and that the NAIRU is the main driver behind the movements in the curve. Thus, the NAIRU is specified as follows:

(12) $$NAIRU = k'_0 + k'_1 \log(km/s)$$

where k'_0 is a constant and k'_1 is negative. Taking logs on the two sides of (11) and substituting (12) into (11), we get the following relation between unemployment, the NAIRU and job vacancies:

(11a) $$\ln[(1-u)/u] = k_0 + k_1 NAIRU + b\ln(v/u),$$

where $k_0 = -k'_0/k'_1$ and $k_1 = 1/k'_1$.

Using the estimated NAIRU from the Phillips curve (estimated in Model 1) and data on the unemployment rate and the vacancy rate, we can estimate the parameters of equation (11a). Ordinary Least Squares estimates of the parameters of equation (11a)[26] are presented in Table 2:

[25] Blanchard claims that the NAIRU is also a function of the bargaining power of workers.

[26] A similar equation was estimated by Dickens (2009). He considered a potential endogeneity of $\ln(v_t/u_t)$ and used instrumental variables (IV) to estimate the equation. He found, however, that the results from IV estimation are very similar to those from the OLS estimation, indicating that the endogeneity bias is negligible. We assert that estimation of the Beveridge curve by OLS should provide consistent estimates even if $\ln(v_t/u_t)$ is endogenous. This is because we actually estimate a cointegration relationship, where all the right hand side variables are integrated of order one. Thus, all parameters should be consistent, irrespective of whether $\ln(v_t/u_t)$ is stationary or not. (A Dickey-Fuller test reveals that the null hypothesis that $\ln[(1-u)/u]$ and $\ln(v_t/u_t)$ are random walk is not rejected for any reasonable level of significance (p-values of 0.77 and 0.35, respectively)). For the relevant test for the NAIRU see footnote 17.

Table 2
Estimates of the Beveridge curve parameters (equation 11a) for the period 1998:Q1–2013:Q4*

k_0 3.48 (86.9)	k_1 -9.79 (20.7)	b 0.22 (8.1)
$R^2 = 0.943$ $DW = 0.728$		

* The t-values of the estimates are in parentheses.

We note that in cointegration equations such as (11a), the distribution of these statistics is not a t-distribution. According to an Engle-Granger test, which examines the cointegration between the variables of an equation by examining the stationarity of the deviations, the hypothesis that the deviations are characterized by a random walk process is rejected at every significance level greater than 0.033.

As can be seen, all the estimates have the correct sign and the fit of the equation is high. From this, and from the existence of a cointegration relationship among the equation variables, it can be concluded that the NAIRU estimated from the Phillips curve helps to identify the negative relationship between the unemployment rate and the vacancy rate. When removing the NAIRU from the equation (an assumption that is the same as assuming the NAIRU is constant), we see that the deviations of the equation are not stationary, which indicates the crucial contribution of the estimated NAIRU to the identification of the Beveridge curve.[27]

4. CONCLUSION

We estimated a time varying NAIRU for Israel's economy using quarterly data for the period 1992:Q1–2013:Q4. As a proxy for expected inflation, we used estimates for 1-year inflation expectations derived from data from Israel's bond market. As a sensitivity analysis, we also estimated the NAIRU using the typical approach of using inflation lags as a proxy for inflation expectations. In both cases, the results are similar, though we found support for added value in using expectations derived from the capital market. The estimations of the inflation equation indicate that the unemployment gap impacts on inflation with a lag of one year. It was also found that the changes in the unemployment gap, with a one-quarter lag, impact on inflation (the speed limit effect), and that changes in the real price of imports have a marked and prolonged effect on inflation (an expression of the openness of the economy).

[27] According to the Engle-Granger test the null hypothesis that the deviations of the equation are non-stationary is rejected only for significance level of 49% (p-value of 0.49). We get much lower values for the R^2 and for the DW as well—0.55 and 0.30, respectively.

An additional finding is that the estimated NAIRU is not constant, and that its development is in line with the low frequency dynamics of the actual unemployment rate. Between 2004 and 2013, the unemployment rate declined by 7.7 percentage points. The NAIRU estimation shows that most of the decline in the unemployment rate, 6.2 percentage points, reflects a decline in the NAIRU. This finding is in line with other assessments regarding a decline in the structural unemployment rate during those years (see Section 2 above). It was also found that the changes in the estimated NAIRU are characterized by a high degree of persistence, which can be interpreted as reflecting a prolonged effect of structural factors. Suhoy and Friedman (2005) assumed from the outset that changes in the NAIRU are white noise. The formulation in this paper includes the above assumption as a specific case that is rejected by the data.

