Israel Economic Review Vol. 8, No. 2 (2010), 67–90

TESTING FOR CONSTANT RETURNS TO SCALE AND PERFECT COMPETITION IN THE ISRAELI ECONOMY, 1980-2006[1]

JOSEPH DJIVRE AND YIGAL MENASHE

Abstract

In this paper we test the hypothesis of constant returns to scale and no market power, implementing Hall's test (1988, 1991) on Israeli annual real business-sector GDP data. The results validate the hypothesis of perfect competition at the aggregate level and are in line with most empirical research in other countries as well. Thus on the one hand they justify the imposition of the CRS restriction when estimating the business-sector GDP production function in Israel and on the other contradict the assumption of CRS with market power and fixed entry costs in the modeling of the Israeli economy.

1. INTRODUCTION

In this paper we test the hypothesis of constant returns to scale and no market power, implementing Hall's test (1988, 1991) on annual real business GDP data for Israel during the years 1979-2006. The test was performed under the assumption of Hicks-neutral technological changes. According to this non-parametric test the acceptance of the null-hypothesis requires that the rate of change of TFP, obtained as a Solow residual (SR) under the null hypothesis, be exogenous to changes in output, whose origin is not in technological changes and which do not affect technological changes.

The issue of divergence from the joint hypothesis of constant returns to scale (CRS) and no market power (hereafter the null hypothesis) has been approached from a second angle as well. This approach originated in the effort to obtain an accurate estimate of the SR by estimating a production function econometrically. The objective of this line of research was to determine the contribution of technological changes embodied in the SR to GDP volatility, thereby empirically testing the RBC theory according to which the procyclicality of the SR mainly reflects productivity shocks. The need for the SR to provide an accurate measure of technological shocks required to account for economies of scale and for input and output mismeasurements in the estimation of the production function. The extent of economies of scale was thus obtained as a by-product of the effort to measure technological

[1] Bank of Israel, Research Department. http://www.boi.org.il; Email: joseph.jdjivre@boi.org.il
Email: yigal.menashe@boi.org.il.
We would like to thank Oshrit Munk and Liron Gat for their research assistance and Roni Frish and the participants in the Research Department seminar for their comments.

change accurately. The empirical research concentrated on the U.S economy, even though some of it related also to other economies.

The results obtained by Hall, using annual value added data between 1953-1984 for 26 industries at the two-digit level, rejected the null hypothesis because his test indicated a correlation between changes in the SR and in exogenous, activity-inducing factors which are uncorrelated with technological change. Following this rejection he estimated the extent of IRS in the U.S economy econometrically. The results obtained justified attributing the procyclicality of SR and its correlation with exogenous activity-inducing factors to the existence of economies of scale. Indeed the derivation of the SR under the erroneous assumption of CRS will give rise to a procyclical residual because output will increase (decrease) more than proportionately to the increase (decrease) of the production inputs in the presence of IRS, giving thereby giving the false impression of positive (negative) productivity shocks.

The rejection of the null hypothesis by Hall in favor of the existence of market power and IRS found support in additional empirical research [Caballero and Lyons (1992) and Roeger (1995)]. In Caballero and Lyons (1992) the origin of the increasing returns was not at the firm level but at the aggregate level and arose from externalities in the form of spillovers.

The empirical research implementing the second methodological approach based on the econometric estimation of the production function did not find evidence of significant returns to scale in the U.S. economy, after accounting for the mismeasurement of output and for fluctuations in the intensity of utilization of labor, as a result of hoarding, and of capital [Burnside et al. (1995), Basu and Kimball (1997) and Basu and Fernald (1997)]. Basu et al (2001) found evidence of limited economies of scale in the production of durables and Basu and Fernald (2001), found evidence of limited economic profits, implying only a small divergence from the perfect competition paradigm in the U.S. economy. This empirical research together with the findings of Burnside and Eichenbaum (1994) reduced the contribution of technological shocks to the volatility of output in the U.S. thus weakening the RBC interpretation of output fluctuations in this economy.

According to Basu et al. (2001) the weakening of the contribution of changes in TFP productivity to the growth of output may have been exaggerated at least in the second half of the 1990s, which was characterized by investment acceleration in information technology. This acceleration seems to have obscured productivity gains as a result of an underestimation of output because of the unmeasured increasing adjustment costs involved in the aforementioned investment acceleration. These costs created a wedge between the higher actual output level and its lower measured counterpart which does not include adjustment costs. [2]

With respect to economies of scale outside the U.S, Paquet and Robidoux (2001) showed, using Canadian aggregate data, that after accounting for capacity utilization, the economies of scale obtained when not adjusting inputs for their intensity of utilization disappeared. Likewise Inklar (2006) obtained that the estimated coefficients for economies

[2] According to Hall (2004) capital and labor adjustment costs for U.S. industries at the two-digit level are relatively small.

of scale in France, Germany and the Netherlands either fell to the vicinity of one or became statistically insignificant after adjusting inputs for their intensity of utilization. Vechi (2000) compared the U.S. and Japanese economies and showed that in Japan procyclical productivity is the result of labor hoarding and that internal returns are higher in the U.S. industries than in Japan; no strong evidence was found of increasing returns in the American economy.

The ability to account for the intensity of input utilization was crucial for the econometric estimation of the production function because its absence and the consequent misspecification of the production inputs give rise to a positive bias in the estimation of the output elasticities with respect to the production inputs, favoring the false rejection of the CRS hypothesis. This is also true of Hall's non-parametric test.

The intensity of input utilization, which is unobservable, was expressed in the empirical literature as a function of observed variables derived from the first order conditions for efficient production. A linear expansion of this function was then introduced in the econometric estimation of the production function instead of the intensity of the input utilization.

With respect to capital utilization Burnside et al. (1995) used electricity consumption as its proxy, assuming that capital services and electricity are used in fixed proportion. Under less stringent assumptions about the production function they showed that the equality of the marginal rate of substitution between labor and electricity, which is also a function of capital services, to their relative price, allows capital services to be expressed as an increasing function of the other two production inputs, labor and electricity consumption, and a decreasing function of the relative price of electricity with respect to labor services. A condition similar to that obtained on the basis of electricity consumption was also derived from the use of intermediate materials when these were considered as production factors alongside labor, capital and energy.

With respect to labor effort Burnside et al. (1993) used as a proxy an expression of income, consumption and average labor hours that was obtained from the first order conditions of optimizing households. Vecchi (2000) used as a proxy for labor utilization the intensity of intermediate materials with respect to hours worked. Basu and Kimball (1997) showed, under the assumptions that labor and capital are quasi-fixed and that the cost of their intensified utilization lies in higher wages for the first and higher depreciation costs for the second, that it is possible to express labor effort as an increasing function of average hours worked, and capacity utilization as a function of observed investment, capital stock, production materials and prices of production materials and new capital (investment goods). They found, however, that the adjustment of capital stock for the depreciation involved in its more intensive utilization was not substantial in the U.S. economy and could be overlooked.

The need to assume the existence of higher costs for higher capacity utilization, in the absence of substantial depreciation costs because of the more intensive use of capital, led Basu et al. (2001) and Basu and Fernald (2001) to assume the existence of shift premia when the capital stock is quasi-fixed in order to obtain an internal solution for capacity utilization. Under this assumption capacity utilization, like labor effort, can be expressed as an increasing function of the average working hours only. This variable was also put forward as a proxy for capacity utilization by Abbott et al. (1998) and has also served us as such in the present paper.

Our findings did not justify the rejection of the null hypothesis for the Israeli economy, after adjusting the data for changes in labor productivity and for variations in the capacity utilization of capital, in line with the majority of the surveyed literature. Our inability to reject the no market power hypothesis justifies the imposition of CRS restrictions when estimating the business-sector GDP production function in Israel, but does not justify the assumption of monopolistic competition among identical firms with fixed entry costs in the modeling of the Israeli economy.

