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Effective Demand, Exogenous Normal Utilization and Endogenous Capacity in the Long Run. Evidence from a CVAR Analysis for the U.S.

October 2013
Working Paper 06/2013
Department of Economics
The New School for Social Research

Effective demand, exogenous normal utilization and endogenous capacity in the long run. Evidence from a CVAR analysis for the US.*

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October 8, 2013

Abstract

Using Cointegrated Vector Auto-Regression analysis, we provide evidence for the US manufacturing sector that production capacities adjust endogenously to current output in the long run. The rate of capacity utilization, i.e. the output-capacity ratio, is found to be stationary since production capacities respond endogenously to changes in current output and not vice versa. Hence, the principle of effective demand in a growth context, by which a permanent demand shock has a permanent growth effect, is consistent with the stylized fact of a stationary rate of capacity utilization since production capacities are endogenous in the long run.

Keywords: Effective demand, stationary utilization rate, endogenous capacity, cointegrated vector autoregression
JEL Classification: E12, E22, C22

*For valuable comments and suggestions, I would like to thank Katarina Juselius and two anonymous referees. The usual caveats apply.

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1 Introduction

Since the contributions by Rowthorn (1981), Dutt (1984) and Taylor (1985), structuralist models have been widely used for studying the determinants of economic growth in the short and long run.[1] One core feature of the structuralist growth theory in the Kaleckian tradition is the principle of effective demand: the steady-state growth rate of the economy and the steady-state rate of capacity utilization, which adjusts savings to investment in the aggregate, are determined by aggregate spending decisions. Fallacies of composition such as the paradox of thrift and the paradox of cost may hold both in the short and long run.

Most traditions of economic thought perceive economic dynamics as demand-constrained in the short run. Yet, disagreements arise regarding the form and relevance of effective demand in the long run. For instance, growth models in the Classical tradition typically maintain that the long-run growth rate of the economy is structurally determined, i.e. independent of aggregate demand and consistent with an exogenous desired rate of capacity utilization.[2] In these models, long-run forces beyond the short-run dynamics come into play bringing the economy back to an exogenous growth rate. The paradoxes of thrift and cost disappear in the long run.

Naturally, a lively debate has emerged on the relevance of effective demand for long-run analysis, a considerable part of which focuses on the role of the rate of capacity utilization as a long-run accommodating variable in Kaleckian growth theory (cf. Lavoie et al. 2004, Hein et al. 2011, 2012, Schoder 2012b, Skott 2012). The critics of the Kaleckian growth model typically point towards a severe empirical shortcoming of the baseline model as objected by (Skott 2012): It predicts a non-stationary rate of capacity utilization if the economy is hit by a series of permanent demand shocks. Put differently, a permanent demand shock, ceteris paribus, implies a permanent change in the equilibrium utilization rate. Without doubt, the decline of the US saving rate since the 1980s constitutes a series of major permanent demand shocks which, according to the baseline model, should have caused a significant increase in utilization. Yet, data on the utilization rate such as the one published by the FED typically indicate stationarity despite long swings. Hence, this empirical observation may be interpreted as an indication for the existence of long-run forces that keep the utilization rate within a rather narrow band which is inconsistent with the predictions of the structuralist benchmark model. Kalecki (1968) may have been wrong in claiming that the long run is nothing but a sequence of short runs.[3]

[1] Much the subsequent literature has focused, in particular, on the growth effects of the re-distribution of income (cf. Bhaduri and Marglin 1990, Naastepad and Storm 2006-07), of monetary policy (cf. Lavoie 1995a, Hein 2007) and of institutional changes on financial markets referred to as *financialization* (cf. Stockhammer 2004, Hein and Schoder 2011). The nexus between growth and distribution has also been discussed in an open-economy context (cf. Blecker 1989).

[2] See, among others, Duménil and Lévy (1999), Shaikh (2007), Shaikh (2009) and Taylor (2012).

[3] In terms of theory, the canonical Kaleckian growth model has also been criticized since it does not require full adjustment, i.e. the consistency of expectations and realizations, in the long run. Hence, critics raised the question why firms should settle on a steady state in which the actual rate of capacity utilization is inconsistent with the desired rate (cf. Committeri 1986, Auerbach and Skott 1988). As a response to this criticism, the desired rate of capacity utilization as well as the secular rate of sales growth have ben

In recent contributions, Schoder (2012a,c) has attempted to reconcile the principle of effective demand in the long run and the stationarity of the rate of capacity utilization within a structuralist framework. This is achieved by introducing an endogenous, pro-cyclical full capacity output-capital ratio to an economy with de-stabilizing Harrodian investment dynamics and stabilizing distribution and debt dynamics. A persistent positive demand shock in the steady-state pushes the utilization rate beyond its target level causing an investment boom. The capacity-capital ratio increases while the utilization rate returns to the target since a rising wage share as well as a rising debt-capital ratio cut into investment demand. With a higher capacity-capital ratio, the long-run growth rate will be higher, too.

Since the assumption of a pro-cyclical capacity-capital ratio is crucial for the principle of effective demand in a structuralist model with output being the accommodating variable and a stationary utilization rate, Schoder (2012a) provides theoretical arguments for an endogenous capacity-capital ratio: First, optimizing firms can be argued to invest in machinery with higher productivity when running low on spare capacity. Second, sustainable full capacity output may depend on the number of shifts employed. With a higher number of shifts, the capacity reported by firms will also be higher which provides a rationale for a pro-cyclical measured capacity-capital ratio. Schoder (2012a) also provides empirical evidence: For various US industrial sectors from 1960Q3 to 2012Q2, the capacity-capital ratio is estimated as a function of the business cycle. The result is a positive response of the growth rate of the capacity-capital ratio to a change in the difference between utilization and trend utilization.

Analyzing the capacity-capital ratio empirically, however, suffers from low quality of capital stock data. Hence, the present paper seeks to complement the previous contributions by approaching the principle of effective demand, the endogeneity of relative capacity growth and the stationarity of the utilization rate from a statistical perspective. We derive an econometric model without normalization through the capital stock from a simple structuralist growth model which has been subject to criticism as of its prediction of a non-stationary utilization rate. We then derive testable hypothesis implied by the principle of effective demand in the context of this model and the stationarity of the utilization rate. In particular, we study the interaction of output, full capacity output and and a composite leading indicator for the US manufacturing sector from 1955Q1 to 2012Q2 employing a Cointegrated VAR model in the I(1) and I(2) analytical framework as developed by Johansen and Juselius (1990), Johansen (1995) and Juselius (2006). We find some evidence that the principle of effective demand by which a permanent demand shock has a permanent growth effect is consistent with a stationary rate of capacity utilization, since production capacities adjust slowly to output.

