Modeling the Dynamics of Welfare Caseload: a New Approach

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Abstract

What are the determinants of welfare caseload fluctuations? To address this crucial policy issue most current research uses either cross-sectional or time-series models. In this study we propose a new two-step latent factor approach. First, a latent-factor model is fitted to the data. Next, time-series techniques are applied to study the relation between the latent factor and control macroeconomic and demographic variables. As an application, we study the dynamics of welfare caseload in Israel over the period 1986-2002. Our major findings are as follows. First, a single latent-factor model accounts for more than 80% of the variance in welfare caseload. Second we find strong evidence that the latent factor and control variables are cointegrated. Third, the cointegration relation has experienced a structural shift during 1994-1995 following new welfare legislations as well as changes in the Israeli labor market. Overall, our findings suggest that both economic and demographic conditions as well as welfare policy regulations play an important role in explaining welfare caseload trends.

JEL code: J21

Keywords: welfare caseload, latent factor, macroeconomic conditions

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1 Introduction

Understanding the dynamics and determinants of trends in welfare caseload has become the issue of harsh debates during recent decades (see Blank, 2000; Bell, 2001; Ayala and Perez, 2005 among others). This is of a crucial importance for policy makers in designing as well as assessing the effectiveness of welfare programs, especially in light of the dramatic changes in welfare caseload. In the United States the total caseload experienced a decline of about 40 percent between the beginning of the 1990 to mid-1997 when the Responsibility and Work Opportunity Reconciliation Act (PRWORA) was implemented nationwide (Wallace, 2007). More recent evidence for fluctuation in welfare caseload comes from Israel where a total number of Income Support recipients has risen dramatically between 1998-2002.

In this study we propose and implement a novel two-step latent factor approach to model the dynamics of welfare caseload. First, the latent factor model is estimated using the Lee-Carter (1992) approach applied to a panel of welfare benefits recipients. Next, we study the links between the latent factor and the observable control variables by means of cointegration and variance decomposition techniques. Using this approach we study the factors affecting the dynamics of Income Support recipients in Israel over the period of 1986-2002.

The contribution of this study is twofold. First, in contrast to existing studies we use the latent factor analysis. Latent factor models are based on the assumption that the comovement between economic series is driven by a finite number of common factors. These models have been extensively used in applied economic research such as modeling the dynamics of business cycles (Stock and Watson, 2002), labor market research (Heckman, Stixrud and Urzua, 2006), forecasting mortality trends (Brouhns, Denuit and Vermunt, 2002) and many more. However, to the best of our knowledge no study implements the latent factor approach in modeling the changes in welfare caseload of different income support groups. This is quite surprising since economic theory predicts that the level of welfare caseload in different income support groups is likely to be affected by common factors such
as financial incentives to work and changes in macroeconomic and demographic conditions. Our approach is also fairly general in the sense that by using the latent factor model we do not make any ex ante assumptions regarding the variables driving the dynamics of welfare caseload. This advantage is particularly significant in light of the unsatisfactory explanatory power of the control variables reported in the previous studies\(^1\). Also, as we shall demonstrate in the following sections, our approach fits our data well and is easy to implement with most existing statistical packages. This makes it a useful tool for further empirical studies.

Second, while the majority of studies examine the trends in the US welfare caseload we focus on the non-US (Israeli) case. During the time-period covered in this study, the Israeli economy experienced significant economic and demographic perturbations. Therefore, focusing on the Israeli case may provide us important clues about the links between the dynamics of welfare caseload and macroeconomic and demographic factors, as well as changes in welfare policy which affect financial incentives to work. Also, in contrast to the previous studies on Israel which focus mainly on one of the abovementioned potential determinants, our approach allows us to control for all of these variables and also to quantify the marginal contribution of each of them to the dynamics of welfare caseload.

Our key findings are as follows. First, we find that a single latent-factor model explains, on average, more than 80% of the dynamics of welfare caseload. Second, we find strong evidence of a cointegration relationship between the estimated latent factor and the macroeconomic, demographic and financial incentive variables. The signs of the cointegration equation parameters are in accordance with economic theory. Finally, we find strong evidence of a structural shift in the relationship between the latent factor and the control variables around the years 1994-1995, a period when the Israeli labor market experienced significant perturbations due changes in welfare policy as well as the massive entrance of foreign labor. Controlling for this shift, we find that changes in the macroeconomic and

\(^1\) As Bell (2001) points out in the context of 1996 US reform "This literature reveals that it is very difficult to untangle the causes of cash assistance caseload declines.....Even when many causal factors are considered at once, none of the several studies reviewed here comes close to explaining all of the unprecedented caseload decline in the 1990s."
demographic conditions and changes in financial incentives to work explain about 50\% of the latent factor dynamics. Curiously, while financial incentives to work seem to affect the "long-run" dynamics of the welfare caseload, its cyclical behavior appears to be mainly driven by changes in the macroeconomic conditions, or the business cycles.

Our findings are of particular interest for policy makers in light of the ongoing Welfare-to-Work program which pursues integration of welfare recipients into the labor market. Benjamin Netanyahu, the Israeli finance minister in 2003-2005, claimed that the income maintenance allowance distorted the motivation to participate in the labor market and created a "poverty trap". Our findings suggest that while the level of welfare caseload is significantly affected by financial incentives to work, other factors, such as conditions on labor market and business cycles should be taken into account by the policy makers as well.

The remainder of this paper is organized as follows. In Section 2 we review some of the existing literature on the subject and discuss the motivation behind the choice of the control variables. In Section 3 we present the model and discuss the methodology. The data is described in Section 4. Empirical findings are reported and analyzed in Sections 5, 6 and 7. Finally, in Section 8 we offer concluding remarks and discuss some directions for further research.

2 Literature Review

As pointed out by Blank (2002), "...... prior to the mid-1990s, there was virtually no published literature in economic journals looking at the movements in caseload over time" (Blank, 2002. pp.1127). However, since the American Welfare Reform legislation was passed in 1996, the literature on caseload changes has rapidly grown. This literature takes multiple economic variables and control variables for policy components and (in the case of US studies) state waiver dummies as explanatory variables (for an extensive review of the findings see Blank, 2002; Grogger and Karoly, 2005). Some studies use either time-series or panel data techniques (Wallace & Blank, 1999; Blank, 2000; Ayala and Perez, 2005) while
other use pure cross-sectional analysis (Grady, 1999; Mead, 2000, Mead, 2003). As noted by Mead (2003), each approach has its own strengths and weaknesses. The time-series or panel data approach which is mostly applied to aggregate data at the national level takes into account the time dimension but at the cost of limiting the number of explanatory variables. On the other hand, a cross-sectional approach which usually treats states (in the case of the US) as distinct units allows inclusion of multiple welfare policy variables but, in turn, gives up the time dimension (Mead, 2003; pp.164-165).