Our main focus in this paper will be on the implementation of Hall's test on business sector GDP production in Israel. We shall therefore refer below to the parametric econometric approach only when necessary for the clarification of our exposition. This paper contains four additional parts: in the second part we present Hall's methodology of testing for the departure from the CRS no market power paradigm, in the third we analyze the data and choose a proxy variable for capacity utilization and instrumental variables exogenous to technological innovations required for the implementation of Hall's test. In the fourth part we present the test results and in the fifth part we conclude.

2. HALL'S TEST

In this section we describe the methodology of Hall's test. The invariance of the Solow residual implies that if a variable exists, hereafter the instrument, which is, say, positively correlated with the rate of growth of GDP per unit of capital, but is not, by construction, correlated with technological change, neither causing productivity shifts nor being caused by them, then its correlation with the rate of change in TFP derived as an SR should be zero under the null hypothesis of CRS and no market power. However under the erroneous assumption that the economy operates under CRS and no market power, while in fact it does not, the derivation of the rate of change of TFP as an SR will include, in addition to the technology component, a factor positively correlated with changes in economic activity which are uncorrelated with technological change, thereby giving rise to a positive correlation between the instrument and the erroneously computed SR. Such a correlation will therefore constitute a proof of the divergence of production from the perfect competition paradigm.

a. The specification of the test

Under the assumption of Hicks-neutral technological changes, the domestic product, Y, produced using capital, K, and labor, L, can be represented by the following functional form:

(1) $Y = A.F(K, L)$.

In the above expression A stands for the technology innovation factor. Expressing the above equation in logs and differentiating with respect to time we obtain that:

$$(2) \qquad \frac{dY}{Y}=\frac{dF}{F}+\frac{dA}{A}=\frac{\partial Y}{\partial K}\cdot\frac{K}{Y}\cdot\frac{dK}{K}+\frac{\partial Y}{\partial L}\cdot\frac{L}{Y}\cdot\frac{dL}{L}+\frac{dA}{A}\ .$$

The sum of the output elasticities with respect to all the production inputs provides a measure of the extent of economies of scale, γ, so that the assumption that price, P_y, is set as a gross markup, μ, over marginal cost, μ.MC=P, allows us to express the aforementioned sum of elasticities, under the assumption that firms are price takers in the inputs market, as follows:

$$(3) \qquad \gamma=\mu\cdot\frac{W\cdot L}{Y\cdot P_y}+\mu\cdot\frac{P_{cr}\cdot K}{Y\cdot P_y}\ .$$

In this expression W stands for the nominal wage and P_{cr} for the cost of capital. Letting the share of the wage bill in the value of domestic product be equal to α_t and the share of the profits above the normal be π_R, we obtain that the share of the capital costs is equal to $(1-\alpha_t-\pi_R)$, and the coefficient of economies of scale is equal to:[3]

$$(4) \qquad \gamma=\mu\cdot\alpha_t+\mu\cdot(1-\alpha_t-\pi_R)=\mu\cdot(1-\pi_R)\,.$$

It transpires from expression (4) that the gross markup, μ, is in general greater than the degree of returns to scale, γ, and that the simultaneous assumption of no market power ($\mu=1$) and CRS (γ=1) implies zero monopolistic rents, $\pi_R=0$ and thereby free competition.[4] It is also apparent from this expression that the existence of IRS (γ>1) must necessarily give rise to market power (μ>1).

According to the above notation and under the assumption that labor and capital services are characterized also by changes in productivity, $\frac{de}{e}$, and in capacity utilization, $\frac{dcu}{cu}$, respectively, over time, then the rate of change of technological innovations is given

[3] A monopolistic rent is sometimes introduced in the guise of fixed costs, say, of entry. In this case gross profits include a monopolistic rent, which is not however appropriated by the capital owners but is used to cover the fixed costs. In this case profits are zero at equilibrium, but the share of the input costs is smaller than one. In all subsequent expressions we have suppressed the time subscript from the gross markup and the labor share in revenue to simplify the exposition. The variability of the labor share over time is consistent, under the assumption that firms are price takers in the inputs market, with a general specification of the production function which is not limited to the Cobb-Douglas type and therefore allows the elasticities of output with respect to the labor and capital inputs to vary over time.

[4] The elasticities of domestic product with respect to labor and capital services are equal to α *and* $(1\text{-}\alpha)$ under the null hypothesis, to $\mu\cdot\alpha$ *and* $(1-\mu\cdot\alpha)$ under CRS and market power, and to $\mu\cdot\alpha$ *and* $(\gamma-\mu\cdot\alpha)$ under IRS.

by expression (5) below:

$$(5)\qquad \frac{dA}{A}=\left[\frac{dY}{Y}-\frac{dK}{K}-\frac{dcu}{cu}\right]-a\cdot(\frac{dL}{L}+\frac{de}{e}-\frac{dK}{K}-\frac{dcu}{cu}).$$

It can be shown that if technological change is erroneously derived on the basis of the RHS of expression (5), when in fact the economy is not characterized by perfect competition, then the rate of change of TFP calculated in this way will be augmented with respect to the true technological change, $\frac{dA}{A}$, by an additional factor, and the measured technological change in the presence of changes in labor productivity and capacity utilization will be equal to:

$$(6)\qquad \frac{dA}{A}+\left[(\mu-1)\cdot a\cdot(\frac{dL}{L}+\frac{de}{e}-\frac{dK}{K}-\frac{dcu}{cu})\right] \text{ under CRS and market power and}$$

to:

$$(7)\qquad \frac{dA}{A}+\left[(\mu-1)\cdot a\cdot\left(\frac{dL}{L}+\frac{de}{e}-\frac{dK}{K}-\frac{dcu}{cu}\right)+(\gamma-1)\cdot\left(\frac{dK}{K}+\frac{dcu}{cu}\right)\right] \text{ under IRS.}$$

If the rate of change of the instruments is, say, positively correlated with the rate of GDP growth per unit of capital services and uncorrelated by selection with the rate of technological change, $\frac{dA}{A}$, it can be shown that it must also be positively correlated with the factors in brackets in expressions (6) and (7).[5] If on the other hand the null hypothesis is correct, the expressions inside the brackets vanish, since then μ and γ are equal to one. As a result a strong correlation between the rate of technological change, derived as a S.R, and the rate of change of the instrument constitutes a proof of departure from the CRS perfect competition paradigm.

It can be also readily deduced from expression (5) that overlooking changes in labor productivity and/or effort and changes in the utilization of capital can lead to the false rejection of the null hypothesis if these changes are procyclical. More precisely, in such a case the measured rate of technological change will include in addition to the actual rate of technological change $\frac{dA}{A}$ an additional component made up by positive expressions of $\frac{de}{e}$ and of $\frac{dcu}{cu}$. If the latter are procyclical, then the rate of change of the instrument will turn

[5] This conclusion is straightforward when the gross markup is assumed to be constant. Hall (1988) has shown under more general conditions that when the gross markup is allowed to fluctuate even though it is on average equal to one, the covariance between changes in the instrument and in the measured technological changes will be zero or slightly negative under the null hypothesis, so that a positive covariance and correlation attests to the divergence from the null hypothesis paradigm.

out to be correlated with the *measured* rate of technological change even though it will not be correlated with the *actual* rate of technological change.

The intuition behind this result may be summarized as follows: if a lower, say demand for goods is accompanied by a lower capacity utilization, CU, and/or lower labor effort, *e*, then having overlooked these two factors we would expect to obtain output higher than the output actually produced. Under these circumstances, if we do not account for the lower capacity utilization and labor effort, the "lost" output is attributed to a negative productivity shock, establishing a positive correlation between the exogenous fall in demand growth, and hence in the rate of growth of the instrument, on the one hand, and the fall, under the null hypothesis, in TFP growth, on the other. Such a result may lead to the false rejection of the null hypothesis.[6] A similar result may be reached if the mismeasurement refers to pro-cyclical labor productivity.