The remainder of the paper proceeds as follows. Section 2 motivates the econometric model consistent with the predictions of a structuralist growth model with instantaneous output adjustment. In section 3, the data used are discussed. Section 4 introduces the econometric specifications estimated, discusses potential misspecification and parameter instability issues and presents our main findings applying I(1) and I(2) analyses. Section 5

endogenized through hysteresis effects implying the economy to be fully adjusted in the long run (cf. Lavoie 1995b, 1996, Dutt 1997, 2009). Schoder (2012b) found some evidence for such hysteresis effects in the US.

concludes the paper.

2 Theoretical considerations

In this section a simple structuralst growth model with instantaneous output adjustment is outlined to study under which conditions the model predicts a non-stationary utilization rate. Then, an econometric variant of the model is derived and testable restrictions imposed in order to study empirically the endogeneity of the capacity-capital ratio which may cause the utilization rate to become stationary.

2.1 A simple structuralist growth model

To motivate the empirical model studied below, let us first consider a simple variant of a structuralist growth model for a closed economy with a government sector.[4] Variants of this model have been used extensively in the theoretical and empirical literature and have been criticized for predicting a non-stationary utilization rate (cf. Skott 2012). Yet, we will argue theoretically and empirically that a stationary utilization rate is consistent with this model.

The model economy comprises representative capitalists and workers as well as a government sector which carries out public investments. One type of good is produced used for investment and consumption. Prices are set according to a constant mark-up on unit costs. There is no overhead labor. Hence, the distribution of income is exogenous. We abstract from depreciation of the capital stock. For simplicity, we assume the propensities to save to be the same for capitalists and workers. In the aggregate, the economy is characterized by the following set of linear equations:

$$g_t^i = \alpha u_t^e + \beta(u_t - \bar{u}), \tag{1}$$

$$g_t^s = s_t u_t \sigma_t, \tag{2}$$

$$g_t^i + g_t^g = g_t^s. \tag{3}$$

Eq. (1) is the reduced-form investment function specifying the accumulation rate, i.e. the investment-capital ratio g_t^i. It is affected by the expected rate of capacity utilization, u_t^e, which is the ratio of expected output, Y_t^e, to capacity output, Y_t^c, and assumed to be exogenous. This term in the investment function may be interpreted as animal spirits. The accumulation rate is also affected by the difference between the current rate of capacity utilization, u_t, which is the ratio of current output, Y_t, to capacity output, Y_t^c, and the *normal* rate of utilization, $\bar{u}$, assumed to be exogenous.[5] If expected utilization goes up,

[4]For a discussion of the structuralist growth model with instantaneous output adjustment in greater detail, see, for instance, Bhaduri and Marglin (1990).

[5]Note that it is reasonable to expect the *normal* rate of capacity utilization to be below full capacity utilization. One rationale for this has been provided by Steindl (1952) arguing that firms need to be able to respond to demand shocks in time in order not to loose market shares to competitors. Another rationale is put forward by Schoder (2012a): For a given capital stock and a given number of shifts employed, the profit-maximizing level of output may be below the full-capacity level of output due to rising marginal costs.

firms raise investment to generate additional capacity. Eq. (2) is the saving function with s_t denoting the propensity to save which may vary over time. Note that aggregate income normalized by the capital stock, Y_t/K_t equals $u_t\sigma_t$ where $u_t = Y_t/Y_t^c$ is the utilization rate and $\sigma_t = Y^c/K_t$ is the capacity output-capital ratio. The macroeconomic balance condition in (3) follows from accounting and states that the sum of private investment and public investment is necessarily equal to aggregate saving. We assume g_t^g and s_t to be stochastic processes since we want to show that the model predicts a non-stationary utilization rate in the case of permanent demand shocks to the economy.

Substituting (1) and (2) into (3), and solving for u_t yields the equilibrium utilization rate, i.e.

$$u_t^* = \frac{\alpha u_t^e - \beta\bar{u} + g_t^g}{s_t\sigma_t - \beta}. \tag{4}$$

Substituting the equilibrium utilization rate, u_t^* into (2) yields the equilibrium accumulation rate, i.e.

$$g_t^* = g_t^i + g_t^g = s_t\sigma_t\frac{\alpha u_t^e - \beta\bar{u} + g_t^g}{s_t\sigma_t - \beta}. \tag{5}$$

Given the parameters of the investment decisions as well as the *normal* utilization rate, the equilibrium utilization and accumulation rates are determined by the animal spirits, public investment, the saving propensity and the capacity-capital ratio. A rise in the animal spirits as well as public investment raises the utilization rate through the multiplier effect and accelerates growth in equilibrium. At a higher capacity-capital ratio, a lower stream of income is required to generate the savings needed for financing investment at a given saving rate. Hence, the utilization rate and the growth rate decrease in equilibrium.[6]

The time series properties, i.e. the order of integration of u_t^* and g_t^* are determined by the time series properties of the time-varying variables, i.e. u_t^e, g_t^g, s_t and σ_t which are assumed to be exogenous in the baseline model.[7] To see this, let us take logarithms on both sides of (4) and (5).[8] We get,

$$\ln u_t^* = \ln(\alpha u_t^e - \beta\bar{u} + g_t^g) - \ln(s_t\sigma_t - \beta), \tag{6}$$

$$\ln g_t^* = \ln(\alpha u_t^e - \beta\bar{u} + g_t^g) - \ln(s_t\sigma_t - \beta) + \ln(s_t\sigma_t). \tag{7}$$

Suppose the economy is hit by a series of permanent demand shocks in any of the exogenous variables g_t^g, s_t and σ_t which could be reflected by specifying these variables as I(1)-processes

[6]The equilibrium utilization rate, u_t^*, will generally be different from the *normal* rate, $\bar{u}$, since there are no mechanism built into this simple model which would align the two measures. This implication of the model has been debated extensively (cf. Committeri 1986, Auerbach and Skott 1988, Lavoie 1995b, 1996, Dutt 1997, 2009, Schoder 2012b).