Based on the current literature we consider three factors to be the potential determinants of welfare caseload dynamics:

- Financial incentives to work,
- Macroeconomic factors,
- Demographic changes.

2.1 Financial Incentive to Work

A common theoretical model used in the literature for evaluating the labor supply and other work incentives effects of means-tested programs is the basic static leisure-consumption model (Moffitt, 2002). According to the model, the consumer maximizes his utility, while the decision of whether to work or to participate in a welfare program depends on two observable factors: the purchasing power of the allowance and the alternative wage in the labor market.

In order to maintain the standard of living of the recipients, the income support allowance is indexed to the average wage in the economy. Besides the allowance, recipients are entitled to in-kind services from the public sector. Gottlieb (2001) demonstrates that the sum of the allowance and the in-kind services can reach up to 96% of the disposable income of a worker with 11-12 years of schooling. Some researchers argue that the generosity of the Israeli welfare system is one of the main reasons for the increase in the number of income support recipients (Gottlieb, 2001; Romanov and Sussman, 2004). The second
reason is that the wages of unskilled workers in Israel have deteriorated. Since the early 1980s inequality in the wages of skilled and unskilled workers has grown owning to technological improvements, relative decrease in the demand for unskilled labor, globalization of the economy, growing share of foreign workers, and the influx of immigrants during the 1990s (Kassir et al. 2000; Dahan, 2002; Sussman, 2004). Similar trends have occurred in the American labor market where the wage rates of unskilled workers have fallen dramatically relatively to skilled workers (Blank et al, 2006). As a consequence of the increase in the welfare benefits and the decrease in the wage for unskilled labor, the relative attractiveness of social benefits has increased, which is positively affecting the decision to participate in the income support program.

Early studies also find that the Aid to Families with Dependent Children (AFDC) program reduced labor supply of single parents by from 10 to 50 percent of non-AFDC levels (Danziger et al. 1981; Moffitt, 1992; Hoynes, 1997). The reduction in labor supply of single mothers due to the negative effects of income support was confirmed in other countries such as the UK (Gregg and Harkness, 2003), Australia (Doiron, 2004) and Israel (Flug and Kasir (Kaliner), 2006; Frish and Zussman, 2008).

2.2 Macroeconomic Factors

The state of the economy is a crucial factor for the chances of being employed and thus avoiding falling into the safety net of the welfare system. Researchers usually include the unemployment rate as a control variable for the state of the economy since reduced economic demand during recessions leads to an increase unemployment rate, and during an economic expansion situation is reversed. Previous studies suggest that in the case of the US the unemployment rate is an important factor which should be taken into account when modeling the dynamics of welfare caseload (Blank, 2002; Grogger and Karoly, 2005). Also, the unemployment rate is reported to be a particularly important labor market indicator for less-skilled workers, since the employment of less-skilled individuals experiences greater fluctuations compared to high-skilled individuals (Hoynes, 1999; Blank et al. 2006).
Controlling for the impact of macroeconomic factors is particularly important in the context of our study. During the years 1986-2002, the Israeli economy experienced three periods of growth, with the longest one lasting 69 consecutive months, and three periods of recession, with the longest one covering the period between October 1996 and August 1999 (Flug and Strawczynski, 2007). Similarly to the US labor market, the employment rate of less skilled workers in Israel tends to fluctuate more with business cycles compared to that of skilled workers (Flug et al. 2000; Arnon and Pressman, 2007). Dahan (2007) argues that high unemployment rates prompts individuals to despair of finding work and to disengage from the labor market, thus, leading to an increase in welfare caseload. On the other hand, some researchers do not take into account the effect of business cycles (Romanov and Zussman, 2004), but, other, argue that the effect of the unemployment rate is not persuasive (Gottlieb, 2001).

2.3 Demographic Changes

Since 1989, Israel has experienced a dramatic demographic change, caused by immigration of massive proportions, mainly from the former Soviet Union. More than a million immigrants have arrived, increasing the population by over 20% through the 1990s. As noted by Freidberg, "no immigration to United States or Western Europe has been comparable in magnitude" (Freidberg, 1996).

Labor economics theory suggests that massive immigration leads to an increase in the number of welfare recipients. First explanation is the "magnetic" effects of a generous welfare system on the decision "where to go" as suggested by Borjas (1999). The Israeli experience demonstrates that the inflow of new claimers for the income support program grew and by the year 2002 about one third of all the recipients were immigrant families (National Insurance Institute, Annual Survey, 2003). Interestingly, while concerns that immigrants might become "public charges" led to a reduction in welfare-related rights of immigrants under the 1996 welfare reform in the US (Borjas, 2002), it has been claimed in Israel that the income support program has served as an effective tool to acclimate
immigrants with neither savings nor pension rights (Swirski, 2004).

A second theoretical explanation concerns the effects of immigration on the native recipients. The textbook model of a competitive labor market predicts that immigrant influx should lower the equilibrium wage and increase the equilibrium employment. Assuming imperfectly elastic labor supply and demand curves, then immigrants should displace some natives in employment and therefore some natives will join the welfare rolls. Borjas (2003) presents evidence suggesting that immigration to US from the 1960-2001 has harmed the employment opportunities of competing native workers. Freidberg (1996) examines a microdata of immigrants from the former Soviet Union to Israel over the period 1989-1994 and shows a negative relation between native employment growth and immigrant entry, which would imply more native recipients on welfare.

3 Methodology

In this section we describe the model and methodology used in this study. The model and the estimation procedures are described in subsection 4.1. In subsection 4.2 we discuss some estimation accuracy and statistical inference issues. Finally, in subsection 4.3 we describe methodology used in this study.