If on the other hand these expressions are countercyclical, then we risk falsely accepting the null hypothesis, since in this case the positive, say, correlation between the rate of technological change $\frac{dA}{A}$ and the rate of change of the instrument will be mitigated by the negative correlation between the latter and the omitted changes in labor productivity and effort and in capacity utilization.

In order to determine whether there is a correlation between the rate of change of the instrument denoted here by *dz*, and the rate of technological change obtained under the null hypothesis as an SR, Hall (1988) ran a regression of the latter on *dz* without considering changes in labor effort and/or changes in the intensity of capital utilization:

$$(8) \qquad \frac{dY}{Y} - \frac{dK}{K} - a \cdot \left(\frac{dL}{L} - \frac{dK}{K}\right) = c_1 + c_2 \cdot dz \ .$$

A statistically significant regression coefficient, c_2, with the same sign as that of the correlation between changes in the instrument and changes in the GDP/capital ratio constitutes a proof of the statistically significant correlation required for the rejection of the null hypothesis.[7]

b. The false rejection or acceptance of the null hypothesis

The false rejection or acceptance of the null hypothesis can be traced to the mismeasurement of the intensity of the utilization and of the productivity of the production inputs, of overhead costs in production and of output. The output mismeasurement in its

[6] A similar outcome is obtained when implementing the parametric methodology of production function estimation. The omission of the change in intensity of utilization of the two inputs creates a positive bias in the estimation of the coefficients of the growth rate of the labor and capital services input because of the positive correlation between the existing regressors, especially in the case of labor input, and the existing regressors, which reflects the pro-cyclicality of input utilization. This positive bias provides an overestimate of the elasticity of output with respect to the production inputs facilitating the false rejection of the null hypothesis in favor of the existence of IRS.

[7] The variables used by Hall (1988, 1991) as instruments were the American President's party and the rate of change of the government's military expenditure, and of the oil price.

turn can be traced to the overlooking of investment adjustment costs, as a result of which the aforementioned correlation between the rate of change of the instrument and the measured rate of growth of the Solow residual is mitigated.

In the previous section we provided both an intuitive and a formal explanation of why the mismeasurement of changes in the utilization intensity and in the productivity of production inputs can lead to either the false rejection of the null hypothesis or to its false acceptance.

It can be shown that overlooking labor overhead costs when implementing Hall's test does not lead to the erroneous rejection of the null hypothesis, contrary to the outcome of the econometric estimation of the production function which will tend to falsely reject the null-hypothesis in the presence of overhead costs. [8]

Because of investment adjustment costs, services rendered within the firm by its labor force may not be considered as an output while the corresponding labor input going into the provision of these services is considered as part of the labor cost. At the aggregate level this shortcoming is reflected in a lower value for investment. If during upturns in economic activity which are not triggered by technological changes economies experience an acceleration in investment reflected in a rising investment to capital ratio, which implies higher investment costs, then the measured domestic product will fall short of its actual level and the one implied by the production inputs, if investment adjustment costs are overlooked. As a result, output mismeasurement will be reflected in a fall in TFP. The procyclicality of the changes in the SR when the null hypothesis is false could be thus obscured by the countercyclical effect on TFP changes of accelerated investment during expansions, leading to the false acceptance of the null hypothesis.

A similar result of underestimation of economies of scale is obtained under the parametric approach of the econometric estimation of the production function. If the adjustment cost factor is omitted from the list of regressors, then the estimates of the coefficients of the rate of growth of the production inputs labor and capital will be biased. The estimation bias should be negative because the correlation between the existing procyclical regressors and the missing one is negative. This negative bias gives rise to underestimates of the output elasticities with respect to the production inputs, thereby supporting the false acceptance of the null hypothesis.

We may therefore conclude from the above analysis that the implementation of Hall's test requires accuracy in the measurement of both output and the effective production inputs.

3. THE TEST AND THE DATA

In this section we present the data and the variables relevant for the implementation of Hall's test. We based our hypothesis testing, like Hall (1988), on a regression in which the

[8] For a more formal analysis of the issues presented in this and in the previous section see: Djivre and Menashe (2009).

rate of the technological change derived as an SR served as a dependent variable and the rate of change of the instrument as a regressor.

We have assumed that production takes place in two stages so that the output of the business sector output is produced using intermediate goods and value added, which in its turn is produced by labor and capital. We have concentrated here on implementing Hall's test on the latter. Hall's original paper of 1988 included 32 annual observations between 1953 and 1984 and covered 26 industries at a roughly two-digit level according to the SIC classification. Our sample period included 28 annual observations between 1979 and 2006 at the aggregate production level.

The econometric estimation of economies of scale based on value added data, when the objective is to determine the deviation from the perfect competition-CRS paradigm with respect to the output function, may introduce a bias [Basu and Fernald (1997)]. This bias is the outcome of the omission of a regressor in the estimation of the production function which depends on the import intensity with respect to output. If this intensity is pro-cyclical, as is usually the case, then the coefficient obtained for economies of scale on the basis of value added data will be biased in favor of IRS. Such a variable will be also missing if we implement Hall's test and want to draw conclusions about the production of output based on value added data. However, this is not the case at hand since we are interested in testing the departure from the perfect competition paradigm of the value added (GDP of the business sector) production function using value added data and not data of the output function.[9]

We describe below in detail the adjustment of the data to prevent mismeasurements of the nature mentioned in the previous sections, and the choice of the exogenous variables, the instruments, which are correlated with economic activity and uncorrelated with changes in technology.

a. Measuring GDP and the labor share

The derivation of the rate of technological change as an SR under the null hypothesis is based on the calculation of the RHS of expression (5), following adjustments to prevent mismeasurements of the GDP and of the production inputs.

In Israel, investment statistics include, according to the CBS data, a large part of the costs accrued to the firm to render the new capital operational, such as transportation and installation costs of new equipment. As a result, the extent of the divergence we pointed to above, between the reported GDP on the one hand and actual GDP inclusive of investment adjustment costs on the other, seems to be limited. We therefore consider the risk of falsely

[9] An additional bias may be introduced when using aggregate data if the assumption of a representative firm is false. Aggregation over diverse sectors and firms gives rise to a factor measuring the reallocation of inputs across firms in the same industry or across industries. This factor is usually omitted when estimation is based on aggregate data because of the underlying assumption of the single representative firm. The divergence from the representative firm paradigm creates a bias in the estimation of the average returns to scale, as a result of the aforementioned factor, whose direction constitutes an empirical issue [Basu and Fernald (1997)]. Paquet and Robidoux (2001) have used directly aggregate GDP data on the Canadian economy without resorting to the aggregation of sectorial data.

accepting the null hypothesis because of the pro-cyclical underestimation of the TFP growth to be limited.

With respect to the measurement of the effective production inputs, we have to account for changes in their productivity and in their intensity of utilization. Our use of capital at constant prices accounts to a large extent for its increasing productivity, leaving us to handle the accurate measurement of labor productivity and the intensity of utilization of labor and capital. According to our exposition, when implementing Hall's test, we should not be preoccupied with the issue of overhead costs which affect the effective labor input.

Our derivation of the rate of technological change as an SR requires as an input the rates of growth of the business sector GDP and of the effective production inputs, adjusted for changes in the capacity utilization of the capital stock and in labor productivity and effort.

In addition to these variables, the labor share in GDP is also required separately for every observation in our sample. A variable labor share over time is consistent with our assumption that the production function is not necessarily of the Cobb-Douglas form.

