Research Department

Bank of Israel

Testing for Constant Returns to Scale and Perfect Competition in the Israeli Economy ,1980–2006*

Joseph Djivre and Yigal Menashe

Discussion Paper No. 2009.07
July 2009

חטיבת המחקר, בנק ישראל ת״ד 780 ירושלים 91007
Research Department, Bank of Israel, POB 780, 91007 Jerusalem, Israel

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. 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, they contradict the assumption of CRS with market power and fixed entry costs in the modeling of the Israeli economy.

Key words: returns to scale, market power, TFP, Solow residual, capacity utilization

JEL: E22, E32, F12, L16

האם הסקטור העסקי בישראל פועל תחת תחרות משוכללת ותשואה קבועה לגודל – בחינה אמפירית לשנים 2006-1980

יוסי גיברה ויגאל מנשה

תקציר

במאמר זה אנו בוחנים את קיומה של ההשערה המשותפת של תשואה קבועה לגודל והעדר כוח שוק על הסקטור העסקי בישראל, תוך יישום מבחן א-פרמטרי שהציע HALL במאמריו משנים 1988 ו-1991. התוצאות האימפיריות המבוססות על נתוני התוצר, תשומות ההון והעבודה (כמות ואיכות) וניצולות גורמי הייצור בתדירות שנתית מאוששות את קיומה של תחרות משוכללת ברמה המצרפית. ניתן לומר כי התוצאות בעבודה תומכות מצד אחד בכפיית ההנחה של תשואה קבועה לגודל באמידת פונקצית הייצור של הסקטור העסקי בישראל ומצד שני מעמידות בספק את ההנחה שבסקטור העסקי בישראל קיים מצב תשואה קבועה לגודל יחד עם כוח שוק מונופוליסטי ועליות כניסה קבועות .

1. Introduction

In this paper we test the hypothesis of constant returns to scale and no market power in the goods market, implementing Hall's test (1988, 1991) on annual real business sector 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 that do not originate in or affect technological changes.

The issue of divergence from the joint hypothesis of constant returns to scale (CRS) and no market power in the goods market (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 the accounting 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 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 relating to the period 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 econometrically the extent of IRS in the U.S. economy. The results obtained justified the attribution of the procyclicality of SR and of 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 would give rise to a procyclical residual because output would rise (fall) by more than the rise (fall) implied by the increase (decrease) of the production inputs in the presence of IRS, thus giving the false impression of output enhancing productivity gains.

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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 they 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) and Basu and Fernald (2001), respectively, found evidence of limited economies of scale in the production of durables and 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, thereby 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 accelerated investment 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 acceleration in investment. These costs created a wedge between the higher actual output level and its lower measured counterpart which does not include adjustment costs. [1]

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 of economies 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

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

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.

Given that the intensity of input utilization is unobservable, it was expressed in the empirical literature as a function of observed variables derived from the F.O.C. for efficient production in the estimation of the production function which were introduced in its stead in the econometric estimation of the production function.

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, and their relative price, allows us to express capital services as an increasing function of the labor input 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, average labor hours and unknown coefficients, obtained from the F.O.C 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 from 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, the capital stock, production materials, and the prices of production materials and of new capital (investment goods). They found however that the adjustment of the 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. 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 model 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 to the parametric approach only when such reference is necessary to clarify our exposition. This paper contains four additional parts: In the second part we present Hall's methodology for 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 we conclude.

2. Hall's Test

In this section we describe the methodology of Hall's test and we subsequently present the economic intuition and a more formal explanation behind the perceived procyclicality of changes in TFP under the null hypothesis. The rationale behind the test is that, under the erroneous assumption that the economy operates under CRS and no market power, while in fact it does not, the derivation of the change in the 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. Such a correlation will therefore constitute a proof of the divergence of production from the CRS and 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) $\dfrac{dY}{Y} = \dfrac{dF}{F} + \dfrac{dA}{A} = \dfrac{dY}{dK} \cdot \dfrac{K}{Y} \cdot \dfrac{dK}{K} + \dfrac{dY}{dL} \cdot \dfrac{L}{Y} \cdot \dfrac{dL}{L} + \dfrac{dA}{A}$.

Our assumption of Hicks neutral technological innovations allows us to rewrite expression (2) in terms of F instead of in terms of the domestic product, Y:

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

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:

(3) $\dfrac{dY}{dK} \cdot \dfrac{K}{Y} + \dfrac{dY}{dL} \cdot \dfrac{L}{Y} = \dfrac{dF}{dK} \cdot \dfrac{K}{F} + \dfrac{dF}{dL} \cdot \dfrac{L}{F} = \gamma$ or $\dfrac{dF}{dK} \cdot \dfrac{K}{F} = \gamma - \dfrac{dF}{dL} \cdot \dfrac{L}{F}$.

The assumption that price, P, is set as a gross mark-up, μ, over marginal cost, MC, allows us to rewrite expression (3), under the assumption that firms are price takers in the inputs market, as follows:

$$(4) \quad \gamma = \mu \cdot \frac{W \cdot L}{Y \cdot P_y} + \mu \cdot \frac{P_{cr} \cdot K}{Y \cdot P_y} \quad .$$

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[2]

$$(5) \quad \gamma = \mu \cdot \alpha + \mu \cdot (1 - \alpha - \pi_R) = \mu \cdot (1 - \pi_R) \quad .$$

It transpires from expression (5) that the gross mark-up, 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 ($\gamma=1$) implies zero monopolistic rents, $\pi_R = 0$, i.e., free entry and thereby free competition. It also transpires from expression (4) that the elasticities of domestic product with respect to the labor and capital services are equal to α *and (1-α)* 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. [3]

Inserting these elasticities in (2') and rearranging terms we obtain that the rate of technological change is given by (6) under the null hypothesis and by (6') and (6"-6"') respectively under CRS with market power and under IRS:

$$(6) \quad \frac{dA}{A} = \left[\frac{dY}{Y} - \frac{dK}{K} \right] - a \cdot (\frac{dL}{L} - \frac{dK}{K}) \quad ;$$

$$(6') \quad \frac{dA}{A} = \left[\frac{dY}{Y} - \frac{dK}{K} \right] - \mu \cdot a \cdot (\frac{dL}{L} - \frac{dK}{K}) = \left[\frac{dY}{Y} - \frac{dK}{K} \right] - a \cdot (\frac{dL}{L} - \frac{dK}{K}) - (\mu - 1) \cdot a \cdot (\frac{dL}{L} - \frac{dK}{K}) \quad ;$$

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[2] 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.
[3] In all of the above and subsequent expressions the gross markup measuring market power, the labor share in revenue, the rate of growth of domestic product, of production inputs and of technology all have time subscripts which have been suppressed to simplify the exposition.

$$(6'') \quad \frac{dA}{A} = \left[\frac{dY}{Y} - \frac{dK}{K}\right] - (\gamma - 1) \cdot \frac{dK}{K} - \mu \cdot a \cdot (\frac{dL}{L} - \frac{dK}{K}) \; ; \text{ or}$$

$$(6''') \quad \frac{dA}{A} = \left[\frac{dY}{Y} - \frac{dK}{K}\right] - a \cdot \left[\frac{dL}{L} - \frac{dK}{K}\right] - (\mu - 1) \cdot a \cdot (\frac{dL}{L} - \frac{dK}{K}) - (\gamma - 1) \cdot \frac{dK}{K} \; .$$

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, $\frac{dY}{Y} - \frac{dK}{K}$, 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. Expressions (6)-(6''') imply that if technological change is erroneously derived on the basis of the RHS of expression (6) when in fact the economy is not characterized by CRS and 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 will in this case be equal to

$$(7) \quad \frac{dA}{A} + \left[(\mu - 1) \cdot a \cdot (\frac{dL}{L} - \frac{dK}{K})\right]$$

if the economy is characterized by CRS and market power, and to

$$(8) \quad \frac{dA}{A} + \left[(\mu - 1) \cdot a \cdot \left(\frac{dL}{L} - \frac{dK}{K}\right) + (\gamma - 1) \cdot \frac{dK}{K}\right]$$

in the presence of IRS.