3.1 A Model

Following the approach, originally proposed by Lee and Carter (1992) in their influential paper we assume that for each period of time $t$ and for each income support group $i$, the total (log-transformed) number of households in each income support group, $h_{i,t}$, is described as the following process

$$h_{i,t} = \alpha_i + \beta_i f_t + \epsilon_{i,t}$$  \hfill (1)

$^2$The description of each income support group is provided in the following section.
Here, $\alpha_i$ and $\beta_i$ are group-specific parameters, $f_t$ is the latent factor which governs the joint dynamics of all the groups and $\epsilon_{i,t}$ is the idiosyncratic shock which is assumed to be uncorrelated with $f_t$. An important feature of this model is that while $f$ is inherently treated as a stochastic factor, it is estimated as an additional "parameter" along with the group-specific vectors of parameters, $\alpha$’s and $\beta$’s. The goal of this method is to decompose the variation in welfare caseload in each group into a "systematic" component, that is, the part which is driven by economic factors captured by the latent factor $f$ and the "individual" or idiosyncratic component. Lee and Carter (1992) propose to estimate the model by the least-squares method

$$[\hat{\alpha}, \hat{\beta}, \hat{f}] = \text{arg min} \sum_{t=1}^{T} \sum_{i=1}^{N} (h_{i,t} - \alpha_i - \beta_i f_t)^2$$  \hspace{1cm} (2)

Lee and Carter show that under current specification the model is unidentified. Therefore, following Lee and Carter’s suggestion we impose the following restrictions

$$\sum_{i=1}^{N} \beta_i = 1$$

$$\sum_{t=1}^{T} f_t = 0$$

Under these identification restrictions $\hat{\alpha}$’s become sample means. On the other hand, $\hat{\beta}$’s will measure factor exposures of each income support group, that is, how sensitive the number of households in each income support group is with respect to changes in the latent factor.

Clearly, this model cannot be estimated via standard statistical packages due to bilinear term $\beta_i f_t$. However, Lee and Carter show that the estimates can be obtained by applying a Singular Value Decomposition (SVD) to the demeaned data matrix of the households. They show that two vectors corresponding to a largest singular value, after imposing an identification constraint on $\beta$’s, provide a unique solution to the least-squares problem.
3.2 Estimation Accuracy and Statistical Inference

While Lee and Carter (1992) propose a straightforward and elegant way of modelling a joint dynamics of panel data, one must be careful in assessing the accuracy of the estimated parameters as well as in testing statistical hypotheses. In particular, since the number of parameters estimates grows as the sample becomes larger, applying standard tools, such as Mean-Value and Central Limit Theorems, to derive asymptotic distribution of the estimators is, to say the least, questionable. Therefore, instead of relying on the asymptotic theory we base our inference on a finite sample distribution by conducting a "wild bootstrap" analysis. The "wild bootstrap" approach is widely used in the literature to handle cases when the asymptotic distribution of the estimators is unknown or non-trivial. With some modifications, this procedure can be also applied to testing various restrictions imposed on the parameters of the model. For instance, an issue of particular interest is whether households from different income maintenance groups have the same exposure to the joint factor. This is particularly important to the policy makers, who would like to forecast which of the groups will be most affected by the introduction of new welfare regulations.

4 Data Description and Control Variables

4.1 Data Description

We use data on the number of households receiving income support allowance, according to four major entitlement groups. The first group includes unemployed (the head of the household) who have to take a work test in the Employment Service, hereafter referred as "UNEMP". The second group, "LOWAGE", includes low-wage workers, who are entitled to an income supplement. The third group, "SINP", includes single mothers who are exempted

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3See Diebold et al., 1996, Ivaldi, 1996 and Miles and Mora, 2003 among others. The details of the "wild bootstrap" procedure used in this study as well as the related Matlab codes are available from the corresponding author.
from a work test as long as the age of their youngest child is less than 7. Finally, the fourth group, "EXEMP", includes individuals who are temporarily exempted from the work test, due to hardships in filling vacancies in the labor market. The source for our data is monthly reports by the National Insurance Institute\textsuperscript{4}. Since the majority of these reports have never been published the data has been hand-collected from the National Insurance Institute archive. The time span of our sample is from November 1986 to March 2002.

In Figure 1 we present some of the observable dynamics of the income support groups. It seems that all groups follow an upward time trend, with the "UNEMP" group tending to fluctuate around its time trend more than the others. This group grew from the late 1980s to the end of 1992. It reached a bottom in the mid 1990s, but amazingly, has steeply grown since. Figure 1 shows a steady grow in the "EXEMP" group until 1998, and then what seems to be a structural break. This dramatic change can be potentially attributed to a new employment service policy which granted recipients a permanent waiver from taking a work test (for further details on this reform see Gottlieb, 2001).

The lower plot in Figure 1 shows the marginal contribution of each group to the total caseload. The most astonishing figure is the growing share of the "EXEMP" group - from a reasonable contribution of 18\% in the late 1980s, to over 40\% in the late 1990s. The increase in the marginal contribution of the "EXEMP" group is in a sharp contrast to the declining share of the "LOWAGE" group - from 34\% at the beginning to 13\% at the end of the time span. The marginal contribution of the "UNEMP" group fluctuates from 7\% in June 1988 up to 39\% in January 1993, then down to 16\% in February 1996 and up again to 41\% in March 2002. The share of the single parents groups, SINP, has also declined from 29\% at the beginning of the year 1987 to about 16\% by the beginning of the year 2002.

\textsuperscript{4}For further details, the reader is referred to the Israeli National Insurance Institute official website www.btl.gov.il.
4.2 Control Variables

- Financial Incentives

A common explanation for the increased caseload is the relative increase in the attractiveness of the income support allowance which relate to financial incentives to work, as discussed in a previous section. To control for the impact of financial incentives on work we calculate the Replacement Rate (RR), meaning the ratio of the entitlement allowance to the alternative wage in the labor market. This variable is constructed as follows:

\[ RR_t = \frac{\delta_t \bar{W}_t}{W_t} \]

Here, \( \bar{W}_t \) is the average monthly wage in the economy published by the Central Bureau of Statistics (CBS). \( \bar{W}^a_t \) is the alternative wage for unskilled labor. Following Gottlieb (2001) we measure \( \bar{W}^a_t \) as the gross wage of 35-44 year old individuals, with 0-10 years of schooling. The source for the data is Income Surveys, which are published annually by the CBS. Finally, \( \delta_t \) is the annual indexation rate to \( \bar{W}_t \) determined by the policy makers. During the time span of our study it ranged between 0.32 to 0.38. Since both \( \bar{W}^a \) and \( \delta \) are available on annual basis only, we assume these variables to be constant within a particular year. Inspection of Figure 2 suggests that the RR was relatively stable until the middle of the year 1994 and then started to trend upwards. The main reason for the increase in the financial incentives to work variable is a substantial decline in the wage for unskilled labor due to the intensive entry of foreign workers whose number more than doubled during the years 1995-2000 (Gottlieb, 2002).