The GDP should be expressed at factor cost (producer prices) and not in market prices since these are the prices relevant for the optimizing firm. The GDP data in fixed producer prices are however available only after 1995 even though GDP data at current producer prices exist for earlier dates. We as a result, were constrained to express GDP growth in terms of constant market prices, since our sample starts from 1979.[10] The labor share, which was obtained as the ratio between the level of the wage bill and GDP, was however available from 1979, since its GDP component, expressed at factor cost, was in terms of current prices. The components of the labor share data are reported by the CBS on an annual basis and we were consequently obliged to limit the implementation of Hall's test to annual observations.

b. Measuring the effective production inputs

(1) Accounting for changes in labor productivity

A substantial part of our sample period was affected by the absorption of an unprecedented wave of immigration from the end of the 1980s leading to a 28 percent increase in the labor force between 1991 and 2006. In view of the fact that this immigration wave adversely affected the productivity of labor [Friedman and Zussman (2008)], during an upswing in economic activity, our overlooking its effect on the rate of growth of the effective labor input could have created a negative bias in the calculation of changes in TFP, thus favoring the acceptance of the null hypothesis, according to our previous analysis. As a result, we adjusted the labor input data to account for the aforementioned fall in productivity.

Our adjustment of the labor input for changes in labor productivity is based on a productivity index derived by Friedman and Zussman (2008) on the basis of wage differentials and the assumption that labor's earnings are commensurate with its marginal

[10] The correlation coefficient between the GDP growth at fixed producer prices and market prices between 1995 and 2006, a period for which GDP data are available under both definitions, is equal to 0.9995 implying that our use of GDP data at fixed market prices for our entire sample, because of the absence of data at fixed producer prices, has not, most probably, impaired the accuracy of our calculations.

productivity. If the wage differentials identified by Friedman and Zussman do not reflect productivity differentials but discrimination against new immigrants, then by adjusting the labor input for a fall in productivity which did not occur, we would have overstated the productivity gains during a period of expansion in economic activity. This implies that the mistakenly measured productivity gains are procyclical and hence are expected to be positively correlated with changes in our instrument, favoring the rejection of the null hypothesis. If under these adverse circumstances our test still cannot reject the null hypothesis, then the case for CRS and competition becomes even stronger.

(2) Measuring production inputs and their intensity of utilization

In this section we describe the characteristics of the variable we chose as a proxy for capacity utilization (CU) and examine its suitability to serve as such by examining whether it is characterized by properties that CU should fulfill and whether its evolution over time resembles that of indices of CU obtained by different methodologies in other countries.

Capacity utilization is in general considered to be pro-cyclical. The economic intuition behind the pro-cyclicality of capacity utilization and labor effort runs as follows: If the production inputs cannot be readily adjusted to their optimal level following an activity augmenting shock, then the, say, higher capital or labor services required to satisfy optimality conditions after the shock will be provided by increasing the intensity of their utilization, even if such an increase is costly, in this case higher costs mitigating the extent of the increase in the intensity of input utilization. This is due to the fact that under efficient production the producer will equate the marginal cost of production expansion by increasing labor employment and/or by raising the capital stock of his/her firm through investment to the marginal cost entailed in longer working hours. In other words higher investment adjustment costs induce efficient firms to increase the intensity of utilization of the production inputs in the margin [Basu and Fernald (2001) and Basu et al. (2001)]. If these adjustment costs are an increasing function of the investment/capital ratio, I/K, we should expect capacity utilization to increase alongside this ratio. Moreover, by the same token, the marginal cost of longer working hours, which carry higher compensation because of the disutility involved in them, should be equal to that induced by the more intensive utilization of, say, capital because of shift premia. In other words capacity utilization should co-vary with the investment/capital-stock ratio and longer working hours per employed person.

Indeed Basu and Kimbal (1997), Basu et al. (2001) and Basu and Fernald (2001) show, under not very restrictive assumptions, that these efficiency conditions imply that labor effort and capacity utilization should be an increasing function of the average number of hours per worker if longer working hours per worker imply higher wages. This allows unobservable variables of intensity of input utilization to be expressed as an increasing function of the average length of the working week, which, in line with the analysis above, should be also procyclical and positively correlated with the I/K ratio similarly to CU.[11]

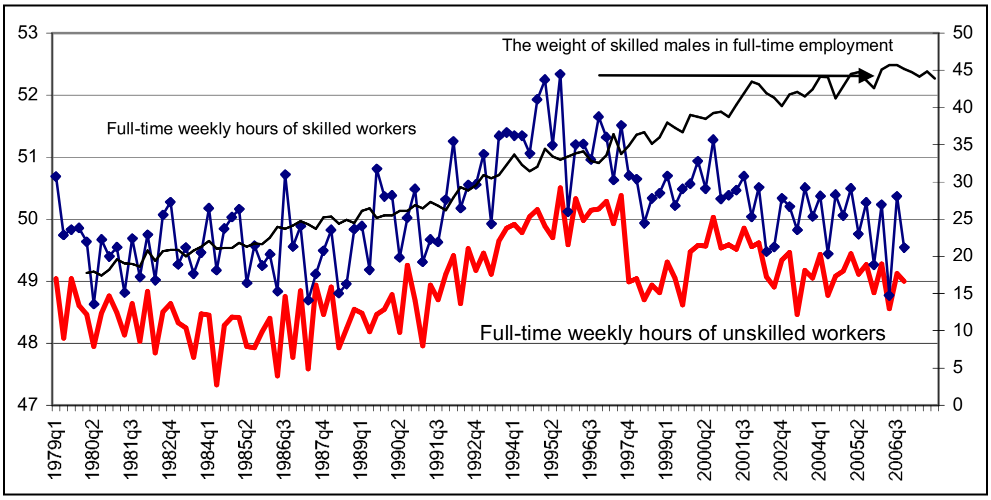
[11] This is true under the assumption that the number of shifts is given, otherwise it is possible that the opening of an additional shift could lead to lower weekly working hours per worker.

If the length of the working week is to capture the effect of changes in the intensity of utilization of inputs, then it must not be "contaminated" by changing trends due to structural changes in the Israeli economy.

The rising share of women in employment over time alongside their lower average weekly hours than men's are expected to have had a negative effect over time on the aggregate average weekly hours. Our use of average working hours as a proxy for capacity utilization should not therefore be a weighted average of the data with respect to gender and full and part-time jobs.

The data indicate that the average number of weekly hours of fully employed males fluctuated around 48.5 hours between 1979 and 1989, and after that date it increased and stabilized after 2002 at a slightly higher level of 49 weekly hours. A closer analysis of the data reveals that this slight rise in the average weekly hours of full-time male employees reflects, among other things, the increasing share of skilled male workers over time (Figure 1) and the fact that in general they work longer hours than the less skilled male workers, and the slight rise in the length of weekly hours worked by unskilled male employees after 2002 compared to the pre-1989 period.[12]

Figure 1
Average weekly hours of full-time skilled and unskilled workers and the weight of skilled males in full-time employment



[12] As skilled workers, in terms of their education, are defined employed persons with 13 years or more of formal education. The hour length of the working week is more or less similar across sectors for fully employed skilled men converging after 2002-03 to its pre-1989 level with the exception of manufacturing and the banking and financial sector. In these two sectors the number of weekly hours stabilized after 2002 at a higher level than the one which prevailed in the late 1980s. The observed rise in full-time skilled man-hours in manufacturing and in the financial sector was only slightly reflected in the average weekly skilled man-hours at the aggregate level as a result of the fall in the weight of manufacturing and of the financial sector over time in full-time employed skilled male employees.

To neutralize this composition effect we chose as a proxy for capacity utilization the average weekly hours of full-time skilled male employees. This variable converged to its pre-1989 level after 2002-03 (Figure 1). When measuring the rate of growth of the labor input to derive the SR we have to include all employed labor; however, when deriving a proxy variable for the intensity of input utilization as we have done here it suffices for this variable to be correlated on theoretical and empirical grounds to this intensity and it does not need to encompass all employed labor. For this reason we relied on the National Accounts statistics which cover Palestinians and foreign workers as well as Israeli workers to calculate the growth in labor input, and we used CBS labor survey quarterly data which cover only the Israeli labor force in order to derive our proxy.

The cyclical characteristics of the proxy variable

According to the analysis in the previous sections, the variable we chose as a proxy for capacity utilization should be pro-cyclical and positively correlated with the investment/capital ratio. We show below that this proxy variable indeed co-varies over time with the business cycle and that it is positively correlated with the investment/capital ratio. However the fluctuations in capacity utilization derived from this variable are more moderate in Israel than the fluctuations in capacity utilization indices in other countries.