If the rate of change of the instruments is positively correlated with the rate of GDP growth and uncorrelated by selection with the rate of technological change, $\frac{dA}{A}$, then it must also be positively correlated with the factors in brackets in expressions (7) and (8) according to expressions (6'-6'''). If the null hypothesis should be rejected because of market power ($\mu > 1$) or because production is characterized by IRS($\gamma > 1$), then the measured rate of technological change under the

erroneous null hypothesis, obtained as an SR, will also include these factors and will be thereby positively correlated with the rate of change of the exogenous instruments.[4]

At the empirical level it is possible to obtain that the rate of capital growth is determined on the basis of decisions made in previous periods, so that its correlation with the instrument may be nil. The impression may thus be created that the null hypothesis may be falsely accepted in spite of the existence of IRS. However the existence of such economies is consistent with economic profits and as a result it requires a gross mark-up which is greater than unity according to expression (5) or, technically, a value for γ greater than unity requires the existence of market power ($\mu > 1$). As a result a positive correlation between changes in the instrument and the growth rate of labor employment per unit of capital will be sufficient to detect the afore mentioned positive correlation, necessary for the rejection of the null hypothesis under IRS even if the rate of growth of capital is uncorrelated with the instrument.[5]

In order to determine the existence of a correlation between the rate of change of the instrument expressed here as dz, and the rate of technological change obtained under the null hypothesis, Hall (1988) ran a regression of the RHS of (6) on dz:[6]

$$(9) \quad \frac{dY}{Y} - \frac{dK}{K} - a \cdot (\frac{dL}{L} - \frac{dK}{K}) = c_1 + c_2 \cdot dz \ .$$

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 product/capital ratio constitutes a proof of the statistically significant correlation required for the rejection of the null hypothesis. The correlation between changes in the instrument and the rate of technological growth under the null hypothesis is affected by mismeasurement of output and of the production inputs leading to its false rejection or

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[4]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.

[5] The performance of Hall's test on the basis of domestic product shares does not allow us to differentiate between the two alternative hypotheses of CRS with market power and IRS with market power. A similar test can be performed calculating a cost-based residual derived from cost-based shares. Such a test can detect the departure from CRS only, because production costs cannot account for the existence of markups, being expressed at factor cost and not in market prices. For this reason such a test can serve to check whether findings of a divergence from the assumptions of CRS and of a no-market-power paradigm are due to a departure from the first or from the second assumption. Unfortunately there are no data available which allow the performance of this test for the Israeli economy at the aggregate level.

[6] 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.

acceptance. Following Hall we define below the conditions under which such outcomes can be obtained, and discuss the ways to avoid them.

b. The false rejection or acceptance of the null hypothesis.

The mismeasurement of production inputs leading to the false rejection of the null hypothesis may arise from the mismeasurement of the intensity of their utilization, of their quality, or of overhead costs in production. Output mismeasurement, in its turn, may be the result of investment adjustment costs being overlooked as a result of which the correlation between changes in the instruments and the measured technological rate of growth is mitigated. In this way input mismeasurement can lead to the false rejection of the null hypothesis, while output mismeasurement may lead to its false acceptance.

In the following section we present the issue of the input and output mismeasurement in some detail. We also show that both the non-parametric methodology we use and the parametric methodology lead to similar results with respect to the false acceptance or rejection of the null hypothesis with the exception of the existence of fixed overhead labor costs. While in this case the application of Hall's test will not lead to the false rejection of the null hypothesis, when overhead costs are overlooked the application of the parametric approach will bias the estimation results in favor of the existence of IRS.

b.1 The mismeasurement of output

The mismeasurement of output arises from the existence of investment adjustment costs, among other things, as a result of which services rendered within the firm by its labor force may not be considered as an output, while the corresponding labor input in the provision of these services is considered as part of the labor force. 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 it is possible that the measured domestic product will fall short from its actual level and from 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. In this way the procyclicality of the changes in the SR when the null hypothesis is false could be obscured by the countercyclical effect on TFP changes of accelerated investment during expansions.

In a more formal manner we can write the observable output Y_{OBS} as a function of the actual output, Y, and of adjustment costs, f, which are an increasing function of the investment capital stock ratio:

(10) $Y_{OBS} = Y \cdot [1 - f(\frac{I}{K})]$, with f'>0 .

The function f provides a measure of the weight of unmeasured adjustment costs in the actual domestic product. Taking logs and differentiating with respect to time we obtain expression (11) below:

(11) $\dfrac{dY_{OBS}}{Y_{OBS}} = \dfrac{dY}{Y} - \dfrac{df}{dt} \, or$.

Under the assumption of IRS the above expression can be rewritten as follows:

(11') $\dfrac{dY_{OBS}}{Y_{OBS}} - \dfrac{dK}{K} - a \cdot \left[\dfrac{dL}{L} - \dfrac{dK}{K} \right] = \left[(\mu - 1) \cdot a \cdot (\dfrac{dL}{L} - \dfrac{dK}{K}) + \dfrac{dA}{A} + (\gamma - 1) \cdot \dfrac{dK}{K} \right] - \dfrac{df}{dt}$.

Had output been properly measured, the procyclical adjustment cost effect, $\dfrac{df}{dt}$, would have appeared on the LHS of expression (11') with a positive sign, and technological growth, obtained under the erroneous null hypothesis, would have been equal to the term in brackets on the RHS of this expression, which is known to be procyclical and hence positively correlated with changes in the so-chosen procyclical instrument. This positive correlation would have allowed the rejection of the null hypothesis when the alternative hypothesis of IRS is correct. As it is, the adjustment cost effect, which is procyclical, appears on the RHS of (11) with a negative sign and thereby mitigates the aforementioned positive correlation. In this way the null hypothesis may be falsely accepted in spite of the existence of IRS.

A similar result of underestimation of economies of scale is obtained under the parametric approach of the production-function econometric estimation. 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, $-\dfrac{df}{dt}$, is negative, given that $\dfrac{df}{dt}$ is in its turn positive and procyclical. This negative bias gives rise to underestimates for the output elasticities with respect to the production inputs, supporting the false acceptance of the null hypothesis.

b.2 The mismeasurement of the intensity of utilization of the production inputs and of labor productivity

The positive correlation between technological growth derived as an SR under the null hypothesis and changes in the instrument may also arise from the mismeasurement of production factor utilization. If a lower 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 a higher level of output than the one actually produced. Under these circumstances 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 in TFP growth under the null hypothesis, on the other. Such a result may lead to the false rejection of the null hypothesis. The above argument may be presented in a more formal way as follows:

Let the production function be given by

(12) Y= F(K.cu,L.e) .

Then, expression (6) above may be re-written as

$$(13) \quad \frac{dY}{Y} - \frac{dK}{K} - a \cdot \left(\frac{dL}{L} - \frac{dK}{K} \right) = \frac{dA}{A} + a \cdot \frac{de}{e} + (1-a) \cdot \frac{dcu}{cu} \; .$$

As a result of overlooking the intensity of utilization of production inputs, changes thereof are erroneously included in the measured technological growth obtained as an SR. If these changes in the intensity of utilization of the production inputs are procyclical they will be also positively correlated

with changes in the procyclical instrument, giving rise to a positive correlation between the latter and changes in the measured TFP, thus leading to the false rejection of the null hypothesis.

A similar outcome is obtained when implementing the parametric methodology of production function estimation. The omission of the change in the intensity of utilization of the two inputs creates a positive bias in the estimation of the coefficients of the growth rates of the labor and capital inputs. The positive sign of the bias is due to the positive correlation between the omitted procyclical regressors, which measure the intensity of input utilization, and the existing regressors, that is the rates of growth of employment and capital stock. 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.

If we relate to $\dfrac{de}{e}$ as a measure of the rate of change of labor productivity, then by the same token overlooking it will give rise to a positive correlation between the SR and changes in the instrument if this change in labor productivity is procyclical, leading to the false rejection of the null hypothesis.