[7]Loosely speaking, a stochastic process is integrated of order k, i.e. I(k), if and only if it is stationary, i.e. I(0), after first-differencing k-times.

[8]Note that neither taking logs nor multiplying by a non-zero constant changes the order of integration of a time series.

such as random walks. Assume that the expected rate of utilization, u_t^e, cannot deviate too much from the realized rate, u_t, therefore sharing the same stochastic trend, i.e. being integrated of the same order. Then, u_t^* will only be stationary if there exists a relationship between g_t^g, s_t and σ_t such that the right-hand-side of (4) is stationary. The equilibrium growth rate, g_t^*, will then generally be I(1) unless both g_t^g and the product $s_t\sigma_t$ are I(0).

In the baseline model, σ_t is typically assumed to be stationary, i.e. an I(0) process, or constant (cf. Taylor 2004, Skott 2012, Hein et al. 2012). Moreover, no relationship between g_t^g and s_t has been derived in the Kaleckian literature such that u_t^* is stationary. Under these assumptions, the model predicts a permanent shift of u_t^* as a response to a permanent shift in g_t^g or s_t. Hence, u_t^* is predicted to be I(1) if the shocks to g_t^g or s_t are I(1). As argued by Skott (2012), this model feature is inconsistent with the empirical observation of a rather stationary utilization rate in the US fluctuating between a narrow band around 80%. This suggests the presence of a long-run adjustment mechanism which brings the utilization rate back to the mean and ensures u_t^* to be I(0).

2.2 Reconciling stationary utilization with permanent shocks

What responses of g_t^g, s_t and σ_t to permanent changes in any of these variables are plausible to close the utilization gap, i.e. the difference between u_t^* and $\bar{u}$, within the model framework outlined above? Consider a fall in s_t or σ_t causing u_t^* to rise beyond its target. g_t^g has to decrease for u_t^* to return to its mean. This implies anti-cyclical fiscal policy as observed empirically. If this is the main mechanism through which the utilization rate returns to its mean, then, in term of our econometric analysis, output will be endogenous to the utilization gap. Yet, our results below indicate that output does not respond to the utilization gap.

Now consider a rise in g_t^g or a fall in σ_t causing u_t^* to move beyond target. Using a slightly more elaborate model Shaikh (2009) has argued that s_t increases to close the utilization gap. This is because firms are assumed to raise the retention ratio on profits, which is a component of the overall saving rate, whenever the accumulation rate exceeds the accumulation rate at normal profits, i.e. whenever the utilization rate exceeds the target utilization rate. Yet, Hein et al. (2011) question the plausibility of this mechanism. While it may be reasonable to assume firms to aim for a *higher* retention rate when accumulation rates are above normal, it is not clear why this rate should keep rising if the utilization rate and the accumulation rate are constant.

Finally, consider a rise in g_t^g or a fall in s_t causing u_t^* to move beyond target. σ_t has to rise to close the utilization gap. In terms of our econometric analysis, this implies that capacity has to be endogenous to the utilization gap.

There are good reasons for a pro-cyclical capacity-capital ratio as pointed out in detail by Schoder (2012a).[9] As argued by Nikiforos (2013) full-capacity output as reported by the FED does not measure the technically feasible capacity but the highest level of output

[9]Regarding the trend of the variable, a large body of literature analyzes if technical change is labor saving or augmenting in the long run, i.e. if the capacity-capital ratio tends to decrease or increase. In industrialized countries technical change has been found to be slightly labor saving in the long run (cf. Foley and Michl 1999, pp. 37-41 and Duménil and Lévy 2004).

that can be produced under *normal* conditions and maintained sustainably. Indivisibilities in the production process such as shift work may then cause capacity output to change endogenously. For every number of shifts in operation, the firm may face different cost curves. Running another shift is associated with additional fixed costs but reduces unit variable costs since over-time labor can be saved. Hence, with a low number of shifts the sustainable capacity output, i.e. the output level for which unit costs are still lower than unit revenues, will be lower than for a high number of shifts. In a boom with high demand expectations some firms may introduce additional shifts and, hence, raise their full-capacity output even though no capital investment need to have taken place. Capacity and, therefore, the capacity-capital ratio will be endogenous.

Moreover, investment induced technical change may affect the capacity-capital ratio pro-cyclically (cf. Schoder 2012a). If a deviation of the utilization rate from the desired rate implies some form of costs arising from an inefficient use of resources, on the one hand, and a lack of flexibility in accommodating demand required to deter market entry of potential competitors, on the other, then a firm will seek to invest in capital which helps realigning the utilization rate to the desired rate. For instance, if utilization is too high, a firm will choose structures and equipment which raise the productivity of capital, since this increases capacity output and, therefore, reduces utilization for a given demand to be accommodated. Hence, the capacity output-capital ratio will move pro-cyclically.

2.3 Deriving testable hypotheses

In the remainder of the paper, we seek to empirically test whether output and capacity adjust endogenously to bring back the utilization rate to the target level. To set up an econometric model based on the theoretical model outlined above, two issues need to be addressed.

First, Schoder (2012a) estimates the adjustment of the capacity-capital ratio in response to a gap between the utilization rate and the target rate for 17 US sectors from 1960Q3 to 2012Q2 and finds a significantly positive adjustment parameter. There are two weaknesses of this approach. First, the target utilization rate is unknown and has to be approximated which is done by using an HP filter and checking the robustness of the results with respect to different smoothing parameters. Second, the data on the capital stock is in general of modest quality. In the present paper, we seek to analyze the endogeneity of capacity without requiring knowledge of the target utilization rate or the capital stock. Cointegration analysis allows us to achieve this by studying whether capacity and/or output adjust endogenously in the long run when the utilization rate deviates too much from its mean.

Second, in the model above we assumed expectations regarding the future output to be backward-looking. In the empirical model, we will use a leading indicator to capture output expectations.

Let us now derive our econometric specification. Suppose there exist two linear relationships between the log of real output, y_t, the log of full capacity output, y_t^c, and the log of

expected demand, y_t^e, of the following form:

$$\beta_{11}y_t + \beta_{21}y_t^c + \beta_{31}y_t^e + \beta_{01} = \mu_{1,t}, \tag{8}$$
$$\beta_{12}y_t + \beta_{22}y_t^c + \beta_{32}y_t^e + \beta_{02} = \mu_{2,t} \tag{9}$$

where $\mu_{1,t}$ and $\mu_{2,t}$ are possibly highly persistent but stationary random disturbance terms.[10] We use empirical stylized facts and the results of the Kaleckian model to motivate parameter restrictions on (8) and (9).