To examine the similarity between the cycles and especially their turning points based on CU and other variables measuring economic activity, we derived the cyclical component of the proxy variable using spectral analysis of quarterly data after subtracting the sample mean and after defining as cyclical frequencies those between 6 and 55 quarters. The reasoning behind the derivation of such a cyclical component is for it to serve as a basis for the construction of a CU index around the mean which reflects full capacity. The cycles on the basis of the cyclical component of CU obtained from spectral analysis did not diverge considerably from cycle classifications with respect to economic activity in the Israeli economy [Melnick (2002)] and Marom et al. (2003)]. Exceptions to this pattern were the expansions and recessions in the first half of the 1980s and the recession between 1990 and 1993 identified by the CU index but not by any other methodology of cycle classification in the Israeli economy.

More precisely, the cycle classification based on CU gives rise to only one recession and expansion between 1979 and 1987 while in the case of Melnick (2002) and Marom et al. (2003) there are two and three expansions and recessions respectively during this period.[13]

The business sector seasonally adjusted GDP growth data on the basis of which we may identify the business cycles in the Israeli economy indicate that the last quarter of 1982 was an outlier characterized by an exceptionally high rate of growth. This outlier may have

[13] Marom et al. (2003) used two approaches in classifying the business cycles in Israel on the basis of the monthly composite index for economic activity in Israel. The first is based on the probability of a recession obtained from the index data, and the second from the deviation of the index from its long term trend. While it is the first methodology which gave rise to the numerous expansion and contraction periods between 1979 and 1987 the second methodology assigned years 1980 and 1981 to expansion periods. This fact is particularly disconcerting because these years were characterized by the second oil shock.

stood behind the classification of the period between 1980 and 1983 and the periods between 1982.10 and 1983.8 and between 1980 and 1988 as expansions by Melnick (2002) and by Marom et al. (1) (2003) and Marom et al. (2) (2003) respectively. If we do not consider this outlier, we will obtain, using Melnick (2002) and Marom et al. (2) (2003), a recession between 1979 and 1984-1985 and an expansion after that date and until 1987. The cycle classification according to the CU index is also consistent with one recession and expansion period between 1979 and 1987, but the end of the recession is anticipated at the second half of 1982 from 1984-85.

The period between two consecutive peaks in 1990 and 1993 identified only on the basis of the CU index, the other classifications having assigned this period to expansions, points to a more profound difference between the CU cycle classification and that of the other methodologies. While deviations of the demeaned CU from zero reflect also deviations of output from its potential level, given the available quantities of labor and capital stock, this is not so with deviations of GDP growth from its trend. This argument becomes clearer when considering the immigration wave in the beginning of the 1990s. When the GDP growth rate is detrended what is obtained, during the immigration years, is a higher level of activity than in normal times, i.e., an expansion. However the influx of immigrants and their absorption into employment immediately increased not only the actual but also the potential output. Under these conditions a fall in the cyclical component of working hours could be consistent with production being below its potential level in spite of the observed high rate of growth. We consider therefore the CU cycle classification to be closer in content to the definition of business cycles in which expansions and recessions refer to periods in which current production is above or below its potential.

While the cyclical component of the full-time skilled man-hours defines cycles in economic activity in Israel similarly to other classifications, the expected positive correlation, according to previous analysis, between this proxy variable for CU and the investment to capital-stock ratio seems to transcend business cycle frequencies. Indeed the low frequency component of the average weekly hours of skilled males, between 55 and 111 quarters, follows an evolution pattern between 1979 and 2006 (Figure 2) resembling that of the long-term evolution of the I/K ratio (Figure 3).

This observation implies that our CU index should not only include the cyclical component of our proxy variable but also its low frequency component, which has co-varied with the long-term evolution of the investment/capital ratio.

Since we performed Hall's test on annual data we used the annual quarterly average of weekly full-time skilled man-hours to construct our proxy for CU. This average filters the seasonal high frequency component of our proxy variable and thereby measures only its cyclical and low frequency component.

The correlation coefficient between the I/K ratio and the length of the working week of skilled employed males in the business sector is 66 percent for equipment and 70 percent for total investment and capital for the period between 1979.q1 and 2006.q4. Between 1995.q1 and 2006.q4, the period for which sectorial-data on investment and capital stock are also available, the corresponding correlations for the business sector are 76 percent and 77 percent respectively, in line with our expectation to find a substantial positive correlation between the two variables. This correlation at the aggregate level is reflected in

Figure 2
The low frequency component of the weekly hours of skilled male workers

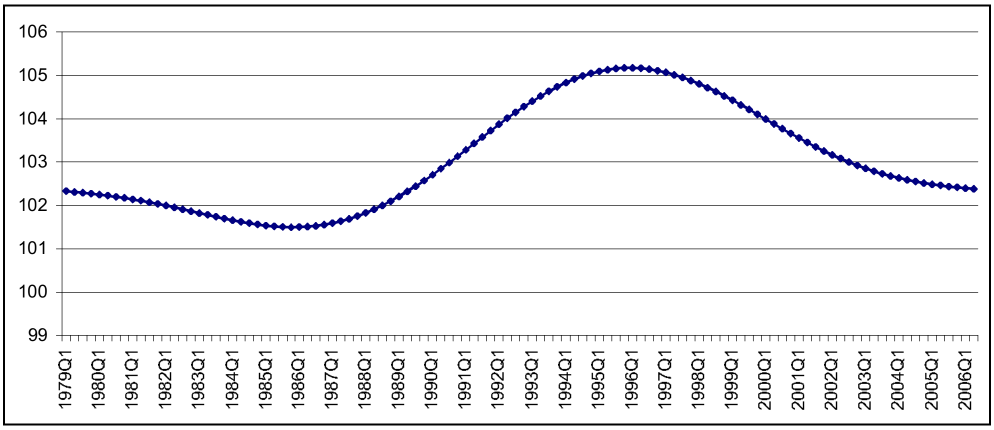
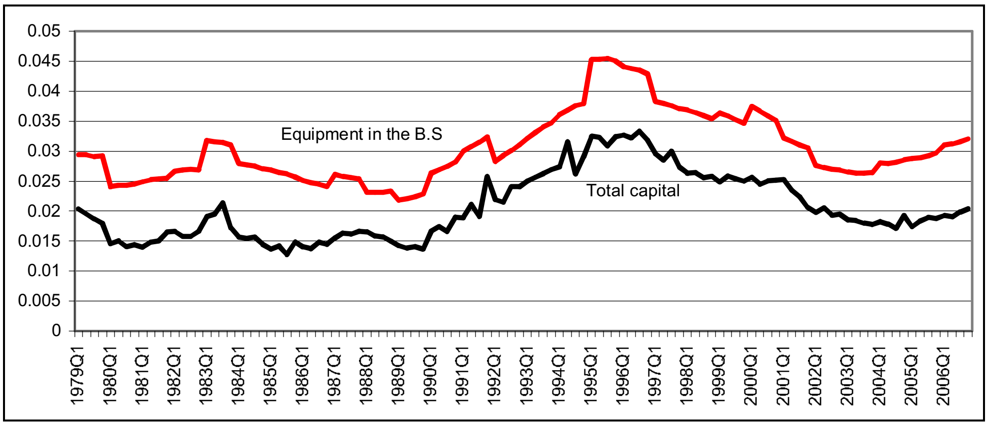


Figure 3
The I/K ratio in Israel (the aggregate quarterly data)



the co-movement between the investment/capital ratio and the average weekly skilled man-hours: the full-time skilled man-hours are characterized by a positive trend since 1988-89 while the trend of I/K becomes positive around the beginning of 1989. This period coincides with the beginning of the end of the downturn after the 1986 stabilization and of the immigration wave. Prior to that date the I/K ratio replicates the cycles between 1979 and 1983, 1983 and 1986 and 1986 and 1989.