Figure 2 approximately here

- Business cycles indicators
We use two measures for business cycles: the unemployment rate $UN$, and the State-of-The-Economy Composite Index, $CI$. Data for the unemployment rate available by the CBS, only in a quarterly frequency, therefore the UN is fixed in each quarter. The CIT is published by the Bank of Israel and is available on monthly frequency. The components of the Index are calculated from the following indicators: the index of manufacturing production (26%); the index of imports (12%); the index of revenue from trade and services (28%); the index of business-sector employment (23%); and the index of export goods (10%) (For a detailed description see Marom et al, 2003).

- **Demographic factors**

To control for the demographic changes we include the total population of the working age population, $POP$, and the monthly inflow of the new immigrants, $IM$. The source is Israel Statistical Journal published on monthly basis by Israeli Central Bureau of Statistics.

## 5 Empirical Findings-Lee-Carter Model

The estimates of the Lee-Carter model applied to household data are presented in Table I, on the left hand side of Panel A. Following Lee and Carter (1992), the data was transformed to natural logarithms and sample means were subtracted from each series. Corresponding bootstrapped 95% confidence intervals are presented beneath each estimate. Since the estimates of $\alpha$’s are the sample means by construction, we turn to the analysis of $\beta$’s, the estimates of factor exposures.

For all household groups the estimates are positive and highly statistically significant. Tight bootstrapped confidence intervals also indicate that the estimates are fairly accurate. Our findings suggest that the UNEMP group is characterized by the highest exposure to the latent factor. On the other hand, a LOWAGE group, that is, the households working at a part-time job and having an income supplement from the state, appeared to have the lowest sensitivity to the factor among the four groups included in our study. The results of a likelihood ratio test suggest that the null hypothesis of equal factor exposures is rejected.
at any reasonable significance level with the corresponding p-value being close to zero. The explanatory power of the Lee-Carter model is quite impressive. A single-factor model explains about 95 percent of total variation for the UNEMP group and about 90 percent for the EXEMP group. The estimates of $R^2$ for the LOWAGE and SINP equations are somewhat more moderate but still substantial (0.62 and 0.86, respectively).

-Table I approximately here-

The estimated factor along with the 95% confidence bands is depicted in Figure 3. Interestingly, the factor exhibited a downward trend until the end of 1988 with an abrupt reversal following afterwards. A downward trend is potentially attributed to the post-recession recovery period of the Israeli economy following a major crisis during the beginning of the 80s. During this period we also observe a consistent improvement in the composite index measure, CI. This pattern can also be attributed to statutory changes, in particular, the minimum wage legislation in 1987 which increased the minimum wage from 33.6% of the average wage in the year 1986 to 41.1% by the end of 1988. This legislation mainly affected the LOWAGE group, who were no longer eligible for the income supplement.

Starting from the year 1989, Israel opened its gates to a massive immigration from former Soviet Union, which reached it peak around 1991-1992 when about 13,000 new immigrants arrived each month (which exceeds by more than twice a long-run monthly average of immigrant arrivals). During the same time period we observe that the factor is following an upward trend until the end of 1992 when the inflow of immigrants substantially subsided. Between 1993-1996 the factor is fairly flat and stable. However, starting from the year 1997 it is upward trending again, a tendency which continues until 2002. Interestingly, this tendency coincides with a rapid increase in the number of foreign workers employed in the Israeli economy, a number which has doubled within six years from 120,000 thousand at the beginning of 1996 to 240,000 thousand in the first quarter of 2002 (Miaari and Sauer, 2006). The massive inflow of foreign labor force has increased the replacement rate and, thus, reduced the incentives to leave the safety net of the income support programs.
For comparison, we estimate the same model with each monthly household observation scaled by the working age population, POP, for that particular month. Here, we model the dynamics of the income maintenance groups measured as shares of a total population instead of levels. The results are presented in the right hand side of Table I, Panel A. Interestingly, for all but one group (UNEMP) we can observe a decline in the explanatory power of the model. This decline is particularly pronounced for the LOWAGE and SINP groups, with R-squared of SINP dropping almost twice from 0.86 to 0.46 while for LOWAGE groups it plummets to almost zero. Also, the estimates of factor exposures, $\beta$'s experience a substantial decline and for LOWAGE group it becomes statistically insignificant. These results provide a preliminary support for our conjecture that demographic changes can partially explain the dynamics of income maintenance household groups. The impact of demographic changes is particularly pronounced for the LOWAGE group, whose dynamics seems to be fully explained by population growth which also may explain low estimate of the LOWAGE $\beta$.

CBS provides data on the number of households in each income support group but does not provide details regarding potential transfers between these groups. In particular, it is possible that changes in welfare caseload in the UNEMP, EXEMP and LOWWAGE groups are partially driven by the households shifting from one group to another. To take this potential effect into account we estimate a two-factor model, with the second factor corresponding to a second largest eigen value. Since factors are orthogonal by construction, adding second factor to our model does not affect the estimates of a single-factor model reported in Table I. The estimate of the UNEMP second factor beta is set to be equal to one. If the second factor captures transfers between the UNEMP, EXEMP and LOWWAGE groups, this would imply that the sum of the second-factor betas of these groups should be equal to zero. Also, since no transfers are expected to occur from these groups to SINP group, the second-factor beta of SINP should be equal to zero as well. However, the estimate of SINP second-factor beta is negative and statistically significant. Also, the analysis of the
corresponding eigen values suggests that a two factor model explains about 86% of the variation in welfare caseload. This finding suggests that even if the transfer of households from one income support group to another does occur and is captured by one or two of the remaining factors, its impact on the dynamics of welfare caseload is not likely to be economically significant.\textsuperscript{5}

6 Empirical Findings—the latent Factor and the Observable Factors

In this section we examine the relation between the estimated factor from the Lee-Carter model and the control variables, discussed in Section 4. In this study we make a distinction between the long-run and the short-run relation between the latent factor and the control variables. The existence of the long-run links between the latent and the observable factors is examined by means of cointegration analysis (Engle and Granger, 1987). The dynamics of the short-run relation between the latent and observable factors is studied within variance decomposition framework (Hamilton, 1994, Lutkepohl, 2006). First, we test for the presence of and examine the long-run relation by using a cointegration analysis. Next, we assess the contribution of each observable variable to the dynamics of the estimated latent factor by means of variance decomposition technique.