It was also found that the estimated NAIRU makes an important contribution to identifying the negative correlation between the unemployment rate and the job vacancy rate, described by the Beveridge curve. This result strengthens the assessment that the NAIRU estimated from the inflation equation is a good estimator for the structural unemployment rate.

Appendix 1
The actual unemployment rate and the NAIRU estimated using several alternatives

Figure A1
The Actual Unemployment Rate and NAIRU Estimated Using the 3 Alternatives in Table 1

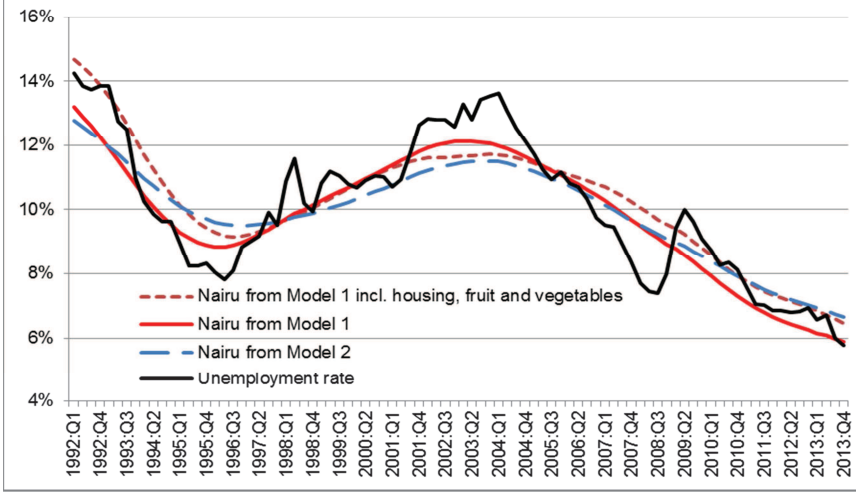
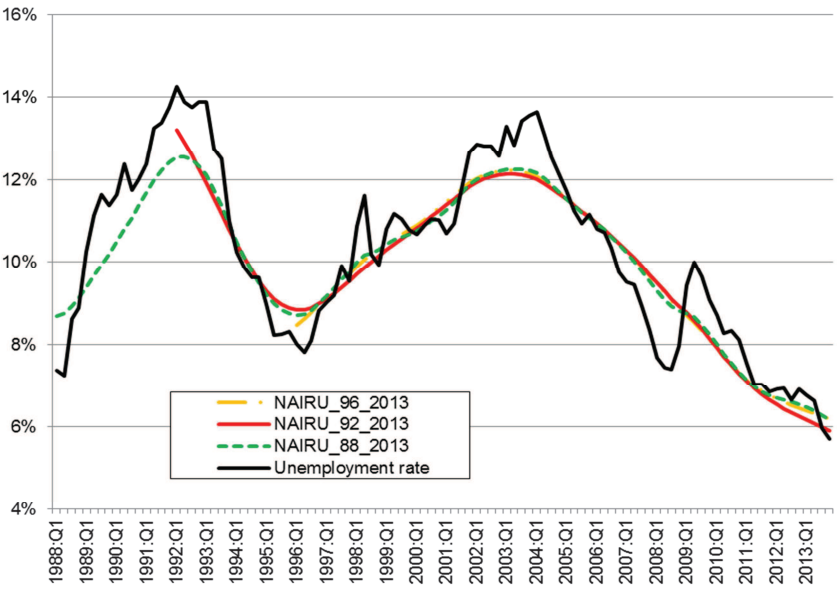


Figure A2
The Actual Unemployment Rate 1988:Q1–2013:Q4, and the NAIRU Estimated Using 3 Alternatives: Beginning in 1988, Beginning in 1992 and Beginning in 1996



Appendix 2: Estimating an extended version of Equation 4

To reach the final specification of the inflation equation, as presented in equation 4a, we began the estimation with a broad formulation (subject to a limitation of degrees of freedom), which includes the following variables: change in shekel price of imports in the present and with a lag of up to 4 quarters, the acceleration in world (dollar) prices of imports in the present, four lags of the inflation rate and five lags of the unemployment gap. The broader formulation of equation 4 follows:

$$\pi_t - \hat{\pi}(market)_{t,t+4} = a_0\Delta epim_t + a_1\Delta epim_{t-1} + a_2\Delta epim_{t-2} + a_3\Delta epim_{t-3}$$
$$+a_4\Delta epim_{t-4} + b_0\Delta\Delta pim_t + c_1\pi_{t-1} + c_2\pi_{t-2} + c_3\pi_{t-3} + c_4\pi_{t-4} +$$
$$d_1 ugap_{t-1} + d_2 ugap_{t-2} + d_3 ugap_{t-3} + d_4 ugap_{t-4} + d_5 ugap_{t-5} + e_t^{\pi}$$