However if changes in labor productivity are negatively correlated with output growth, then they will most probably be negatively correlated also with changes in the instrument. This negative correlation will mitigate the positive correlation between changes in the procyclical instrument and the SR when the null hypothesis is false. This mitigation effect may in its turn allow the false acceptance of the null in spite of the existence of IRS.

b.4 Labor overhead costs

If workers allocate a certain number of hours to non-productive activities at work regardless of the length of their shift, then longer shifts, arising from demand-induced upturns in economic activity, are consistent with an increase in the effective productive hours of work. This increase is relatively more substantial than the corresponding increase in the reported working hours, which include also hours spent on non-productive activities. As a result longer shifts will induce a greater output expansion than that implied by the growth rate of reported working hours, since output expansion varies

according to the higher growth rate of the effective working hours and not to the lower growth rate of reported working hours. Therefore, if we use the lower values of the latter as a regressor in the econometric estimation of the production function instead of the higher values of the former, then under the null hypothesis we will obtain an unexplained increase in output. In order to account for this higher output, the estimation results will provide an overestimate for the output elasticity with respect to the labor input, biasing the estimation result in favor of the IRS hypothesis and of the rejection of the null. This intuition is not however valid when implementing Hall's test, as is shown below.

Let the production function be expressed by

(14) $Y_t = A_t.F[L_t \cdot (h - \phi), K_t]$.

The above production function is similar to (1) but it differs from it in that whereas total working hours are equal to h, productive hours are equal to $(h - \varphi)$, a fixed number of hours φ being spent on non-directly-productive activities. For reasons of simplicity we assume here that the employment level, L, in terms of the number of workers is fixed, so that the rate of change of the productive labor input is measured by $\dfrac{dh}{(h - \varphi)}$. Under the null hypothesis the rate of product growth is given by

(15) $\dfrac{dY}{Y} = \left[\dfrac{d\mathbf{F}}{dK} \cdot \dfrac{K}{F} \right] \cdot \dfrac{dK}{K} + \left[\dfrac{d\mathbf{F}}{dh} \cdot \dfrac{h - \varphi}{F} \right] \cdot \dfrac{dh}{h - \varphi} + \dfrac{dA}{A}$.

Given that φ is constant, the second term in the RHS of expression (15) measures, under the null hypothesis, the contribution to output growth of the growth rate of effective working hours. This term may be rewritten on the basis of expression (3) above as follows:

(16) $\left[\dfrac{d\mathbf{F}}{dh} \cdot \dfrac{h - \varphi}{F} \right] \cdot \dfrac{dh}{h - \varphi} = \left[\dfrac{A \cdot d\mathbf{F}}{dh} \cdot \dfrac{h - \varphi}{A \cdot F} \right] \cdot \dfrac{dh}{h - \varphi} = \left[\dfrac{w}{P} \cdot \dfrac{h - \varphi}{A \cdot F} \right] \cdot \dfrac{dh}{h - \varphi} = \left[\dfrac{w}{P} \cdot \dfrac{h}{Y} \right] \cdot \dfrac{dh}{h}$.

Expression (16) implies that the contribution to output growth of the growth rate of effective working hours, $\dfrac{dh}{h - \varphi}$, and of reported working hours, $\dfrac{dh}{h}$, are equal provided they are weighted by their corresponding shares in the value of output. As a result, no bias is introduced in the calculation of the

change in TFP under the null hypothesis when implementing Hall's test, if we do not differentiate between effective and reported working hours.

3. The Test and the Data

In this section we present the data and the variables relevant to the implementation of Hall's test.

We have assumed that production takes place in two stages so that the output of the business sector is produced using intermediate goods and value added, which in 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 SIC classification. Our sample period included 28 annual observations between 1979 and 2006 for the test 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 procyclical, 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 missing also 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 for the output function.[7]

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[7] 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 this factor, whose direction constitutes an empirical issue [Basu and Fernald (1997)]. Paquet and Robidoux (2001) have used directly aggregated GDP data on the Canadian economy without resorting to the aggregation of sectoral data.

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

We based our hypothesis testing, like Hall (1988) on a regression estimation in which the 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. As will be shown below, the test results did not allow the rejection of the null hypothesis especially after correcting for changes in the productivity of labor and in the capacity utilization of capital.

a. Measuring GDP and the labor share

The derivation of the rate of technological change as an SR under the null hypothesis based on the calculation of the RHS of expression (5) requires the calculation of the rates of growth of GDP, production inputs and labor share. In addition to these variables, the rate of change of the production inputs has to be adjusted for changes in their intensity of utilization and for changes in the productivity of labor.

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. As a result, we were constrained to expressing GDP growth in terms of constant market prices since our sample starts from 1979.[8] 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 expressed 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.

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[8] The correlation coefficient between the GDP growth at fixed producer 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 may not have impaired the accuracy of our calculations.

In Israel investment statistics include, according to the CBS data, a large part of the costs incurred by 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 noted 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 accepting the null hypothesis because of the procyclical underestimation of the TFP growth to be limited.

According to our exposition we should not be preoccupied with the issue of overhead costs which affect the effective labor input when implementing Hall's test. With respect to the measurement of the effective production inputs, we were thus left to cope with their qualitative aspects and with their intensity of utilization. Our use of capital in 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.

b. Measuring the effective production inputs

b.1 Accounting for changes in labor productivity

A substantial part of our sample period was affected by the absorption of an unprecedented influx of immigrants at 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 adversely affected the productivity of labor [Friedman and Zussman (2008)], during an upswing in economic activity, ignoring 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, favoring the acceptance of the null hypothesis according to our previous analysis. We therefore adjusted the labor input data to account for the 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 productivity. If the wage differentials identified by Friedman and Zussman do not reflect productivity differentials but discrimination against new immigrants, then, as we show below, by adjusting the labor input for a fall in productivity which did not occur, we would have inflated the productivity gains by a factor which is

procyclical in our sample period and is thus expected to be positively correlated with changes in our instrument. We would have thus created a bias 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.

If the quality of the labor input has indeed been deteriorating at a rate of $\frac{de}{e} < 0$, the true estimate of the rate of technological change under the null hypothesis will be equal to expression (17) below instead of being equal to the RHS of (6).

$$(17) \quad \frac{dY}{Y} - \frac{dK}{K} - a \cdot (\frac{dL}{L} + \frac{de}{e} - \frac{dK}{K}) = \frac{dA}{A} \ .$$

If on the other hand the fall in labor productivity is fictitious, the true rate of technological change is given by the RHS of expression (6). If under these circumstances we erroneously derive the rate of technological change as being equal to the LHS of (17), while (6) is correct, we will obtain a biased rate of technological change, which will be equal to $\frac{dA}{A} - a\frac{de}{e}$. The bias is positive, because of the assumed fall in productivity, and equal to $-a\frac{de}{e}$. It is also procyclical, since the immigration influx induced an upswing in activity and should thus be positively correlated with changes in our instrument, leading to the false rejection of the null hypothesis.

b.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.