6.1 A Cointegration Analysis

We start with the time-series properties of our data. The results of the Augmented Dickey-Fuller test suggest that both the estimated factor and the observable variables are non-stationary in levels but stationary in their first differences\textsuperscript{6}. Hence, as a next step we proceed with testing for the presence of one (or more) common stochastic trends. We apply a test proposed by Johansen (1991) corrected for the small-sample size distortion, following

\textsuperscript{5}The estimates of the two-factor model are available from the corresponding author.

\textsuperscript{6}The results are available from the authors upon request.

The results are reported in Table II. The trace and max-eigen value statistics corrected for a small-sample bias along with the corresponding critical values are presented in Panels A and B, respectively. Overall, we find strong evidence that the series are cointegrated. More specifically, both trace and max-eigen value tests reject the null hypothesis of no cointegration in favor of at most one cointegration equation. However, our findings become more ambiguous when we test for the presence of more than one long-run relation. One should note, however, that while corrected for a small-sample bias the tests may still be oversized due to the non-normality of VAR innovations (Cheung and Lai, 1993). Indeed, we find evidence of both statistically significant skewness and excess kurtosis.

-Table II approximately here-

Taking these considerations into account we decide to proceed while assuming the existence of a single cointegration equation. Following Engle and Granger (1987) recommendations, the parameters of the cointegration equations are estimated by using least-squares. The estimated cointegrating equation is presented below with the Newey-West standard errors being reported beneath each estimate, respectively. The sign of all the estimates appears to be consistent with economic theory. There is a positive and statistically significant relation between the factor and the financial incentives to work (RR), a finding which suggests that the number of income support households is expected to increase when labor market prospects become less attractive. This can be due to either an increase in $\delta$, a share of the average salary paid as the income maintenance, or the declining alternative salary.

$$f_t = -371.46 - 0.04 t + 1.48 RR_t - 1.77 CI_t + 5.29 UN_t + 25.27 POP_t + 0.15 IM_t$$ (3)

Turning to the demographic factors, namely the total working age population and the inflow of immigrants, we find that both variables are positively associated with the factor, with the link between working age population, POP, and the factor being statistically
significant as well. These findings suggest that the observed upward trend of the estimated factor can be at least partially attributed to demographic changes in the economy. Finally, our results suggest that the estimated factor is positively linked to the unemployment rate, UN, while being negatively related to the level of state-of-the-economy composite index, CI. Though statistically insignificant, the sign of these estimates is still consistent with the economic intuition that labor market prospects become more (less) attractive during economic expansion (recession).

6.2 A Variance Decomposition Analysis

To quantify the contribution of the economic and demographic variables to the dynamics of the latent factor, we use a variance decomposition technique. Since all variables appear to be I (1), we use first differences instead of levels in estimating VAR. The lag order of the VAR is selected based on Wald test. The contribution of each variable to the variance of latent factor is presented under the corresponding headings, respectively.

The results of the variance decomposition analysis are reported in Table III. Overall, the observed variables seem to explain about 40% of a "long-run" (Period 60) variation of the latent factor. In the "long run" (period 60) financial incentives to work explain about 5.4% of the latent factor variance. Demographic factors (changes in the working age population, $\Delta POP$, and inflow of immigrants,$\Delta IM$) contribute about 12.3% to the variance of the income maintenance households. On the other hand, changes in business cycles indicators (changes in the state-of-economy composite index, $\Delta CI$, and the unemployment rate, $\Delta UN$) appear to explain more than 20% of the dynamics of the latent factor, suggesting that changes in macroeconomic conditions of the economy seem to play a dominant role in the dynamics of the income maintenance households compared to demographic and negative labour incentive factors.

-Table III approximately here

Overall, the results of cointegration and variance decomposition analysis suggest that
the dynamics of the income maintenance households can be, at least partially, attributed
to changes in economic and demographic conditions. However, it seems surprising that the
impact of these variables on the dynamics of welfare caseload, in particular, the link between
the latent factor and the financial incentives to work have limited economic significance.
While the choice of how the variables are measured may contribute to this puzzling result,
it may also be attributed to the structural shifts in the dynamics of the factor. This issue
is addressed in the following section.

7 A Temporal Stability Analysis

In this section we examine whether the relation between the latent factor driving the dy-
namics of the income maintenance households and the control variables is stable over time.
First, we test for the presence of structural shifts in the cointegration relation. Next, we
examine structural stability of VAR system.

Cointegration equation

We start with the analysis of the cointegration equation. In our analysis we adopt
the approach of Gregory and Hansen (1996) who propose a residual-based test to test for
the presence of a structural shift in the cointegrating equation. Following the notations
of Gregory and Hansen (1996), let \( y_t = (y_{1,t}, y_{2,t}) \) be the observed data where \( y_{1,t} \) is real-
valued and \( y_{2,t} \) is an \( m \)-vector. Also, let \( \tau \) be the unknown parameter \( \in (0, 1) \) and let \( \varphi_{t,\tau} \)
take value 0 if \( t \leq n\tau \) and 1 if \( t > n\tau \). We test the null of no
cointegration versus the alternative that \( (y_{1,t}, y_{2,t}) \) are cointegrated with the regime shift in
the parameters of cointegrating equation which occurred at \( t = n\tau \)

\[
H_1 : y_{1,t} = \mu_1 + \mu_2 \varphi_{t,\tau} + \beta_1 y_{2,t} + \beta_2 y_{2,t} \varphi_{t,\tau} + \varepsilon_t
\]

with \( \varepsilon_t \) being \( I(0) \).\(^7\) In context of our study, \( y_{1,t} \) is the latent factor and \( y_{2,t} \) are the
observable control variables. Gregory and Hansen (1996) propose to estimate the model for

\(^7\)This specification corresponds to the \( (C/S) \) model in the original paper of Gregory and Hansen (1996).
each \( \tau \) and to compare the minimum ADF statistic to the corresponding critical values.