Logarithmized capacity utilization, υ_t, is defined as $\upsilon_t \equiv y_t - y_t^c$. The FED provides data on the rate of capacity utilization which is consistent with the definition of υ_t. Despite long swings, it can be supposed that capacity utilization, structurally determined by the firms optimization problems, is stationary (Skott 2012) in the long run. Our results below confirm this suggestion. We assume υ_t to be an I(0)-process around a constant mean, $\bar{\upsilon}$. Hence, we can impose the restrictions $\beta_1 1 = 1$, $\beta_2 1 = -1$, and $\beta_3 1 = 0$ for (8) to represent this utilization relation. Then, we obtain

$$y_t - y_t^c = -\beta_{01} + \mu_{1,t}, \tag{10}$$

where $\beta_{01} \equiv -\bar{\upsilon}$. The log of the utilization rate is equal to a constant plus a random stationary disturbance term.

An output relation can be obtained by slightly modifying the structuralist model above. Let us consider an economy characterized by the following set of equations:

$$I_t = f_1(Y_t^e) + f_2(Y_t - \bar{Y}), \tag{11}$$
$$S_t = s_t Y_t, \tag{12}$$
$$I_t + G_t = S_t, \tag{13}$$

where I_t, S_t, Y_t and Y_t^e are investment, savings, output and expected demand, respectively, in levels. $f_1(\cdot)$ and $f_2(\cdot)$ are functions. Solving (11) to (13) for Y_t yields a non-linear function in Y_t^e and G_t. Log-linearizing the result leads to the output relation which is equivalent to (9) with $\beta_{12} = 1$ and $\beta_{22} = 0$. Hence,

$$y_t = -\beta_{32}y_t^e - \beta_{02} + \mu_{2,t}, \tag{14}$$

where we expect $\beta_{32} < 0$ and $\mu_{2,t}$ captures the demand shocks caused by changes in public investment and the saving rate.

Given the structure of the simple model characterized in (11) to (13), the principle of effective demand in the long run combined with the stylized fact of a stationary utilization rate imply the following predictions which will be tested in the econometric section below.

Hypothesis 1. *There exist two cointegrating relationships between y_t, y_t^c, y_t^e and a constant of the form CI(1,1) or CI(2,2).*[11] *The first one, the utilization relation, is characterized*

[10] Additional relationships between these variables may exist. Yet, we only consider two since this is the number of relationships obtained in the cointegration analysis below.

[11] Variables are *cointegrated of order* CI(d,p) if each variable is integrated of order d and there exists a linear relationship, i.e. a cointegration relation, between the variables which is integrated of order d-p.

by the vector $\boldsymbol{\beta_1'} = (1, -1, 0, \beta_{01})$. *The second one, the output relation, is characterized by* $\boldsymbol{\beta_2'} = (1, 0, \beta_{32}, \beta_{02})$.

For the empirical analysis, the assumption that the sum of all demand shocks and parameter shifts is non-stationary may be too strong. Despite some considerable shocks and shifts in the US since the 1950s, modeling them as a non-stationary process may be premature as they cannot wander off arbitrarily in the long run. Recall that, in the structuralist framework outlined above, the growth rate of output in equilibrium is either I(0) or I(1) and, therefore, output (and, hence, capacity output and expected output) either I(1) or I(2). Hence, our empirical analysis will cover both cases: y_t, y_t^c and y_t^e being I(1) and I(2).

The restrictions on the cointegration vectors $\boldsymbol{\beta_1'}$ and $\boldsymbol{\beta_2'}$ follow from the theoretical considerations leading to (10) and (14), respectively.

Hypothesis 2. *Since* y_t *is fully characterized by* y_t^e, y_t *is error-correcting to the output relation but not to the utilization relation.*

In the model above, output is solely driven by expected demand apart from an exogenous shock caused by changes in public investment and the saving rate. Hence, it should respond to a disequilibrium in the output relation in an error-correcting way. For instance, if output undershoots expected demand, it will increase towards closing the gap. Yet, output should not respond to a gap between the utilization rate and its mean since, in the structuralist model with instantaneous output adjustment, output is not the variable bringing back the utilization rate to its mean, despite anti-cyclical fiscal policy.

Hypothesis 3. *Since* y_t^c *is endogenous it should be error-correcting to the utilization relation.*

The hypothesis of long-run output moving independently of the utilization rate as stated in hypothesis 2 combined with the assumption of a stationary utilization rate requires capacity to be the accommodating variable. It is capacity output which brings back the utilization rate to the mean, not realized output. Note that this view contradicts much of the orthodox literature which typically perceives capacity as exogenous and as a strong attractor for realized output.

Hypothesis 4. *Since* y_t^e *is predetermined it should be weakly exogenous.*

In the model, we have assumed expected demand to be exogenous which implies that it does not respond to disequilibria in any of the two suggested cointegration relations and is, therefore, weakly exogenous.

To sum up, we have established the conditions for the principle of effective demand and the stationarity of the rate of capacity utilization to hold at the same time: y_t^c has to adjust endogenously to the stochastic trend in y_t which, in turn, follows y_t^e which, in turn, is independent of y_t^c.