Choosing the Proxy Variable

Capacity utilization is in general considered to be procyclical. The economic intuition behind the procyclicality 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 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 mitigate the extent of the increase in the intensity of input utilization.
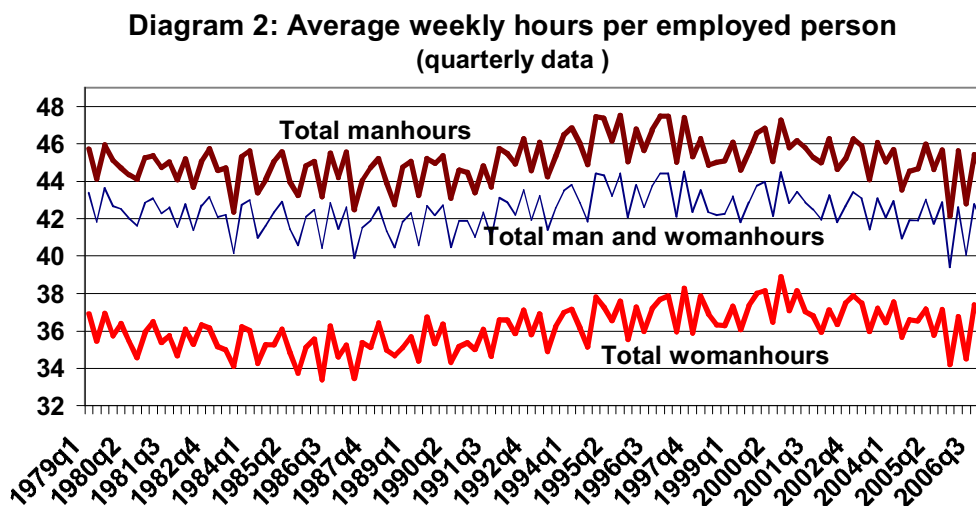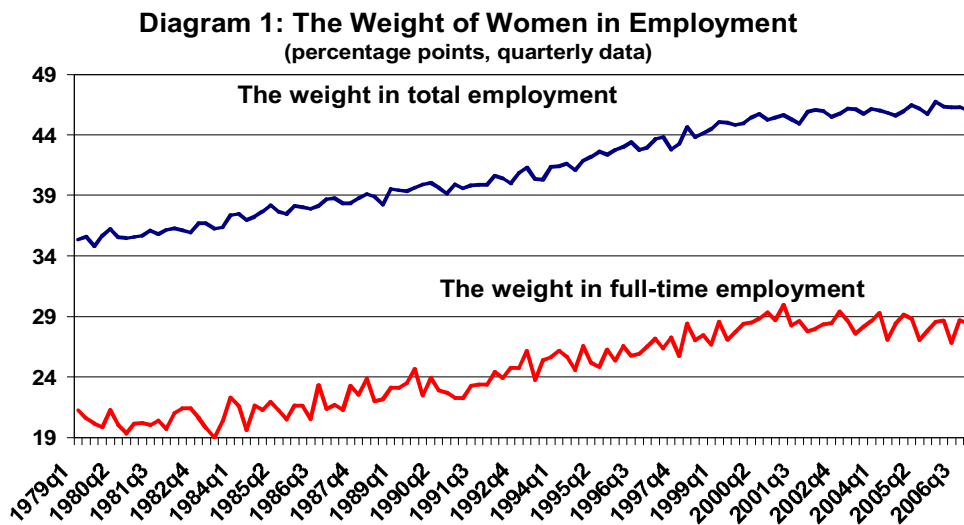
By the same token, according to Basu and Fernald (2001) and Basu et al. (2001), higher investment adjustment costs induce an efficiently producing firm to increase its capacity utilization in the margin when the demand for investment increases. If these costs are an increasing function of the I/K ratio, then we would expect CU to be positively correlated with the latter. As a result, CU would be expected to be higher the greater the deviation of the capital stock from the level satisfying the firm's optimality conditions, since the I/K ratio should be an increasing function of the capital stock deviation from its optimal level. This conclusion is in line with the spirit of the analytical results obtained by Greenwood et al. (1988).
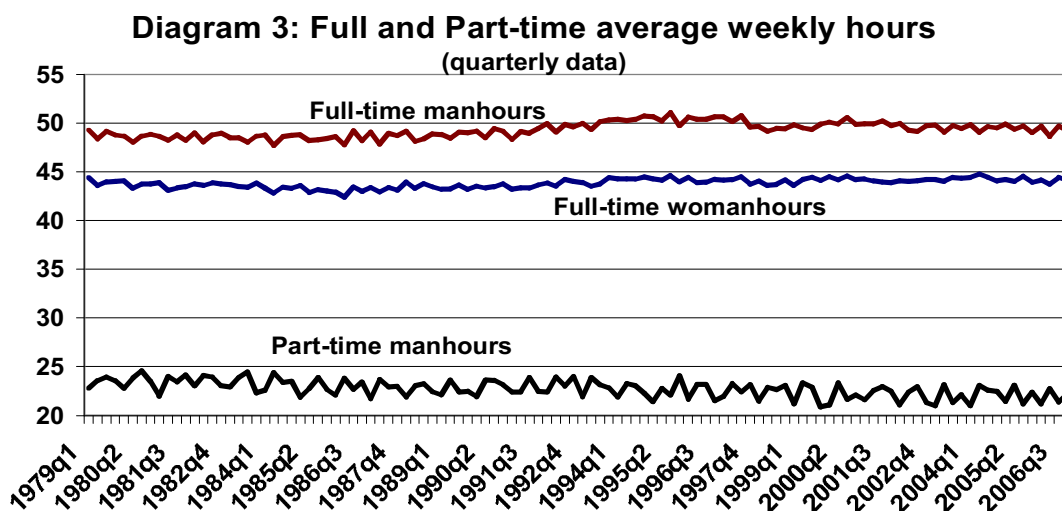
These efficiency conditions imply that CU should be procyclical and positively correlated with the I/K ratio and as a result our proxy variable for it should be also characterized by similar properties. Our choice of this proxy variable is based on the fact that efficient production requires marginal costs to be equal across all production inputs. As a result 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 capital because of shift premia. 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.[9]

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[9] 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 working weekly 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 increasing or decreasing trends resulting from structural changes in the Israeli economy.

The rising share of women in employment over time (Diagram 1) alongside their lower average weekly hours than men's (Diagram 2), as a result of their consistently lower level of full-time working hours (Diagram 3), are expected to have had a negative effect over time on the aggregate average weekly hours.

Diagram 1: The Weight of Women in Employment
(percentage points, quarterly data)

Diagram 2: Average weekly hours per employed person
(quarterly data)

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Diagram 3: Full and Part-time average weekly hours
**(quarterly data)**

This implies that our use of average working hours as a proxy for capital capacity utilization should not 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, after that date it increased, and it stabilized after 2002 at a slightly higher level of 49 weekly hours (Diagram 4).

Diagram 4: Full-time average weekly manhours
(quarterly data)
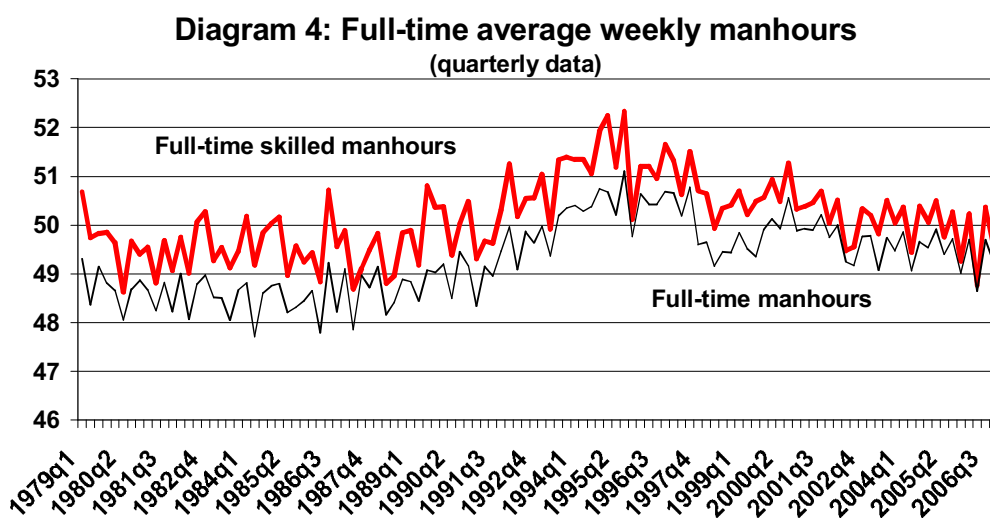
A closer analysis of the data reveals however that this slight rise in the average weekly hours of full-time male employees reflects, among other things, the increasing weight of skilled male workers over time (Diagram 5) and the fact that in general they work longer hours (Diagram 4). The slight rise in the length of weekly hours worked by unskilled male employees after 2002 compared to the pre-1989 period (Diagram 5) also contributed to the rise in the average weekly working hours of male workers.