The results of the estimation of the inflation equation (estimated simultaneously with the other equations of the model) are presented below:

	Estimate	s.d.	t-stat
a0	0.0752	0.0259	2.8999
a1	0.0291	0.0199	1.4677
a2	0.0231	0.0208	1.1094
a3	-0.0148	0.0188	0.7852
a4	0.0307	0.0193	1.5897
b0	0.0494	0.0187	2.6455
c1	0.2088	0.0735	2.8397
c2	-0.2962	0.0800	3.7038
c3	-0.0785	0.0753	1.0422
c4	-0.0335	0.0643	0.5207
d1	-0.2702	0.0932	2.8986
d2	0.1648	0.1235	1.3344
d3	0.1969	0.1123	1.7541
d4	-0.2262	0.1126	2.0093
d5	-0.0475	0.0995	0.4779

As explained previously, we gradually added constraints to this specification, which were derived from the sign and significance of the estimators and from the requirement that there will not be a long-term tradeoff between inflation and unemployment, or between inflation and changes in the real price of imports. After several iterations, we reached the specification in 4a.

Appendix 3: Structural changes in the labor market in 2002–07 and their possible effect on the NAIRU

In 2002 and 2003, a government plan was formulated and adopted, continuing in 2004–07 as well, which led to a prolonged structural change in the labor market.[28] Among the goals of the plan was making the labor market more flexible, increasing the employment rate and reducing the share of people relying on transfer payments. The steps taken included reducing the number of people employed in public services and their wages, taking actions to reduce the number of foreign workers, reducing the negative incentives to work that are inherent in various transfer payments, reducing the cost of labor and reducing tax rates on wage income.

Regarding transfer payments, it was decided to cut the income support allowance and related benefits, to reduce payments to people receiving child allowances, to increase the stringency of the terms of eligibility for unemployment benefits and to reduce the scope of the payment.

One of the expressions of increased strictness of the criteria for receiving unemployment benefits was a marked reduction in the share of unemployment-benefit claims that were approved. The share of claims approved relative to total job seekers in 1995–2002 was 0.65, on average. In 2003, the share declined to 0.42 and in 2004 there was a further decline, to 0.32.[29] It is reasonable to assume that this also contributed to a decline in the structural unemployment rate. Another way to see the negative impact on the unemployed is to examine the number of unemployment benefit recipients vis-à-vis the number of unemployed people—in 2002–04, the number of benefit recipients declined by a cumulative 53 percent, while the number of unemployed people increased by a cumulative 18 percent.

Another expression of the plan is the marked reduction of tax rates on wages. Between 2002 and 2008, tax rates on wages declined from 32 percent to 26 percent, and reached 24 percent in 2011. This reduction almost certainly acted to increase the participation and employment rates, and possibly to reduce the structural unemployment rate.

An additional factor that may be attributed to the results of the plan is the flexibility seen in Israel's labor market as it responded to the global crisis of 2008. As described in Chapter 5 of the Bank of Israel Annual Report for 2009, the recovery of demand in the market was accompanied by a renewed trend of decline in the unemployment rate, in contrast to many advanced economies in which the recovery of activity was not accompanied by an increase in employment. A possible explanation for this is a decline in the friction in the labor market (that is, improved matching between employment seekers and suppliers), expressed in a decline in the structural unemployment rate. Corroboration

[28] For an extended discussion on the steps and their outcomes, see the chapter on the labor market in the Bank of Israel's annual reports for the relevant years as well as the annual surveys of the National Insurance Institute for the relevant years.

[29] In the years 2005-08, the rate remained 0.32 and then increased somewhat to 0.36 in 2009-13.

for this can be found in the leftward shift of the Beveridge curve in 2006 and 2007. Additional evidence of the structural change that occurred in Israel's labor market is the extended decline of the rate of people moving from employment to unemployment, which occurred in 2002–07.[30]

[30] For a detailed discussion on these issues, see Box 5.1 in the Bank of Israel Annual Report for 2011.

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