One could object that a y_t^c being affected by y_t is not inconsistent with an economy featuring endogenous utilization in the short run but exogenous utilization in the long run such as Duménil and Lévy (1999) and Shaikh (2009) due to the capacity building effect of

investment. Yet, in heterodox macro models the capacity effect is typically super-fast. A rise in investment simultaneously leads to a higher capital stock and, therefore, a higher capacity. In reality, a rise in the flow (investment) leads to a change in the stock (capital) with some delay. In our CVAR analysis, this effect will be captured by the short-run dynamics of the econometric model. The cointegrating relation will capture the long-run effect of output on capacity through changes in the capacity-capital ratio which is on a different time scale than the relatively fast effect of output on capacity through a higher level of capital.[12]

3 Data

We employ quarterly data from 1955Q1 to 2012Q2. For y_t and y_t^c, we use the logs of the production index and the full-capacity index, respectively, for the US manufacturing sector provided by the FED. y_t^e is approximated by the trending Composite Leading Indicator provided by the OECD which is an average of business and consumer confidence indicators. All variables are seasonally adjusted.[13]

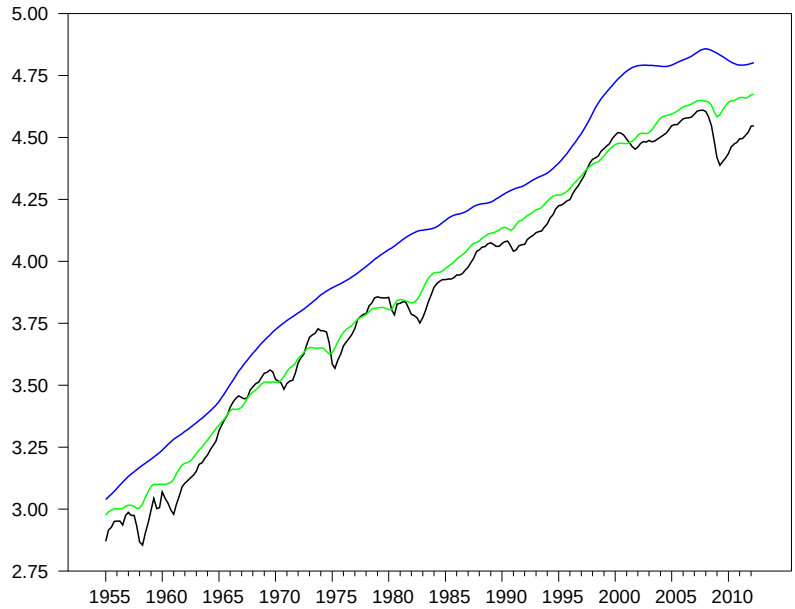

Figure 1: The logs of the production index (black), capacity index (blue) and the composite leading indicator (green) for the US in levels.

[12]DeLong and Summers (2012) have analyzed the short-run capacity effect of utilization (triggered by changes in the level of capital rather than its productivity) by regressing the growth rate of capacity output on the two years lagged utilization rate (both in percent) and find a slope coefficient of 1.88.

[13]All three series are also available in monthly frequency. Yet we chose to use quarterly data since the quality of the Composite Leading Indicator is not satisfactory in the first part of the sample with constant values for several months followed by sudden changes.

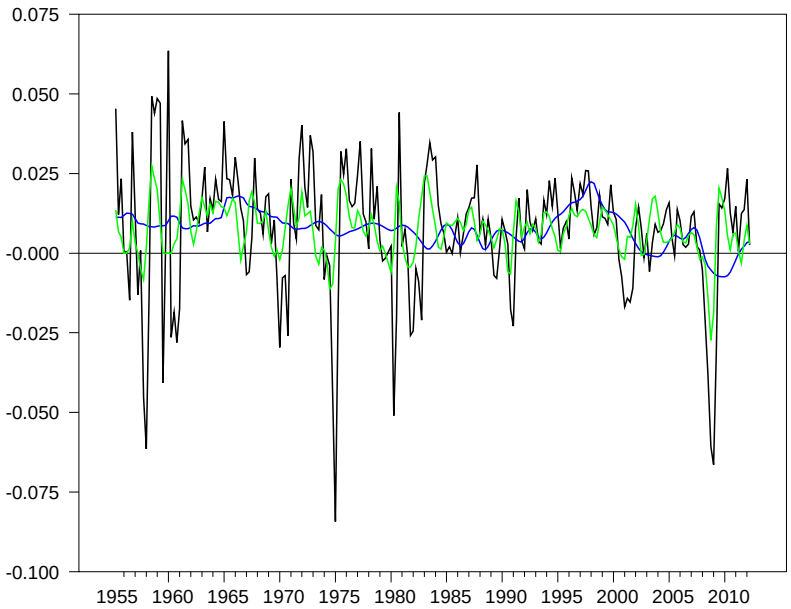

Figure 2: The logs of the production index (black), capacity index (blue) and the composite leading indicator (green) for the US in first differences.

The three time series used are plotted in levels and first differences in Figures 1 and 2, respectively. Graphical inspection yields the following noteworthy insights:

First, all variables follow the same stochastic trend. Even though the plot in first differences indicates that, apart from y_t^c which is rather smooth, y_t and y_t^e may well be I(1), one could interpret the stochastic trend as an I(2) process since the volatility of these series may blur the picture and make their first differences appear more stationary than they are.

Second, as confirmed by the plot in differences, y_t^e is leading and closely correlated with y_t.

Third, the level plot provides some indication of y_t^c adjusting slowly to changes in y_t. Note, further, that capacity output seems to have been smoothed by the FED. This will affect the short-run dynamics of the model but not the cointegrating relations.

4 Econometric analysis

Since y_t, y_t^c and y_t^e may be perceived as I(1) or I(2) processes, we conduct both I(1) and I(2) analyses. An I(1) model is much simpler to interpret than an I(2) model. Note that, even if our variables followed an I(2) stochastic trend, the I(1) analysis would yield consistent estimates of the $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}'$ matrices. Below, we will check the robustness of our result by pursuing an I(2) analysis.

4.1 I(1) analysis

To test the hypotheses 1 to 4 posed above, we apply a cointegration analysis developed by Johansen and Juselius (1990), Johansen (1995) and Juselius (2006) and estimate the following VAR model in VECM representation:

$$\Delta \mathbf{x}_t = \boldsymbol{\alpha} \begin{bmatrix} \boldsymbol{\beta}' & \boldsymbol{\beta_0}' \end{bmatrix} \begin{bmatrix} \mathbf{x}_{t-1} \\ c \end{bmatrix} + \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \Delta \mathbf{x}_{t-1} + \boldsymbol{\Phi} \mathbf{D}_t + \boldsymbol{\varepsilon}_t, \tag{15}$$

where $\mathbf{x}_t = \begin{bmatrix} y_t & y_t^c & y_t^e \end{bmatrix}'$, $\mathbf{D}_t$ is a matrix of deterministic variables and $\boldsymbol{\varepsilon}_t \sim \mathbf{I}\mathcal{N}_p(\mathbf{0}, \boldsymbol{\Omega})$ is a vector of disturbances. We include a constant term restricted to the cointegrating space since utilization fluctuates around a constant mean. $\mathbf{D}_t$ comprises dummies for the following quarters since they feature large outliers: transitory dummies for 1959Q3 to 1959Q4 and for 1965Q1 to 1965Q2 as well as a permanent dummy from 1975:1. We chose $k = 4$ as the optimal lag length following the suggestion of the SBC information criteria. Moreover, including fewer lags in the model would lead to severe serial correlation problems.[14]