Diagram 5: Average weekly hours of full-time unskilled workers and the weight of skilled males in full-time employment
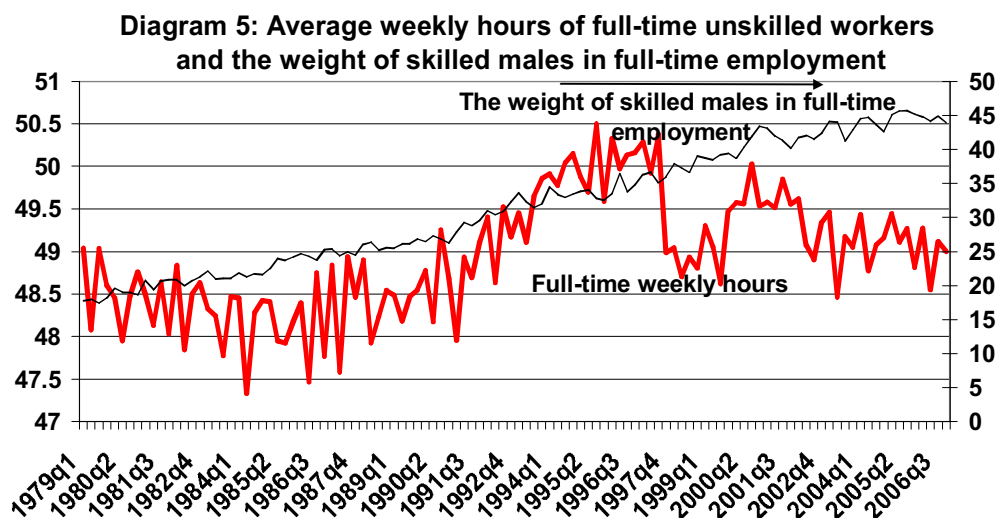
To neutralize this composition effect we chose as a proxy for capacity utilization the average weekly hours of full-time skilled male employees, defined as employed persons with 13 years or more of formal education. [10] This variable converged to its pre-1989 level after 2002–03 (Diagram 4). When measuring 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 need not necessarily 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 to derive our proxy.

The cyclical characteristics of the proxy variable

In this section we examined the extent to which our choice of the proxy variable chosen has properties compatible with those of capacity utilization (CU), and whether its evolution over time resembles that of indices of CU obtained by different methodologies in other countries.

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[10] The length of the working week in hours 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 sector 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 the full-time skilled man-hours in manufacturing and in the financial sector was not 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 the number of full-time employed skilled male employees.

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 de-trended component is for it to serve as a basis for the construction of a CU index around the mean which will reflect full capacity. The cycles obtained 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 (Table 1) but not by any other methodology of cycle classification in the Israeli economy.

Table 1: The cycles in economic activity in Israel and in average weekly working hours[(1]

Melnick Dates*		Marom et al. (1) Dates*		Marom et al. (2)Dates*		CU Dates**	
Recession	1975.5–1980.7	Recession	1979.4–1980.7	Recession	1979–1980.5	Recession	1979–1982.2
Expansion	1980.8–1983.9	Expansion	1980.8–1981.7	Expansion	1980.6–1988.3	Expansion	
Recession		Recession	1981.8–1982.9	Recession		Recession	
Expansion		Expansion	1982.10–1983.8	Expansion		Expansion	1982.3–1987.1
Recession	1983.10–1985.10	Recession	1983.9–1984.3	Recession		Recession	
Expansion	1985.11–1987.8	Expansion	1984.4–1987.5	Expansion		Expansion	
Recession	1987.9–1989.12	Recession	1987.6–1990.2	Recession	1988.4–1989.8	Recession	1987.2–1989.1
Expansion	1990.1–1996.4	Expansion	1990.3–1996.2	Expansion	1989.9–1996.2	Expansion	1989.2–1990.2
Recession		Recession		Recession		Recession	1990.3–1993.2
Expansion		Expansion		Expansion		Expansion	1993.3–1997.2
Recession	1996.5–1999.7	Recession	1996.3–1999.3	Recession	1996.3–1999.1	Recession	1997.3–1999.4
Expansion	1999.8–2000.9	Expansion	1999.4–2000.9	Expansion	1999.2–2000.10	Expansion	2000.1–2001.4
Recession	2000.10–2003.7	Recession	2000.10–2003.10	Recession	2000.11–2003.10	Recession	2002.1–2003.1
Expansion	2003.8–2008.5	Expansion	2003.11–2008.5	Expansion	2003.11–2008.5	Expansion	2003.2--?

1) On the basis of the probability of recession.
2) On the basis of the level of the composite index of economic activity in the Israeli economy.
* Monthly data on the basis of the composite index of economic activity in the Israeli economy.
** Quarterly data on the basis of average weekly working hours of skilled male employees.

More precisely, the cycle classification based on CU gives rise to one only 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.[11] The business sector seasonally adjusted GDP growth data on the basis of which we may identify the business cycles in the Israeli

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[11] 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 one 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 into expansion periods. This fact is particularly disconcerting because these years were characterized by the second oil shock.

economy indicate that the last quarter of 1982 was an outlier characterized by an exceptionally high rate of growth (Diagram 6).

This outlier may have stood, apparently, 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 1984–85 and an expansion after that date and until 1987.

Diagram 6 : The quarterly rate of growth of business sector GDP (Deviations from the mean, sa data)
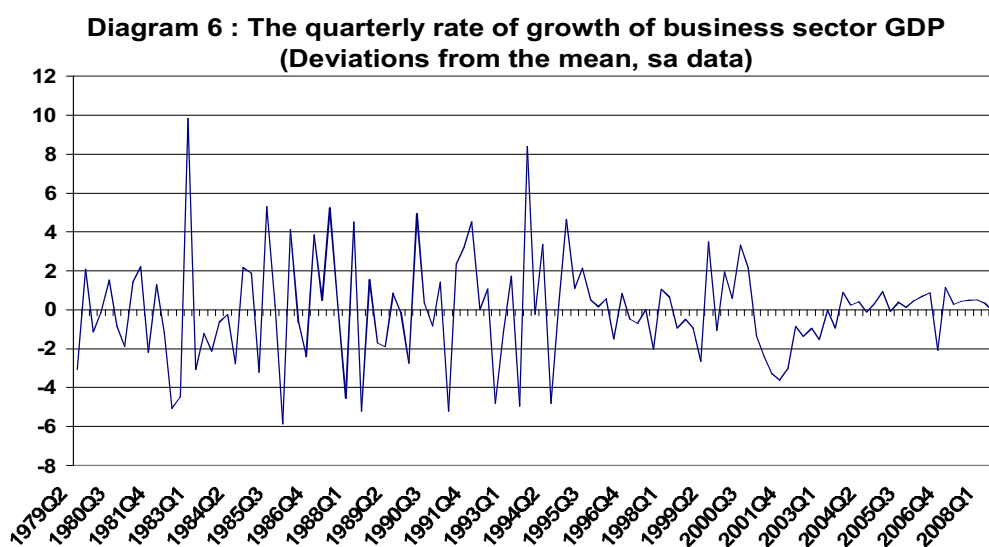
The cycle classification according to the CU index is also consistent with one recession and expansion period between 1979 and 1987 but the end of the recession is anticipated at the second half of 1982 from 1984–85.

The recession which has been identified between 1990 and 1993 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 wave of immigration in the beginning of the 1990s. When the GDP growth rate is de-trended, what is obtained for the years of immigration is a higher level of activity than in normal times, i.e., an expansion. However the inflow of immigrants and their absorption into employment increased overnight not only the actual but also

the potential output and 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 therefore consider 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.

The classification of cycles on the basis of the CU cyclical component succeeds, therefore, in delineating cycles in the Israeli economy between 1979 and 2006 in a way which is similar to that of other classifications. Moreover, in the instances in which the two classifications do not coincide, the difference is either not very substantial or indicates that the CU classification is more precise.