The results of the Gregory-Hansen (1996) test are presented in Figure 4. We consider four different specifications of the cointegration equation by varying the number of control variables, \( m \). More specifically, we consider a specification with a single control variable \( (y_2 = RR) \), two \( (y_2 = (RR, IM)) \), three \( (y_2 = (RR, IM, CI)) \) and four \( (y_2 = (RR, IM, CI, POP)) \) control variables. The resulting ADF statistics are plotted against the time line where each point on a time line is considered as a potential date when the structural shift occurred. The minimum ADF statistic is compared to the 10 and 5% critical values denoted by the dotted lines.

We find no evidence of a structural shift for \( m = 1 \) and \( m = 2 \) specifications. However, our conclusions substantially change as we examine the results of the test for the models with three and four control variables. For the \( m = 3 \) specification, after we allow for a structural shift in the loading of the state-of-economy variable, \( CI \), the null hypothesis of no structural shift is strongly rejected with the minimum ADF statistic being significant at 5% level. As we extend the model to \( m = 4 \) specification, which also allows for a structural shift in the population loading, the null is still marginally rejected at 10% significance level. It seems that allowing for a shift in the \( POP \) loading does not seem to have any significant effect on the minimum ADF statistic. On the other hand, it appears that a structural shift in the cointegration equation can be, at least partially, attributed to a structural break in the relation between the latent factor and the state of economy. For both \( m = 3 \) and \( m = 4 \) specifications a breakpoint is detected at March 1994.

Below we report the estimated cointegration equation based on the \( m = 3 \) specification. That is, we allow for a structural shift in the intercept and the loadings of \( RR, IM \) and \( CI \) based on the results of the Gregory-Hansen test. Based on our findings, we choose March 1994 as the month when a structural shift occurred. The corresponding Newey-West standard errors are reported below each estimate, respectively.
Our estimation results support the results of the Gregory-Hansen test. First, we find strong evidence of an upward shift in a level of the latent factor with the estimate of intercept becoming significantly larger. This could be attributed to the presence of the upward trend in the factor over the time span of our study. However, our findings indicate that the loadings of the control variables have undergone a structural shift as well.

In particular, we find strong evidence of the structural shift in the RR and CI loadings. More specifically, prior to March 1994 the estimate of the RR loading is negative and statistically insignificant. In other words, prior to the year 1994 we find no evidence of a long-run relation between the number of the income maintenance households and the financial incentives to work. However, the picture dramatically changes as we turn to the post-structural break period when the estimate of the RR loading becomes positive and highly statistically significant, suggesting that the impact of the financial incentives to work on the dynamics of the income maintenance households has substantially increased. On the other hand, we also see a sharp and statistically significant decline in the estimate of a CI loading which becomes almost eight times smaller and statistically insignificant in the post-March 1994 period. Our results remain qualitatively the same after including the population and unemployment control variables.

Variance decomposition

The matrix of the variance decomposition elements is linked to the structural form VAR via the equations’ parameters, $A$ and the innovations covariance matrix, $\Sigma_u$. Therefore, examining the stability of the variance decomposition matrix is equivalent to testing for the presence of a structural shift in one of the abovementioned elements. Since $\Sigma_u$ enters the variance decomposition formula merely as a orthogonalizing factor, we find it more useful as well as intuitive to focus our attention on the structural stability of the VAR parameters.
Based on the results of Gregory-Hansen (1996) test we choose March 1994 as a potential breakpoint date. Clearly, allowing for structural shift in all parameters would make estimation infeasible due to relatively large dimension of VAR. Therefore, in the following analysis we allow for structural shift in the first-lag loadings only. The following system of equations is estimated

$$ z_t = (A_{1,t<03/94} + A_{1,t=03/94})z_{t-1} + \sum_{i=2}^{p} A_i z_{t-i} + \epsilon_t \quad (5) $$

where $A_{1,t<03/94}$ ($A_{1,t=03/94}$) are the matrices of the first loadings in the pre-(post) March 1994 periods, $z_t$ is vector of the endogenous variables, and $p$ is the lag order of the reduced form VAR. As a next step, a recursive scheme (Luthkepol, 2006) is applied to estimate the variance decomposition matrix of the pre-and-post March 1994 periods using the estimates of $A_{1,t<03/94}$ and $A_{1,t=03/94}$, respectively.

The results of the variance decomposition analysis while accounting for a potential structural shift in VAR parameters are presented in Table IV. The results of the variance decomposition analysis for the pre-and post March 1994 periods are presented at the left-and-right hand sides of the panel, respectively. Overall, our results suggest that prior to the beginning of the 1994 demographic and macroeconomic factors accounted for about 45% of the total variance of the latent factor. Among these variables, the unemployment rate factor, $UN$, appears to have the major impact on the dynamics of the income maintenance households, explaining about 16% of the "long-run" (period 60) variance of the latent factor. Changes in the demographic factors (monthly inflow of the immigrants, $IM$ and the working age population, $POP$) explain about 18% of the latent factor total variance while changes in the state-of economy composite index, $CI$, account for about 3.5% of the latent factor’s dynamics. The financial incentives to work, $RR$, explain 7% of the latent factor variance.

Table IV approximately here

Our findings substantially change as we examine the results of the variance decompo-
sition during the post-March 1994 period. First, it appears that the overall explanatory power of the observable variables has increased from 45% to approximately 50% after the breakpoint date of March 1994. Secondly, and more important, we find that the contributions of the observable variables to the latent factor variance have significantly changed. The explanatory power of the financial incentives to work variable has more than doubled from 7.1% to 14.6%. Also, the explanatory power of the business cycle variables and, in particular, the state-of-the-economy composite index has increased as well. Overall, in the post-March 1994 period changes in the business cycle indicator variables explain about 25% of the total factor variance compared to 20% during the years 1986-1993. On the other hand, the impact of the inflow of immigrants variable on the latent factor dynamics has significantly declined from almost 11% to 4.7%. Since it is financial incentives to work and the inflow of the immigrants which seem to drive the structural shift it would be illuminating to examine the dynamics of these factors around the breakpoint date.

Which events may have triggered a structural shift in March 1994? First, as Figure 2 clearly shows, there is a pronounced structural shift in the dynamics of the RR factor around the years 1994-1995. While prior to 1994 the RR variable seems to fluctuate around some fairly constant level, it exhibits a significant upward trend in the post-1994 period with a pronounced upward shift around the years 1998-1999. This shift may have been driven by a number of factors. First, the indexation rate, \( \delta \), has increased as part of the Single Parent Family Law (1992) and the Reduction of Poverty and Income Inequality Law (1994-1995) legislations \(^8\)(Single Parent Family Law, 5752 and amendments, 5754). Second, this shift can also be partially driven by the deterioration of the alternative wage for the unskilled labor due to the inflow of foreign labour, starting from 1993.