4.1.1 Misspecification tests

Table 1 reports the tests for residual normality, independency and homoskedasticity. While there is still some evidence for first-order autocorrelation the Ljung-Box test as well as the LM test for second-order autocorrelation reject the null. Normality as well as homoscedasticity of the residuals are rejected. This may indicate non-modeled non-linear effects. The fact that our variables exhibit strong persistence after taking first differences may also contribute to non-normality and heteroskedasticity. Including more dummies for outliers allows us to accept the null of normality and homoscedasticity but inflates the model enormously without changing the estimation results below. Hence, we only report the results of the more parsimonious specification. Some caution when interpreting the significance tests below is in order. As shown in the section presenting the I(2) analysis, in which non-normality and heteroskedasticity cease to be a problem since variables are differentiated twice, the results of the I(1) analysis are robust.

The recursive and backwards recursive tests of $\boldsymbol{\beta}(t)$="known beta", of beta constancy and of eigenvalue fluctuation indicate parameter stability for the unrestricted model as well as for all restrictions considered below.

4.1.2 Rank test

Our theoretical considerations suggest two cointegrating relationships between our three variables since they all follow the same stochastic trend which seems to be confirmed by the level plot in Figure 1. Table 2 reports the rank test statistics.

Note the large difference between the p-values of the trace test and the Bartlett corrected trace test for 1 and 2 ranks. This suggests that our variables may be I(2) which is also

[14]The econometric analysis has been conducted using CATS 2.0 for RATS 8.2.

Table 1: Tests for autocorrelation, residual normality and homoskedasticity

Trace Correlation:		=	0.793	
Tests for Autocorrelation				
Ljung-Box(56):	ChiSqr(468)	=	444.421	{0.777}
LM(1):	ChiSqr(9)	=	18.861	{0.026}
LM(2):	ChiSqr(9)	=	5.938	{0.746}
Test for Normality:	ChiSqr(6)	=	46.630	{0.000}
Test for ARCH				
LM(1):	ChiSqr(36)	=	112.304	{0.000}
LM(2):	ChiSqr(72)	=	184.097	{0.000}

Notes: p-values in curly brackets.

Table 2: Rank test

p-r	r	Eig.Value	Trace	Trace*	Frac95	P-Value	P-Value*
3	0	0.166	65.080	40.505	35.070	0.000	0.011
2	1	0.069	24.111	9.934	20.164	0.012	0.651
1	2	0.035	7.967	1.941	9.142	0.085	0.785

confirmed by the estimated roots of the companion matrix. Regardless which rank is selected, the first unrestricted root is always larger than 0.98, indicating an I(2) stochastic trend. In this case, the trace tests become unreliable. Nevertheless, according to the uncorrected trace test, we can accept the hypothesis of $r = 2$ which is consistent with the theoretical prior.

4.1.3 Testing restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}'$

Using the I(1) analysis, Table 3 reports the estimates of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}'$ for different set of restrictions. The model including the restrictions derived from theory is reported in model (a). Note there is one overidentifying restriction which the LR test cannot reject. Note further that we leave the constant in the output relation between y_t and y_t^e unrestricted. This constant is restricted in model (b). The two overidentifying restrictions are now rejected. For some reason, however, the estimate of the constant, 0.389, in the utilization relation between y_t and y_t^c in model (a) is inconsistent with the data. It implies a long-run equilibrium of $y_t - y_t^c = -0.389$ which is equivalent to $u_t = 0.677$ in equilibrium. Yet, the mean around which u_t fluctuates with some persistence is 0.802. To correct for this inconsistency, we additionally restrict the constant in the first cointegrating relation to $0.096 = -\log(0.802)$. The estimates are reported in model (c). Note that the LR test rejects the overidentifying restrictions in this case. Model (d) reports the estimate of the model restricting the constant of the first relation to 0.096 and the constant of the second relation to 0. Yet the LM test

Table 3: Estimation results for restricted I(1) models

	Model (a) $\chi^2(1) = 0.060$ {0.807}				Model (b) $\chi^2(2) = 8.007$ {0.018}			
	y	y^c	y^e	*const.*	y	y^c	y^e	*const.*
The cointegrating relations $\boldsymbol{\beta}$								
$\boldsymbol{\beta_1}'$	1.000	−1.000	0.000	0.389 [13.130]	1.000	−1.000	0.000	0.015 [0.392]
$\boldsymbol{\beta_2}'$	1.000	0.000	−1.093 [−34.698]	0.864 [5.680]	1.000	0.000	−1.089 [−58.116]	0.000
The adjustment coefficients $\boldsymbol{\alpha}$								
$\boldsymbol{\alpha_1}'$	0.020 [0.829]	0.003 [3.406]	−0.014 [−1.852]		0.039 [1.685]	0.003 [3.445]	−0.012 [−1.599]	
$\boldsymbol{\alpha_2}'$	−0.029 [−3.480]	−0.001 [−3.163]	0.006 [2.254]		−0.005 [−0.398]	−0.002 [−3.295]	0.005 [1.138]	

	Model (c) $\chi^2(2) = 8.187$ {0.017}				Model (d) $\chi^2(3) = 8.951$ {0.030}			
	y	y^c	y^e	*const.*	y	y^c	y^e	*const.*
The cointegrating relations $\boldsymbol{\beta}$								
$\boldsymbol{\beta_1}'$	1.000	−1.000	0.000	0.096	1.000	−1.000	0.000	0.096
$\boldsymbol{\beta_2}'$	1.000	0.000	−1.066 [−16.245]	−0.389 [−1.331]	1.000	0.000	−1.086 [−123.343]	0.000
The adjustment coefficients $\boldsymbol{\alpha}$								
$\boldsymbol{\alpha_1}'$	0.004 [0.157]	0.003 [3.319]	−0.012 [−1.635]		0.022 [0.905]	0.003 [3.506]	−0.013 [−1.693]	
$\boldsymbol{\alpha_2}'$	0.010 [1.678]	−0.001 [−3.051]	0.002 [0.951]		0.008 [0.788]	−0.001 [−3.253]	0.003 [1.042]	

Notes: t-statistics are in brackets, p-values in curly brackets.

rejects the overidentifying restrictions.