Capacity utilization and the investment capital-stock ratio.

While the cyclical component of 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 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 (Diagram 7) resembling that of the long term evolution of the I/K ratio (Diagram 8).

This observation suggests that our CU index should also include the low frequency component of our proxy variable, implying that the protracted deviation of the average weekly hours and capacity utilization from its steady state may have been the result of the protracted adjustment of the economy to the influx of immigrants in terms of optimal capital stock. Since we perform 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.

Diagram 7: The Low-frequency component of the index of skilled male weekly manhours
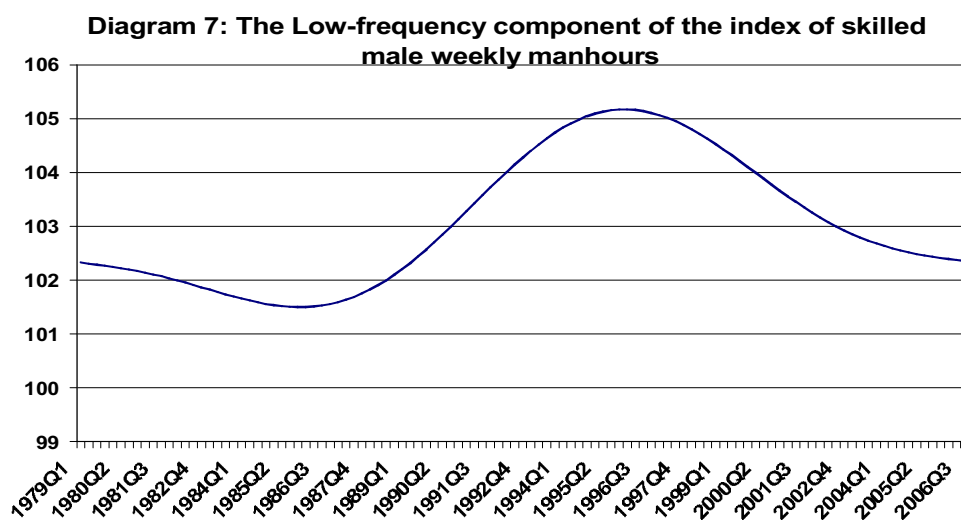
Our choice is consistent with our including in the CU index only the low and business cycle components of the weekly man-hours, even though we use quarterly data which include the high frequency component of weekly working hours as well. This is because averaging the quarterly data filters the high frequency component of our proxy variable.

Diagram 8: The I/K ratio in Israel
(percentage points, quarterly data)
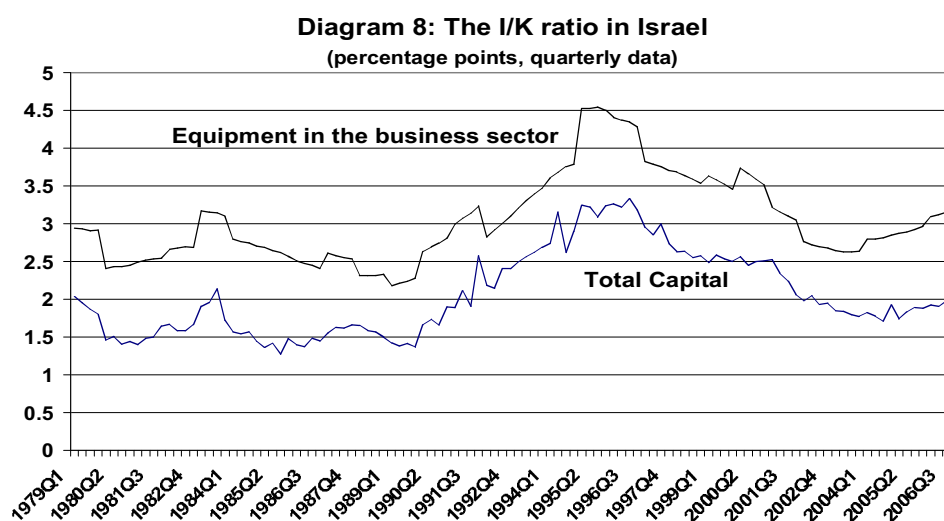
The correlation coefficient between the I/K ratio and the length of the working week of skilled employed males in the business sector is 66 percent for equipment and 70 percent for total investment and capital for the period between 1979:Q1 and 2006:Q4. Between 1995:Q1 and 2006:Q4, the period for which 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.

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This correlation at the aggregate level is reflected in 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 (Diagram 4) while the trend in the I/K data becomes positive around the beginning of 1989 (Diagram 8). This period coincides with the beginning of the end of the downturn after the 1986 stabilization program and the beginning of the immigration influx. Prior to that date the I/K ratio replicates the business cycle 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 and the average skilled man-hours stabilized between 1998 and mid-1999, increased slightly in 2000, renewed its negative trend after 2000, stabilized between 2003 and 2004, and increased 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, and fell after 2000 to its 1989 level. Even though the evolution of I/K replicates that of economic activity during the surveyed period, the upturn observed after 2003 has not been reflected in the length of the working week.

It should be mentioned at this point, however, that the interdependence between the two 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 at the sectorial level fluctuate between 40 percent and 47 percent in manufacturing with respect to investment in equipment and aggregate investment respectively and between 50 percent and 70 percent in services. [12]

Intercountry comparison

Our CU index refers to the business sector as a whole while the available indices from foreign countries refer to capacity utilization in manufacturing only. CU indices for the manufacturing

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[12] It should be noted however that 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.

industry in the Israeli economy did not however substantially change the results, which indicate an amplitude of fluctuations that is rather small for Israel compared to that of CU indices in other economies.

Table 2
Domain of fluctuation in capital capacity utilization in selected developed economies
(percentage points)

Israel	U.S.A	EU	Germany	France	U.K	Small open Eu economies		
						NL	IRL	B
5.4 (2.3)* 6.5** (3.0)*	20.5**	9.6**	15.5**	13.1**	11.6**	7.4**	9.9**	10.3**

  * In parentheses we report the cyclical component data on the basis of spectral analysis.
** 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.

In previous work on the Israeli economy, Bregman and Marom (1998) used electricity consumption per unit of capital in manufacturing as a proxy for capacity utilization in the industry. However its use 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 for 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. Moreover the electricity consumption data of the business sector point to a negative structural break at the end of 1997 both at the aggregate business sector level and in manufacturing, rendering problematic the use of this variable as a proxy for CU.[13]

In spite of the fact that our proxy variable is characterized by properties which are in line with economic theory (similar timing with business cycles measured by fluctuations in economic activity and correlation with the I/K ratio), the range of fluctuations of the CU index derived on the basis of this variable is lower than in the EU area and the small EU economies and much lower than in the U.S.

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[13] Since at optimum capital services are an increasing function of energy consumption and a negative function of its relative price with respect to labor, then the combination of lower consumption in conjunction with the apparently higher relative energy prices after 2003 would tend to imply lower rather than higher capital services during the upturn in economic activity.

c. The Selection of the Instruments.

To choose the best-suited instruments 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. [14] We thereby satisfied the condition that the rate of change of the instruments be correlated with the growth rate of output per unit of capital.

If the null hypothesis is false and we calculate the change in TFP on the basis of expression $\frac{dY}{Y} - \frac{dK}{K} - a \cdot (\frac{dL}{L} - \frac{dK}{K})$, disregarding changes in capital utilization and labor productivity when they have actually occurred, then the measured S.R will be equal to the RHS of expression (17) below, the last component in the RHS of this expression reflecting the effect of IRS when the latter exists. If production is characterized by CRS $(\gamma = 1)$, then the corresponding factor in expression (17) vanishes:

$$(17) \quad \frac{dY}{Y} - \frac{dK}{K} - a \cdot (\frac{dL}{L} - \frac{dK}{K}) = \frac{dA}{A} + a \cdot \frac{de}{e} + (1-a) \cdot \frac{dcu}{cu} + (\mu-1) \cdot a \cdot (\frac{dL}{L} + \frac{de}{e} - \frac{dK}{K} - \frac{dcu}{cu}) + \\ + (\gamma-1) \cdot (\frac{dK}{K} + \frac{dcu}{cu})$$
.