On the other hand, by the year 1994 massive immigration from the former Soviet Union, which reached its peak in the years 1991-1992 subsided. About a million immigrants arrived in the years 1990-1991. This figures become more moderate in the years 1992-1993 with

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\(^8\)For instance, after the Single Parent Family Law the regular rate for household with at least two children increased from 35 to 47.5% of the average wage, while the special rate (those receiving income maintenance for more than two years) increased from 42.5 to 52.5% of the average wage.
less than 80,000 immigrants arriving each year and continued to plunge to less than 60,000 immigrants in 1994 (Israel Statistics Journal, 1990-1995). Consequently, the impact of the immigrants on the labor market become more moderate and the impact of the IM variable on the latent factor dynamics has declined.

8 Summary and Conclusions

As pointed out by Blank (2007), "the history of welfare reform has made it clear that the effects of policy are impacted by the particular time in which it happens to be implemented" (Blank, 2007. pp.23). Thus, a proper understanding of how and to what extent macroeconomic and demographic factors affect the dynamics of the welfare caseload is the issue of major importance both for the economists and the policy makers.

In this paper we develop a new two-step approach for modeling trends in welfare caseload. First, a Lee-Carter (1992) latent factor model was fit to the panel of the income support recipients. Next, we study the relation between the latent factor and changes in the macroeconomic and demographic conditions by the means of time-series techniques. We apply this approach to the data on welfare caseload in Israel over the period 1986-2002.

Our key findings are as follows. First, a single latent factor model provided an exceptionally good fit to the data by explaining, on average, more than 80% of the welfare caseload dynamics. Second, we found that the latent factor and the control variables are cointegrated, which suggests the existence of a long-run "equilibrium" relation between welfare caseload and macroeconomic and demographic conditions. The signs of the cointegrating equation parameters were in accordance with economic theory. Finally, we found that the cointegration equation has undergone a structural shift around the years 1994-1995, following new welfare legislations as well as changes in the labor market structure. Controlling for this shift the financial incentives to work along with the macroeconomic (business cycles) and demographic factors explained about 50% of the dynamics of the latent factor which drives the dynamics of the welfare caseload.
Our findings are of particular interest for Israeli policy makers in light of recent changes in welfare policy. In 2003, the State of Israel prompted a welfare reform that fundamentally reduced the level of cash payments, decreased in-kind services and goods supplied by the public sector and toughened work requirements for recipients of income maintenance. Also, in 2004 the State of Israel prompted a Welfare-to-Work program whose official agenda was to facilitate the transition of welfare recipients from dependence on the social security network to economic and social independence. These changes in Israel’s welfare policy went into effect during a favorable period of the business cycle. The Israeli GDP (per capita) grew 5.2% (3.4%) on average during the time span of 2004-2007 (Bank of Israel Annual Report 2007). Our results indicate the importance of variables such as business cycle in explaining variances in the dynamics of income support recipients, contrary to other studies which have mainly focused on the role of financial incentives to work. Our findings suggest that appropriate welfare policy critically depends on correctly understanding the causal factors affecting the dynamics of welfare caseload. Focusing on only one potential determinant of welfare caseload is likely to provide a misleading picture.

Though our approach relies on the latent factor model and, thus, is fairly general, it is still based on a number of assumptions, some of which may seem too restrictive. Therefore, as robustness checks, we considered a number of alternative model specifications. First, since a Lee-Carter (1992) model implicitly assumes normal innovations, for comparison, we estimated a Poisson bilinear regression model as in Brouhns, Denuit and Vermunt (2002). The estimated factor and the factor sensitivities remained qualitatively the same. Second, the assumption that the trends in the welfare caseload are driven by a single factor may seem too restrictive. Therefore, we also estimated the latent model with two factors. Adding a second factor leaded only to a moderate improvement of the explanatory power of the model. Also, a second factor exhibits no time trend. This finding suggests that a one-factor model appropriately captures a time-trend component of welfare caseload.

There are several potential directions for further research. First, it would be interesting to assess the effectiveness of the recent reforms in reducing welfare caseload. The approach
proposed in this study may be a useful empirical tool for such a purpose. Second, further extensions of our model may be considered, in particular a partially-segmented model, where in addition to a joint factor, the dynamics of each subgroup is driven by the individual factor. This extended model may be a useful framework for studying the dynamics of native versus new immigrant caseload or, in the US case, welfare caseload trends across different states.

References


[49] National Insurance Institute, Annual Survey, various Years (Hebrew).


Table I

Estimates of the Lee-Carter model

<table>
<thead>
<tr>
<th></th>
<th>Levels</th>
<th>Share of population</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>UNEMP</td>
<td>EXEMP</td>
</tr>
<tr>
<td>$\alpha$</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>9.12**</td>
<td>9.49**</td>
</tr>
<tr>
<td></td>
<td>[9.08,9.15]</td>
<td>[9.45,9.52]</td>
</tr>
<tr>
<td>$\beta$</td>
<td>0.42**</td>
<td>0.34**</td>
</tr>
<tr>
<td></td>
<td>[0.39,0.45]</td>
<td>[0.31,0.36]</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.94</td>
<td>0.89</td>
</tr>
<tr>
<td></td>
<td>&lt;0.0001</td>
<td>&lt;0.0001</td>
</tr>
</tbody>
</table>

In This Table we report the estimates of the Lee-Carter (1992) model as described in Sections 4.1 and 5.1. The model is estimated separately with caseload measured in levels and as a share of a total working age population. The estimates are reported under "Levels" and "Share of Population" headings, respectively. Bootstraped 95% confidence intervals are reported in parentheses. The null hypothesis that all $\beta$’s are equal is tested by using a likelihood ratio test. The corresponding bootstrapped $p$-values are reported under $\beta = 0.25$ headings. Asterisks ** and * denote significance at 5% and 10% levels, respectively.
Table II