If not indicated otherwise, the following results hold for all specifications: First, in equilibrium a one-percent increase in y_t^e is associated with a statistically significant more-than-one-percent increase in y_t with coefficients ranging from 1.066 to 1.093. This might indicate a long-run multiplier effect.

Second, y_t is not error-correcting to the utilization rate, i.e. the first cointegrating relation, since the corresponding loadings, i.e. the elements of the matrix $\boldsymbol{\alpha}$, are all insignificant and have the wrong sign. In the first specification, y_t error-corrects the second relation. That means, excess output implies a reduction of output in the succeeding period.

Third, in all specifications, y_t^c is error-correcting to the utilization rate with a small but significant coefficient of 0.003. A positive deviation of utilization from its long-run mean leads to a slow acceleration of full-capacity output. The change in y_t^c is also affected by the output relation. A y_t exceeding its equilibrium level causes y_t^c to decrease slightly. This finding is not easy to interpret since one would expect an output disequilibrium at a given utilization rate not to affect capacity. Note that a disequilibrium in the first relation is highly

Table 4: Tests of weak exogeneity

	Model (a)	Model (b)
	$\chi^2(1) = 0.060$ {0.807}	$\chi^2(2) = 8.007$ {0.018}
$H_0 : \alpha_{11} = \alpha_{12} = 0$	$\chi^2(3) = 32.290$ {0.000}	$\chi^2(4) = 30.913$ {0.000}
$H_0 : \alpha_{21} = \alpha_{22} = 0$	$\chi^2(3) = 9.693$ {0.021}	$\chi^2(4) = 13.929$ {0.008}
$H_0 : \alpha_{31} = \alpha_{32} = 0$	$\chi^2(3) = 3.949$ {0.267}	$\chi^2(4) = 17.056$ {0.002}
	Model (c)	Model (d)
	$\chi^2(2) = 8.187$ {0.017}	$\chi^2(3) = 8.951$ {0.030}
$H_0 : \alpha_{11} = \alpha_{12} = 0$	$\chi^2(4) = 32.030$ {0.000}	$\chi^2(5) = 32.081$ {0.000}
$H_0 : \alpha_{21} = \alpha_{22} = 0$	$\chi^2(4) = 13.442$ {0.009}	$\chi^2(5) = 15.151$ {0.010}
$H_0 : \alpha_{31} = \alpha_{32} = 0$	$\chi^2(4) = 13.131$ {0.011}	$\chi^2(5) = 13.952$ {0.016}

Notes: p-values in curly brackets.

correlated with a disequilibrium in the second relation. Hence, one can ask the question in what direction an increase in y_t causes y_t^c to change, in equilibrium. Since the loading of the utilization relation is larger than the loading of the output relation, y_t^c will increase and, hence, stabilize the system.

Fourth, the loadings to y_t^e are insignificant in almost all specifications which may suggest weak exogeneity of the Composite Leading Indicator. In fact, the LR test of weak exogeneity of y_t^e in the completely unrestricted model (not reported) cannot be rejected, whereas weak exogeneity can be rejected for all other variables. Table 4 reports the test results of the LR test of overidentifying restrictions for the models considered if weak exogeneity of y_t, y_t^c and y_t^e, respectively, is additionally imposed. Compared to the benchmark test statistics of the models without restrictions on $\boldsymbol{\alpha}$, restrictions on the loadings to y_t^e cause the smallest increases in the test statistics. Hence, there seems to be some evidence for y_t^e to be exogenous.

Overall, we have found evidence in support of the hypotheses 1 to 4 postulated above using the I(1) analytical framework. In the following, we will apply the I(2) framework to analyze these hypotheses.

4.2 I(2) analysis

Equivalently to the I(1) analysis, we estimate the following model:

$$\Delta^2 \mathbf{x}_t = \boldsymbol{\alpha}\{\begin{bmatrix}\boldsymbol{\beta}' & \boldsymbol{\rho_0}'\end{bmatrix}\} \begin{bmatrix}\mathbf{x}_{t-1} \\ t\end{bmatrix} + \begin{bmatrix}\boldsymbol{\delta}' & \boldsymbol{\gamma_0}'\end{bmatrix} \begin{bmatrix}\Delta\mathbf{x}_{t-1} \\ c\end{bmatrix} + \boldsymbol{\zeta} \begin{bmatrix}\boldsymbol{\beta}' & \boldsymbol{\rho_0}' \\ \boldsymbol{\beta}'_{\perp 1} & \tilde{\boldsymbol{\gamma}}_{\mathbf{0}}'\end{bmatrix} \begin{bmatrix}\Delta\mathbf{x}_{t-1} \\ c\end{bmatrix} + \sum_{i=1}^{k-2} \boldsymbol{\Gamma}_i \Delta\mathbf{x}_{t-1}^2 + \boldsymbol{\Phi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t, \tag{16}$$

where, as above, $\mathbf{x}_t = \begin{bmatrix}y_t & y_t^c & y_t^e\end{bmatrix}'$, $\mathbf{D}_t$ is a matrix of deterministic variables and $\boldsymbol{\varepsilon}_t \sim \mathbf{I}\mathcal{N}_p(\mathbf{0}, \boldsymbol{\Omega})$ is a vector of disturbances. To include a constant term but no linear trend in the cointegration space, we restrict $\boldsymbol{\rho_0}' = 0$ and leave $\boldsymbol{\gamma_0}'$ unrestricted.

Table 5: Rank test

Approximate 95% fractiles					
$p-r$	r	$s_2=3$	$s_2=2$	$s_2=1$	$s_2=0$
3	0	89.020	69.376	53.921	42.770
2	1		48.520	34.984	25.731
1	2			20.018	12.448

If we followed the advice of the I(2) rank test reported in Table 5, there would be no cointegrating relationship and one stochastic I(2) trend. This may be because both the utilization relation and the output relation are highly persistent. Since, our theoretical prior implies two cointegrating relationships and one stochastic trend, we select $r = 2$ and $s_2 = 1$.