The beginning of the falling trend observed in the average weekly skilled man-hours between 1996 and 1997 was accompanied by a fall in the I/K ratio during the same period. Following its fall between 1996-97 the I/K ratio stabilized between 1998 and mid-1999, increasing slightly in 2000 and renewing its negative trend after 2000, stabilizing between 2003 and 2004 and increasing thereafter. Similarly, after its fall between 1996 and 1997, the length of the working week of full-time skilled men stabilized at a lower level between

1998 and 1999, increased slightly between 1999 and 2000 falling after 2000 to its 1989 level. While the evolution of I/K replicates that of economic activity during the whole surveyed period, the length of the working week does not account for the upturn after 2003.

It should be mentioned here, however, that the interdependence between the two aforementioned variables is weaker at the sectorial level for the period 1995-2006, in spite of the high correlation at the aggregate level. The corresponding correlations fluctuate between 40% and 47% in manufacturing with respect to investment in equipment and aggregate investment respectively, and between 50% and 70% in services. [14]

Inter-country comparison

Our CU index refers to the business sector as a whole during our sample period. This is also true about the U.S.A., where in addition to manufacturing there is also a capacity utilization index covering all industries. In the EU on the other hand the data available refer to capacity utilization in manufacturing only. In table 1 below we compare the maximal peak to trough amplitude of the capacity utilization index in Israel and in selected developed economies. In the case of Israel the data we report refer only to the cyclical component of our index derived from the average weekly man-hours of skilled workers and do not include the low frequency component of the average weekly hours which exhibited a prolonged cycle, most probably as a result of the protracted adjustment of capital stock following the huge influx of immigrants. The data for Israel cover both the business sector and the manufacturing industry, as do the capacity utilization data for the U.S. economy. The data for the EU economies cover only the manufacturing industry. The data reported indicate a maximal amplitude of CU fluctuations in Israel that is small compared to that of indices in other economies even in terms of the manufacturing sector in which the extent of

Table 1
Domain of fluctuation in capital capacity utilization in selected developed economies
(percentage points)

Large economies						Small open EU economies		
U.S.A*	EU	Germany	France	U.K	Israel	NL	IRL	B
6.5-16.9					2.3			
8.4-14.6**	7.5**	12.0**	10.3**	9.0**	3.0	5.9**	7.1**	7.7**

* The first number for the U.S. economy reflects the fluctuation between the peak in 1988-89 and the trough in 1994-95, while the second number the corresponding difference between the peak in 1994-95 and the trough in 2008-09.
** Capacity utilization in manufacturing industries. Sources: The Federal Reserve Board for the U.S.A. data and the Business and Consumer Survey of the European Commission for Economic and Financial Affairs. The registered peak for the EU countries reflects activity in the period 1987-91 and the trough in the period 1991-95.

[14] The observed correlation reflects the co-movement between the I/K ratio and the length of the working week and not that between adjustment costs, which depend on the I/K ratio, and the length of the working week. Moreover this correlation can also reflect the effect of other economic factors which may have mitigated the correlation arising from the direct interdependence between the two variables.

fluctuations in Israel is more substantial.[15] This difference is rather pronounced if we take into consideration that the reported peaks and troughs reflect average CU values during the corresponding periods.

c. The selection of the instruments

We chose our instruments from a list of variables which are ex ante expected to be correlated with GDP growth and uncorrelated with the rate of technological change. This list included the rates of change of the inflow of new immigrants, of an index of terror attacks, of tourist inflow and of its lagged value, of the oil price inflation, of the volume of world trade and of the volume of government civilian purchases.

To choose the best-suited instruments from the above list, we estimated regressions in which the rate of growth of GDP per unit of capital served as a dependent variable and the instruments as a regressor and chose the instrumental variables on the basis of the significance of their coefficient. [16]

If the rate of technological change is calculated as an SR on the basis of the RHS of expression (5), that is under the null hypothesis which is not however valid, then, as we have shown in section 2, the measured rate of technological change will include a factor which will be correlated with the rate of change of the instrument to the same extent that the latter is correlated with the rate of change of GDP per unit of capital.

Even though this correlation between the measured rate of technological change and the rate of change of the instrument derives, theoretically at least, automatically from the latter's correlation with the rate of change of the GDP per unit of capital, we also made sure that the rates of change of the chosen instruments positively correlated with the rate of change of labor per unit of capital adjusted for changes in capacity utilization and labor productivity, so that a value for μ greater than unity will give rise ex post to a positive correlation between the *measured* rate of technological growth in expression 6 and the rate of change of the instrument, allowing the rejection of the null hypothesis. We did this by estimating regressions in which the growth rate of labor per unit of capital, adjusted or unadjusted for changes in labor productivity and capital utilization, served as dependent variables and the instruments under consideration as regressors.

[15] If we relate also to the low frequency component of our index, the amplitude of the fluctuations in Israel increases to 5.4 percent for the business sector and to 6.5 percent for the manufacturing sector, which are still low relative to other developed economies. Bregman and Marom (1998) used the electricity consumption per unit of capital as a proxy for capacity utilization in the manufacturing industry. However, the use of electricity per unit of capital relies on the rather strong assumption of a Leontief production function. Under production functions other than Leontief, energy consumption does not constitute an accurate measure of capital CU because the flow of capital services does not depend only on energy consumption but also on the relative price of electricity with respect to labor services.

[16] Our definition of capital includes buildings, equipment, vehicles and intangible assets in the high-tech industries as of 1995.

The instruments which passed the test, whose results are reported in Appendix 1 (Table A.1.1), was the rate of growth of tourist inflow, hereafter *Dtour* and the rate of growth of government civilian purchases, hereafter *D_govcicil*.[17]

Based on these two variables we also created "two-stage" instruments using the fitted values of regressions in which we run the rate of change of the GDP per unit of capital on both of the two above instruments. We calculated two such "two-stage" instruments. The first one was obtained by calculating the fitted value of GDP per unit of capital without adjusting for changes in labor productivity and capacity utilization, while the second was the fitted value obtained for the GDP per unit of capital adjusted for changes in capacity utilization and labor productivity. We performed all the tests using both instruments but given the very high correlation between them, 99.9%, and the similar results, we report below only the results relating to the latter, hereafter *Dinstr*.

When the test is performed using the SR obtained from the business sector GDP growth rate per unit of capital without adjusting it to changes in capacity utilization, the,*measured* technological change in expressions (6) and (7) under the erroneous null hypothesis will be further augmented by expression $(1-a)\frac{de}{e}$, which is positively correlated with the instruments, especially with *D_govcivil*, and will thus lead to rejection of the null hypothesis, but for the wrong reason.

4. THE TEST RESULTS

We performed Hall's test twice. Once when the derived technological growth between 1979 and 2006 was not adjusted for changes in labor productivity and capacity utilization and a second time after making the required adjustments in the labor and capital services, so that overall we estimated six regressions. We report the test results in Table 2 below and in a more detailed manner in appendix 2.

Table 2
The significance of the instrument coefficient

Instrument	Non-adjusted Δ(TFP)	Adjusted* Δ(TFP)
Dtour	t=1.00, P=16.5%	t=0.79, P=22%
D_govcivil	T=0.96, P=17%	t=0.46, P=32%
Dinstr	t=1.36, P=9.25%	t=0.92, P=18%

* The rate of technological growth has been adjusted for changes in labor productivity between 1988 and 2005 and capacity utilization. The P-values reported are for a one-tailed test.

The test results for the whole sample do not justify the rejection of the null hypothesis regardless of whether we adjusted the data to changes in capacity utilization and labor

[17] The variable of tourist inflow is not a completely exogenous variable because it is also affected by endogenous factors such as the real exchange rate which is affected by technological growth in the tradables sector.

productivity. After such an adjustment the test results distanced themselves from the rejection domain even more, as expected, the significance of the instruments' coefficients falling further relative to their level for the unadjusted SR growth rate.