Results of the Johansen (1991) cointegration tests

Panel A. Trace test

<table>
<thead>
<tr>
<th>No. of CE’s</th>
<th>Statistic</th>
<th>5% cr. value</th>
<th>1% cr. value</th>
</tr>
</thead>
<tbody>
<tr>
<td>None</td>
<td>111.24***</td>
<td>94.15</td>
<td>103.18</td>
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<tr>
<td>At most 1</td>
<td>68.27</td>
<td>68.52</td>
<td>76.07</td>
</tr>
<tr>
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<td>32.1</td>
<td>47.21</td>
<td>54.46</td>
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<tr>
<td>At most 3</td>
<td>13.79</td>
<td>29.68</td>
<td>35.65</td>
</tr>
<tr>
<td>At most 4</td>
<td>5.28</td>
<td>15.41</td>
<td>20.04</td>
</tr>
<tr>
<td>At most 5</td>
<td>0.36</td>
<td>3.76</td>
<td>6.65</td>
</tr>
</tbody>
</table>

Panel B. Maximum-eigen value test

<table>
<thead>
<tr>
<th>No. of CE’s</th>
<th>Statistic</th>
<th>5% cr. value</th>
<th>1% cr. value</th>
</tr>
</thead>
<tbody>
<tr>
<td>None</td>
<td>42.98**</td>
<td>39.37</td>
<td>45.1</td>
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<tr>
<td>At most 1</td>
<td>36.17**</td>
<td>33.46</td>
<td>38.77</td>
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<tr>
<td>At most 2</td>
<td>18.3</td>
<td>27.07</td>
<td>32.24</td>
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<tr>
<td>At most 3</td>
<td>8.51</td>
<td>20.97</td>
<td>25.52</td>
</tr>
<tr>
<td>At most 4</td>
<td>4.92</td>
<td>14.07</td>
<td>18.63</td>
</tr>
<tr>
<td>At most 5</td>
<td>0.36</td>
<td>3.76</td>
<td>6.65</td>
</tr>
</tbody>
</table>

In this Table we report the results of the Johansen (1991) trace (Panel A) and maximum eigen value (Panel B) tests applied to the latent factor and the control variables. Under the null hypothesis of "None" the variables are not cointegrated. Test statistics are compared to the small-sample adjusted critical values as in Cheung and Lai (1993). Asterisks *** and ** denote significance at 1% and 5% levels, respectively.
### Table III

**Variance decomposition analysis**

<table>
<thead>
<tr>
<th>Period</th>
<th>$\triangle f$</th>
<th>$\triangle CI$</th>
<th>$\triangle IM$</th>
<th>$\triangle POP$</th>
<th>$\triangle RR$</th>
<th>$\triangle UN$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>93.4</td>
<td>0.12</td>
<td>0.05</td>
<td>0.26</td>
<td>0.93</td>
<td>5.24</td>
</tr>
<tr>
<td>10</td>
<td>66.9</td>
<td>2.56</td>
<td>6.24</td>
<td>3.92</td>
<td>4.85</td>
<td>15.58</td>
</tr>
<tr>
<td>20</td>
<td>62.22</td>
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<td>7.01</td>
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<td>16.2</td>
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<tr>
<td>30</td>
<td>61.53</td>
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<td>4.62</td>
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<tr>
<td>60</td>
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<td>5.07</td>
<td>7.44</td>
<td>4.75</td>
<td>5.38</td>
<td>16.23</td>
</tr>
</tbody>
</table>

In this Table we report the results of the variance decomposition analysis of the latent factor. The lag-order of the corresponding VAR’s has been chosen by using a Wald test. The contribution of each variable to the variance of latent factor is reported in percentage points under the corresponding heading.

### Table IV

**Variance decomposition-sub-period analysis**

<table>
<thead>
<tr>
<th>Period</th>
<th>$\triangle f$</th>
<th>$\triangle CI$</th>
<th>$\triangle IM$</th>
<th>$\triangle POP$</th>
<th>$\triangle RR$</th>
<th>$\triangle UN$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>85.9</td>
<td>1.9</td>
<td>0.37</td>
<td>0.09</td>
<td>0.17</td>
<td>11.5</td>
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<td>10</td>
<td>59.2</td>
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<td>9.1</td>
<td>6.3</td>
<td>6.8</td>
<td>16.1</td>
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<td>56.5</td>
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<td>9.9</td>
<td>7.1</td>
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<td>30</td>
<td>55.6</td>
<td>3.4</td>
<td>10.5</td>
<td>7.2</td>
<td>7.0</td>
<td>16.3</td>
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<td>60</td>
<td>55.2</td>
<td>3.4</td>
<td>10.8</td>
<td>7.4</td>
<td>7.1</td>
<td>16.2</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Period</th>
<th>$\triangle f$</th>
<th>$\triangle CI$</th>
<th>$\triangle IM$</th>
<th>$\triangle POP$</th>
<th>$\triangle RR$</th>
<th>$\triangle UN$</th>
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</thead>
<tbody>
<tr>
<td>1</td>
<td>94.7</td>
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<td>0.29</td>
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<tr>
<td>60</td>
<td>51.3</td>
<td>8.5</td>
<td>4.7</td>
<td>5.2</td>
<td>14.6</td>
<td>15.8</td>
</tr>
</tbody>
</table>

In this Table we report the results of the variance decomposition analysis for the pre-and-post structural shift periods. The contribution of each variable to the variance of the latent factor is measured in percentage points. The findings for each sub-period are reported under the corresponding headings, respectively.
Figure 1: In this Figure we present the time-dynamics of the welfare caseload for different income support sub-groups. For each sub-group the data is presented in levels (upper plot) and as a share out of a total caseload (lower plot).
Figure 2: Replacement rate over the period of November 1986-March 2003. Replacement rate is calculated as $RR_t = \frac{\delta_t \bar{W}_t}{\bar{W}_t}$ where $\bar{W}_t$ is the average monthly wage in the economy, $\bar{W}_t$ is the alternative wage for unskilled labor and $\delta_t$ is the annual indexation rate to $\bar{W}_t$. 
Figure 3: Lee-Carter latent factor estimated using a Singular Value Decomposition as described in Section 3.1. Dotted lines denote 95% confidence interval estimated using a bootstrap method.
Figure 4: In this Figure we plot the estimated ADF statistics of the Gregory-Hansen (1996) structural stability test. Each point on a time line is considered to be a potential date when a structural shift occurred. Minimum value of ADF is compared to the 10 and 5% critical values, denoted by the dotted lines. ADF statistic is denoted by the solid line.