If changes in TFP account for changes in capacity utilization and labor productivity, then the measured rate of technoplogical change obtained as S.R will be equal to expression (18) below:

$$(18) \quad \frac{dY}{Y} - \frac{dK}{K} - \frac{dcu}{cu} - a \cdot (\frac{dL}{L} + \frac{de}{e} - \frac{dK}{K} - \frac{dcu}{cu}) = \frac{dA}{A} (\mu-1) \cdot a \cdot (\frac{dL}{L} + \frac{de}{e} - \frac{dK}{K} - \frac{dcu}{cu}) \\ + (\gamma-1) \cdot (\frac{dK}{K} + \frac{dcu}{cu})$$
.

It is evident from the RHS of the two expressions above that for our test to detect market power, changes in the chosen instrument should be positively correlated ex ante 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 and the rate of change of the instrument. In the absence of a positive correlation between the rate of growth of labor per unit of capital, adjusted for changes in capacity

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

utilization and in labor productivity, it is impossible to consider a statistically significant correlation between changes in the instrument and the measured rate of technological change as emanating from a false null hypothesis, as it is required for its rejection. For this reason we also estimated similar regressions in which the growth rate of labor per unit of capital or of labor only served as dependent variables and the instruments under consideration as regressors.

Moreover, even though the existence of IRS cannot be detected independently from the existence of market power it will support the rejection of the null hypothesis if the rate of change of the instrument is positively correlated with the rate of growth of the capital stock adjusted for changes in capacity utilization. It is also evident from the RHS of expression (17) that if our calculation of the rate of TFP growth does not account for the growth rates of capacity utilization and labor productivity, then a positive correlation between the rate of growth of the instrument and the rate of growth of capacity utilization can lead to the false rejection of the null hypothesis, as noted earlier, while a negative correlation between the former and the growth rate of labor productivity can lead to the false acceptance of the null hypothesis in spite of the presence of market power. To detect the existence of these correlations we estimated regressions in which the rates of growth of capacity utilization and of labor productivity served as dependent variables and the instruments as their regressors.

We chose our instruments from a list of variables whose rate of change is positively correlated with that of the domestic product per unit of capital but which do not affect and are not affected by technological changes. The instrument which passed the test, whose results are reported in Appendix 1, was the rate of growth of tourist-inflow, hereafter *dtour*. The second instrument we used was the rate of growth of government civilian purchases, hereafter *d_govcicil*. The variable of tourist inflow is not a completely exogenous variable because it is affected by both endogenous and exogenous (e.g., geopolitical) factors. For this reason we should also expect this variable to give rise to statistical results which come closer to the rejection of the null hypothesis.

The estimation results on which we based our choice of the instrumental variables indicated that the change in the tourist inflow exhibited positive and very significant coefficients in the regressions of output growth per unit of capital adjusted for changes in capacity utilization and labor

productivity, and of the growth rate of the labor input per unit of capital adjusted for changes in labor productivity and capacity utilization multiplied by the labor share.[15] The rate of growth of *d_govcicil* was found to exhibit highly significant coefficients in the regressions of the rate of growth of capacity utilization, but its coefficients were at the borderline of significance with respect to the rate of growth of business sector GDP growth per unit of capital adjusted and unadjusted for changes in capacity utilization. Based on these results we also created two-stage instruments using the fitted values of regressions in which we used as regressors the variables *dtour* and *d_govcivil*. We calculated two such instruments. The first was obtained by calculating the fitted value of output per unit of capital without adjusting for changes in labor productivity and capacity utilization while the second was the fitted value obtained for the same variable 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 percent, and the similar results, we report below only the results relating to the latter, hereafter *dinstr*. The rate of growth of the instrument was found to have a highly significant coefficient with the rate of GDP growth without adjustments, with the rate of growth of capacity utilization and with the rate of growth of the labor input adjusted for changes in capacity utilization and labor productivity. None of the instruments was found to have a significant coefficient with respect to the rate of growth of labor productivity and the rate of growth of the capital stock.

These results imply that when implementing Hall's test, the rate of growth of tourist inflow and the two-stage instrument are expected to detect the existence of market power in view of their highly significant coefficients in the regressions of the growth of the labor input adjusted for changes in productivity and capacity utilization. On the other hand the significant coefficient of *d_govcivil* with respect to the rate of growth of capacity utilization implies that its implementation in Hall's test will help to detect the existence of IRS even if the rate of growth of this instrument was not found to be correlated with the growth rate of the capital stock. Moreover when the test is performed using the S.R. residual obtained from the business sector GDP growth rate without adjusting it to changes in capacity utilization, the aforementioned significance should create a bias in favor of the false rejection of the null hypothesis according to the RHS of expression (17).

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[15] Highly significant coefficients were also obtained when the growth of labor was not multiplied by the labor share.

In total we used three different instruments: *dtour*, *dgovcivil*, and the two stage instrument whose derivation was described in detail above. These instruments were chosen from a longer list of variables exogenous to economic activity such as the rate of change of the inflow of new immigrants, of an index of terror attacks, of the lagged value of tourist inflow, of oil price inflation and of the volume of world trade. These potential instrumental variables did not give rise to significant results, failing the test of a suitable instrumental variable.

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 then after making the required adjustments in the labor and capital services as in the LHS of expression (18). We did not however adjust the data to account for changes in the labor effort. We report the test results in Table 3 below and in a more detailed manner in Appendix 2. The number of the regressions estimated for each calculation of the technological growth is equal to the number of the instruments.

The test results for the whole sample reported in Table 3 below do not justify the rejection of the null hypothesis regardless of whether we adjust the data to changes in capacity utilization and labor productivity. After adjusting the TFP growth rate for these changes, the test results distance themselves from the rejection domain even more, as expected, and the significance of the instruments' coefficients falls further relative to its level for the unadjusted S.R. growth rate.

Table 3
The significance of the instrument coefficient
(probabilities in percentage points)

Instrument	Non-Adjusted Δ(TFP)	Adjusted* Δ(TFP)
dour	t=1.00, P=16.5	t=0.79, P=22
dgovcivil	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.

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 fixed 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 percent confidence interval. The only instrument for which we obtained a statistically significant positive coefficient was the rate of growth of government civil expenditures. However, the significance of the corresponding coefficient fell below the level consistent with a 5 percent significance level after we adjusted the SR for changes in capacity utilization and productivity. We therefore attribute the rejection of the null hypothesis for the unadjusted SR to the high correlation between the rate of change of this instrument and the rate of change of capacity utilization which, according to our analysis, may lead 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 no-market-power and/or CRS hypotheses, the null hypothesis would have been rejected also adjusting our SR for changes in capacity utilization and labor productivity.[16]

5. Conclusions

We consider our results as providing evidence of the existence of CRS with no market power in the Israeli economy. Because of the occasional proximity of the results to the rejection limit we consider that additional research is required in the direction of the accurate measurement of capital services and effective labor input. Our findings provide support for the existence of CRS and perfect competition at the aggregate GDP production level in Israel, and create some skepticism about the

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[16] 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 however that this variable failed the instrumental variable test, because it was not found to be correlated either with the business sector GDP growth rate 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. While the *dtour* variable is positively correlated with the GDP growth rate of the business sector, over the sample period, it was negatively correlated with the latter during the years 1990 and 1991 because the period of economic expansion coincided with the Gulf War and the tensions leading to it which negatively affected the inflow of tourists. It is possible, therefore, that the null hypothesis was falsely accepted because the positive correlation between the rate of growth of the instrument, under the false null hypothesis, was mitigated by their negative correlation between 1990 and 1991. The introduction of a dummy variable for this period in the SR regression did not however justify this hypothesis.

assumption of monopolistic competition with entry costs among identical firms with CRS and a

constant mark-up of prices over marginal costs imposed on aggregate macroeconomic models of the

Israeli economy.