In order to examine the robustness of our results we estimated two types of rolling regressions for each of the specifications whose estimation results are reported in Table 3. In the first type we allowed for the sample period to contract by one year each time leaving the end period fixed at 2006. In this way in addition to the original regression we estimated six additional regressions whose starting points ranged from 1980 to 1985. In the second type of rolling regression estimation we defined a sample period of twenty one observations, the initial sample starting in 1980 and ending in 2000 shifting the whole sample each time by one period so that the last sample covered the period between 1985 and 2006. The estimation results are presented graphically in Appendix 2 with a 95% confidence interval.

The only instrument for which we obtained a statistically significant positive coefficient was *D_govcivil* and this only in the case when the SR was not adjusted for changes in labor productivity and in capital utilization. In fact the significance of the corresponding coefficient fell below the level consistent with a 5% significance after we adjusted the SR for changes in the capacity utilization of capital and in labor productivity. We attribute, as a result, the rejection of the null hypothesis for the unadjusted SR to the high correlation between the rate of change of this instrument and the expression $(1-a).\frac{de}{e}$ which may lead, according to our analysis, to the false rejection of the null hypothesis, if the TFP growth rate obtained as an SR has not been adjusted for changes in capacity utilization. Had the rejection reflected a correlation with the residual because of factors measuring deviations from the perfect competition paradigm, the null hypothesis should also have been rejected after adjusting our SR for changes in capacity utilization and labor productivity.[18]

5. CONCLUSIONS

We consider that our results provide evidence of the existence of CRS with no market power in the Israeli economy at the aggregate level. Because of the proximity, sometimes, of the results to the rejection limit we consider that additional research is required in the direction of the accurate measurement of effective capital services and labor input. Our inability to reject the perfect competition hypothesis at the aggregate level, in line with the majority of the surveyed literature, justifies imposing the CRS restriction when estimating the aggregate production function for GDP in Israel econometrically, but creates some skepticism about the assumption of monopolistic competition among identical firms with entry costs and CRS characterizing DSGE macroeconomic models of the Israeli economy.

[18] The test results for the lagged value of the rate of change of tourist inflow gave rise to a significant correlation with the SR. Given that this variable failed the instrumental variable test, not being found to be correlated either with the business sector GDP growth rate per unit of capital or with the rate of change of the labor input per unit of capital adjusted for changes in capacity utilization and labor productivity, we could not attribute this correlation to a departure from the perfect competition paradigm.

APPENDIX 1

Table A.1.1
The Selection of the Instrumental Variables:
The Regression Coefficient of the Instruments 1979-2006*
(Quarterly Data)

The Dependent Variable	The Instrumental Variable		
	Dtour	D_Govcivil	Dinstr
DY/Y-DK/K	t=3.05, P=0.25%	t=1.65, P=5.5%	t=3.75, P=0.04%
DY/Y-DK/K-Dcu/cu	t=2.31, P=1.47%	t=1.32, P=10%	t_{dtour}=2.34, P=1.4% t_{dGov}=1.41, p=8.5%**
(1-Labor_share).Dcu/cu	t=1.60, P=6.16%	t=2.30, 1.5%	t=2.7, P=0.6%
Labor_share.De/e	t=0.26, P=40%	t=1.2, P=12.1%	t=0.96, P=17.2%
Labor_share.(DL/L+De/e-DK/K-Dcu/cu)	t=2.69, P=0.6%	t=1.52, P=7.0%	t=3.3, P=0.13%

* t- values in parenthesis. The probability P is for one tailed tests.
** The significance of the coefficients of dtour and D_govcivil in the regression serving to construct Dinstr as the fitted value of the rate of growth of GDP per unit of capital adjusted for changes in capital utilization.

APPENDIX 2

Definition of the instrumental variables.

D_govcivil: The rate of growth of government civil purchases.

Dtour: The rate of growth of tourist inflow.

Dinstr: The fitted value of the rate of growth of the Business Sector GDP per unit of capital, adjusted for CU and labor productivity, in the regression with the rate of growth of tourist inflow and government civil purchases as regressors.

APPENDIX 2

The 5% confidence interval around the regression coefficients

Contracting rolling regression results

Figure A1.1: The D_govcivil coefficient with unadjusted technological growth rate for changes in CU and labor productivity

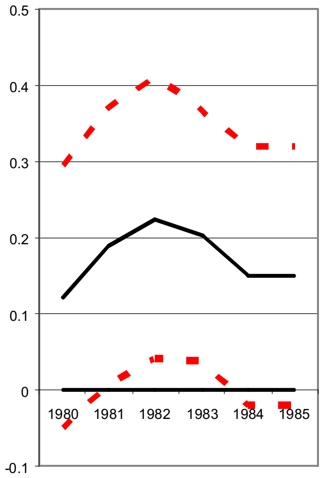


Figure A1.2: The Dtour coefficient with unadjusted technological growth rate for changes in CU and labor productivity

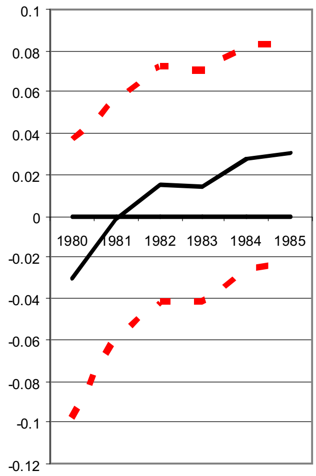
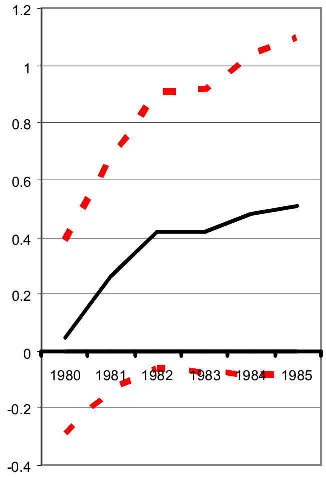


Figure A1.3: The Dinstr coefficient with unadjusted technological growth rate for changes in CU and labor productivity



Shifting rolling regression results

Figure A2.1: The D_govcivil coefficient with unadjusted technological growth rate for changes in CU and labor productivity

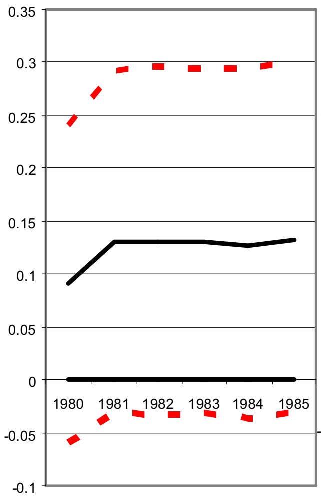


Figure A2.2: The Dtour coefficient with unadjusted technological growth rate for changes in CU and labor productivity

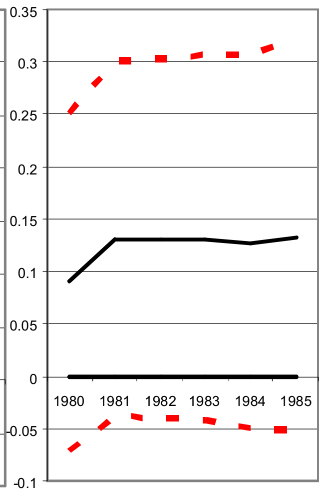
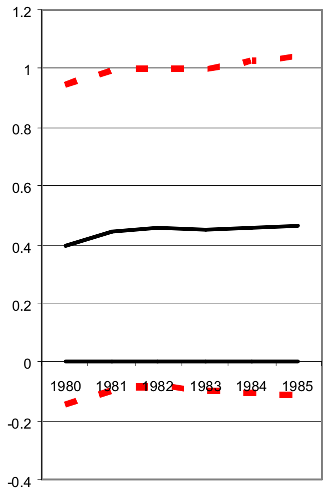