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Appendix 1

Table A.1.1: The Selection of the Instrumental Variables:
The Regression Coefficient of the Instruments 1979–2006*
(Quarterly Data, Probabilities in percentage points)

The Dependent Variable	The Instrumental Variable		
	dour	**dGovcivil**	**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 parentheses. The probability P is for one tailed tests.

** The significance of *dtour* and *d_govcivil* in the regression serving to construct the two-stage instrument.
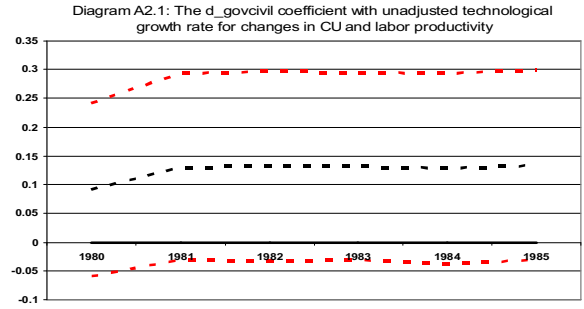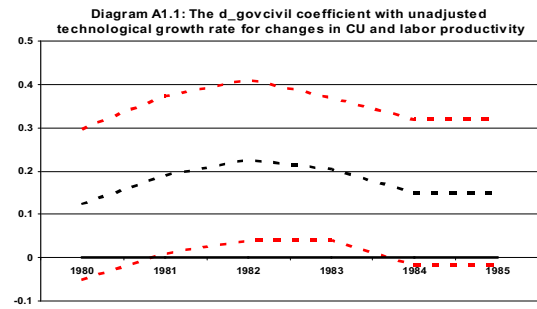

Appendix 2

Definition of the instrumental variables.

d_govcivil: The rate of growth of government civil purchases.

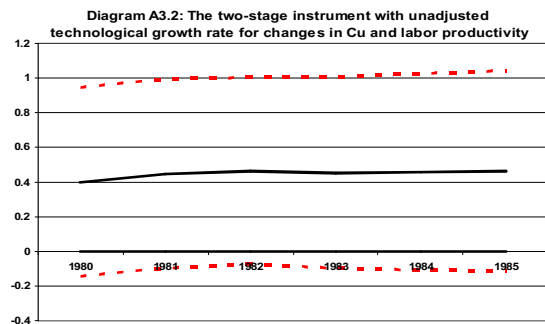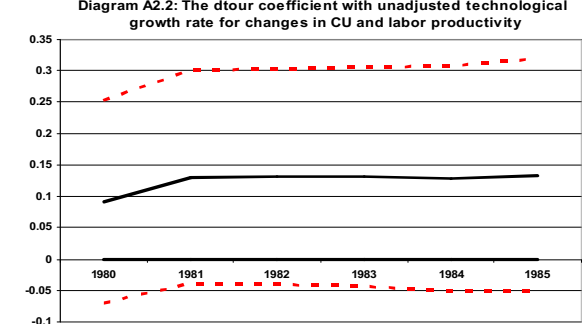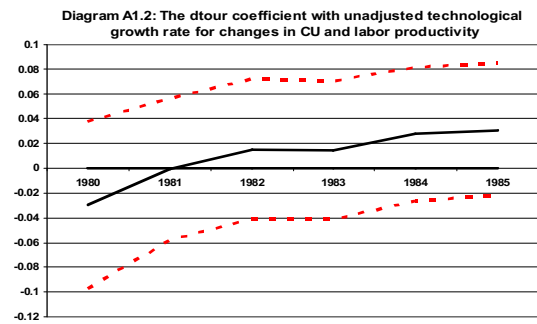**dtour:** The rate of growth of tourist inflow.

Dinstr: tted 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 percent confidence interval around the regression coefficients

Contracting rolling regression results

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

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

Diagram A3.1: The two-stage instrument with unadjusted technological growth rate for changes in Cu and labor productivity

Shifting rolling regression results

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

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

Diagram A3.2: The two-stage instrument with unadjusted technological growth rate for changes in Cu and labor productivity

Contracting rolling regression results

Diagram B1.1: The d_govcivil coefficient with adjusted technological growth rate for changes in CU and labor productivity
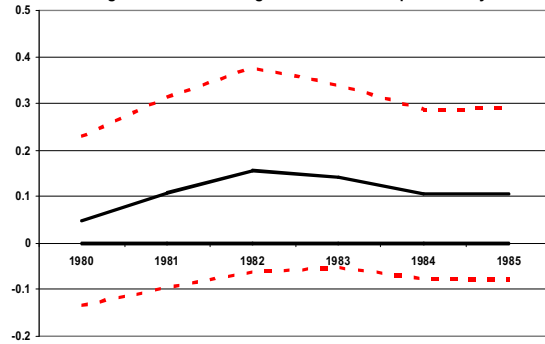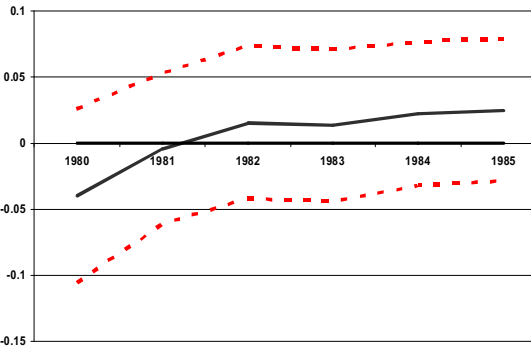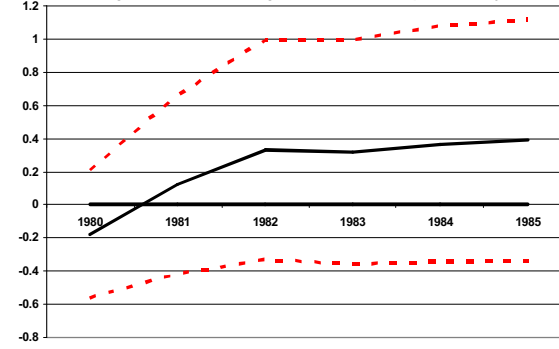
Diagram B1.2: The dtour coefficient with adjusted technological growth rate for changes in CU and labor productivity

Diagram B3.1: The two-stage instrument with adjusted technological growth rate for changes in Cu and labor productivity

Shifting rolling regression results

Diagram B2.1: The d_govcivil coefficient with adjusted technological growth rate for changes in CU and labor productivity
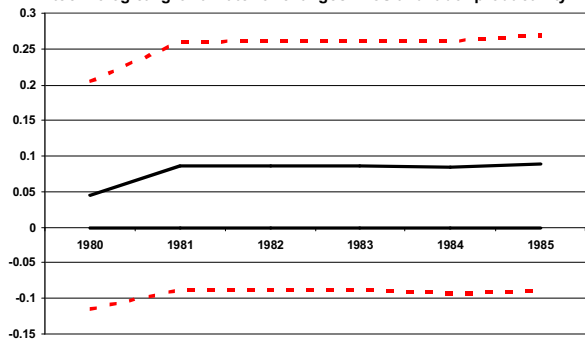
Diagram B2.2: The dtour coefficient with adjusted technological growth rate for changes in CU and labor productivity
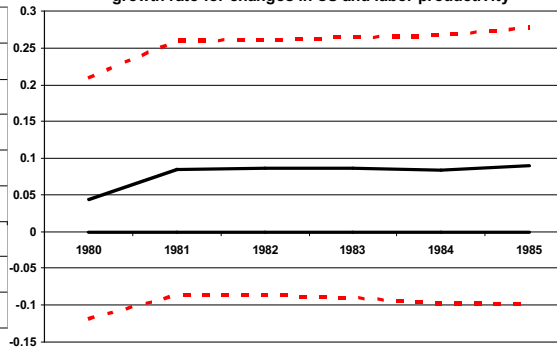
Diagram B3.2: The two-stage instrument with adjusted technological growth rate for changes in Cu and labor productivity

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