Figure A2.3: The Dinstr coefficient with unadjusted technological growth rate for changes in CU and labor productivity



Contracting rolling regression results

Figure B1.1: The D_govcivil coefficient with adjusted technological growth rate for changes in CU and labor productivity

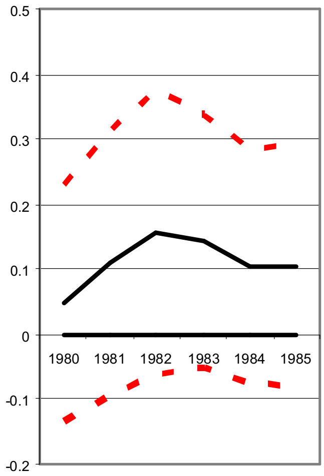


Figure B1.2: The Dtour coefficient with adjusted technological growth rate for changes in CU and labor productivity

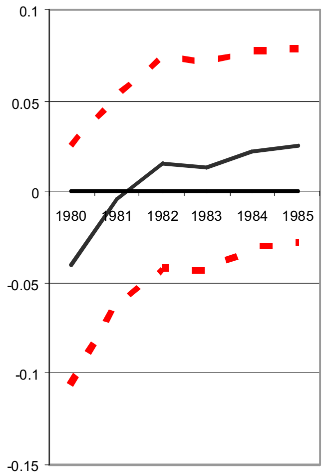
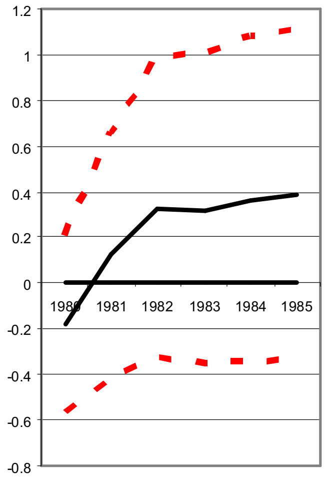


Figure B1.3: Dinstr coefficient with adjusted technological growth rate for changes in CU and labor productivity



Shifting rolling regression results

Figure B2.1: The D_govcivil coefficient with adjusted technological growth rate for changes in CU and labor productivity

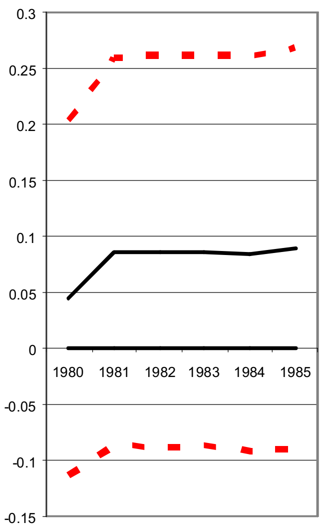


Figure B2.2: The Dtour coefficient with adjusted technological growth rate for changes in CU and labor productivity

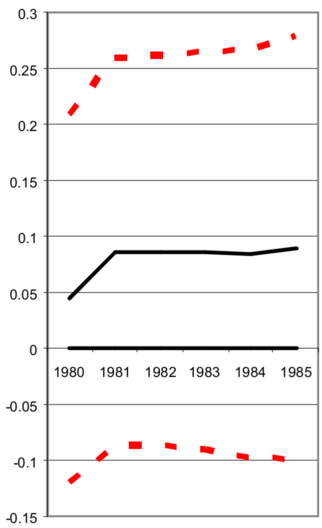
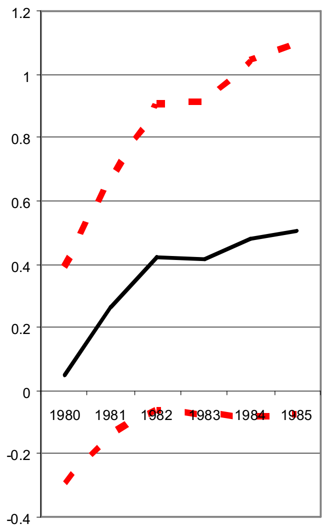


Figure B2.3: The Dinstr coefficient with adjusted technological growth rate for changes in CU and labor productivity



REFERENCES

Basu, Susanto, and Fernald, John H. (1997). "Returns to Scale in U.S Production: Estimates and Implications," *Journal of Political Economy* 105, 249-283.

Basu, Susanto, and Fernald, John H. (2001). "Why is Productivity Procyclical? Why Do We Care?" in *New Developments in Productivity Analysis, Hulten, Charles R. (ed)*, 225-296, Chicago: The University of Chicago Press for NBER.

Basu, Susanto, and Kimball, M.S.(1997). "Cyclical Productivity with Unobserved Input Variation" NBER Working Paper no. 5915.

Basu, Susanto, Fernald, John, and Shapiro Matthew D. (2001). "Productivity Growth in the 1990s: Technology Utilization or Adjustment?" *Carnegy-Rochester Conference Series on Public Policy 55*, 117-165.

Burnside, Craig, Eichenbaum, Martin, and Rebelo Sergio T. (1993). "Labor Hoarding and the Business Cycle," *Journal of Political Economy* 101, 245-272.

Burnside, Craig, Eichenbaum, Martin, and Rebelo Sergio T. (1995). "Capital Utilization and Returns to Scale,"in *NNBER Macroeconomics Annual 1995, Bernanke, Ben S and Rotemberg, Julio J. (eds)*, 67-110, Cambridge and London: M.I.T Press, 1995.

Burnside, Craig, and Eichenbaum, Martin (1996). "Factor Hoarding and the Propagation of Business Cycle Shocks," *American Economic Review 86*, 1154-1174.

Caballero, Ricardo J, and Lyons, Richard K. (1992). "External Effects in U.S Procyclical Productivity," *Journal of Monetary Economics* 29, 209-226.

Djivre, Joseph, and Menashe, Yigal (2009). "Testing for Constant Returns to Scale and Perfect Competition in the Israeli Economy, 1980-2006," Bank of Israel, Discussion Paper no 2009.07

Hall, Robert E. (1988). "The Relation between Price and Marginal Cost in U.S. Industry," *Journal of Political Economy* 96, 921- 9.

Hall, Robert E. (1991). "Invariance Propertiew of Solow's Productivity Residual," in *Growth, Productivity Unemployement,* Peter Diamond (ed.), pp. 71-112, Cambridge: The MIT Press.

Hall, Robert E. (2004). "Measuring Factor Adjustment Costs," *The Quarterly Journal of Economics,* August 2004, 119(3), 899-927.

Inklaar, Robert E. (2006). "Cyclical Productivity in Europe and the United States, Evaluating the Evidence on Returns to Scale and Input Utilization," *CEPR Discussion Paper 5501.*

Marom , Arie, and Bregman, Arie (1998). "The Productivity and its Sources in the Israeli Manufacturing Sector, "Bank of Israel, Discussion Paper no 1998.03

Marom, Arie, Menashe, Yigal, and Suchoy Tanya (2003). "The State of the Economy Index and the Probability of Recession: the Markov Regime-Switching Model," Bank of Israel, Discussion Paper no 2003.05

Melnick, Rafi (2002). "Business cycles in Israel," The Economic Quarterly, 219-244 (Hebrew).

Paquet, Alain, and Robidoux, Benoit (2001). "Issues on the, measurement of the Solow Residual and the Testing of its Exogeneity: Evidence for Canada," *Journal of Monetary Economics* 47, 595-612.

Roeger, W. (1995). "Can Imperfect Competition Explain the Difference between Primal and Dual Productivity Measures? Estimates for U.S Manufacturing," *Journal of Political Economy,*103(2),316-330.

Vecchi, Michela (2000). "Increasing Returns, Labor Utilization and Externalities: Procyclical Productivity in the United States and Japan," *Economica, 67,*229-244.

Zussman, Noam, and Friedman Amit (2008). "The Quality of the Labor Force in Israel," Bank of Israel, Discussion Paper no 2008.01.