The estimation results for the model with restrictions on the trend in the cointegration space as well as the restrictions derived from theory are reported in Table 6. Even though, we have to reject these restrictions according to the LR test, the results are very similar to the I(1) analysis.[15] Again, there is an accelerator effect of y_t^e on y_t in the long run.

The analysis of the error correction mechanism is more complicated in the I(2) framework. If $\alpha_{ij}\delta_{ij} < 0$ than $\Delta^2 x_{i,t}$ is error correcting to $\Delta x_{i,t}$ and if $\delta_{ij}\beta_{ij} > 0$ than $\Delta x_{i,t}$ is error-correcting to $x_{i,t-1}$. Hence, comparing the estimates for $\boldsymbol{\beta_1}$ and $\boldsymbol{\delta_1}$ reveals that a change in y_t equal to a change in y_t^c leaving the utilization rate constant has no effect on the acceleration of the variables. Yet, Δy_t is not error-correcting to capacity utilization, whereas Δy_t^c is. Comparing $\boldsymbol{\delta_1}$ and $\boldsymbol{\alpha_1}$ reveals that only $\Delta^2 y_t^c$ is significantly error-correcting to the utilization relation. For the output relation the estimates imply that Δy_t is not error-correcting to output, whereas Δy_t^e is. Moreover, as in the I(1) analysis, both $\Delta^2 y_t$ and $\Delta^2 y_t^c$ are error-correcting to output.

Hence, the I(2) analysis confirms the result that y_t is error-correcting to the output relation but not to the utilization relation, that there is some endogenous error-correcting adjustment of y_t^c from the utilization relation and that there is no significant feedback of the utilization relation on y_t^e.

The I(2) analysis, however, reveals some additional insights. First, the estimates of $\beta'_{\perp 1}$ suggest that there is almost a perfectly proportional relationship between all variables in the medium run which is not surprising given the similarity of the time series used. Second, the estimates of $\tilde{\boldsymbol{\beta}}'_{\perp 2}$ indicate that the I(2) trend affected all variables equally. Third, the estimates of $\boldsymbol{\alpha}'_{\perp 1}$ suggest that the common I(2) trend has primarily been generated by the twice cumulated shocks to capacity output and to a lesser extent to the composite leading indicator. Fourth, the last row in the table confirms that the disturbances to y_t and y_t^e have a higher standard deviation than the ones to y_t^c.

[15] Note that the restrictions cannot be rejected, if we allow for a linear trend in the cointegration space which only changes the results of the output relation. Since, however, the restrictions imposed are not the subject of debate, we report the results of the model which is most consistent with our theoretical priors.

Table 6: Estimation results for restricted I(2) model

$\chi^2(3) = 8.567$ {0.036}	y	y^c	y^e	*const.*	*trend*
The cointegrating relations β					
β_1'	1.000	−1.000	0.000	—	0.000
β_2'	1.000	0.000	−1.083 [−8.596]	—	0.000
The cointegrating relations δ					
δ_1'	−7.484	−7.484	−6.909	0.395	—
δ_2'	−18.586	−18.586	−17.158	0.810	—
The adjustment coefficients α					
α_1'	0.020 [0.857]	0.003 [3.489]	−0.013 [−1.728]		
α_2'	−0.030 [−3.386]	−0.001 [−3.214]	0.007 [2.285]		
The adjustment coefficients $\beta_\perp$					
$\beta_{\perp 1}'$	1.000	1.000	0.923		
$\tilde{\beta}_{\perp 2}'$	0.006	0.006	0.006		
The common trends $\alpha_\perp$					
$\alpha_{\perp 1}'$	1.000	42.809 [35.023]	11.352 [56.559]		
σ_ε	0.010	0.000	0.003		

Notes: t-statistics are in brackets, p-values in curly brackets.

5 Concluding remarks

Schoder (2012c) attempts to reconcile the principle of effective demand and the stationarity of the rate of capacity utilization within a structuralist framework with instantaneous output adjustment by introducing an endogenous capacity-capital ratio. Some theoretical arguments and empirical evidence for a pro-cyclical behavior of this ratio have been put forward by Schoder (2012a). The present paper has sought to complement these other contributions by analyzing the long-run behavior of capacity without normalizing the variables by the capital stock which facilitates theoretical reasoning but aggravates empirical analysis due to the low quality of capital stock data.

Using the Cointegrated VAR framework of Johansen and Juselius (1990), Johansen (1995) and Juselius (2006), we provide evidence that the principle of effective demand by which a permanent demand shock has a permanent growth effect is consistent with a stationary rate of capacity utilization, since production capacities adjust slowly to output.

We show that the principle of effective demand in a structuralist framework with instantaneous output adjustment implies output to follow an I(1) process in case of I(0)-shocks to the economy and an I(2) process in case of I(1)-shocks. Moreover, we take the stationarity

of the rate of capacity utilization relating output and full-capacity output as a stylized fact which is confirmed by our results. Using the composite leading indicator as a proxy for demand expectations exogenous from output and capacity output, we derive two steady state relations between output, capacity output and demand expectations. The principle of effective demand and the stationarity of the utilization rate can then be shown to be consistent with each other if capacity output adjusts endogenously.

We apply the cointegrating VAR framework to study this question for the US manufacturing sector from 1955Q1 to 2012Q2. We find evidence that there are two cointegrating relationships between the three variables of the form CI(1,1), meaning that there exists a linear relationship between I(1)-variables which is I(0), or of the form CI(2,2), meaning that there exists a linear relationship between I(2)-variables which is I(0): The utilization relation is between output and full-capacity output. The output relation is between output and the composite leading indicator.

We find that output is fully characterized by the composite leading indicator, since output is error-correcting to the output relation while it is not error-correcting to the utilization relation. Hence, output deviating from expected demand will be corrected by an adjustment of output. Yet, output deviating from capacity output plus some constant will not be corrected by an adjustment of output. Instead, we find evidence for capacity output to be error-correcting to the utilization relation, i.e. to be endogenously adjusting. This means, capacity adjusts in order to bring back the utilization rate to its mean. Finally, we find evidence for the composite leading indicator to be weakly exogenous.

One core implication of this analysis for the debate between advocates and critiques of the principle of effective demand in structuralist growth models is that it should not focus on the question of stationarity of the utilization rate. We provide some evidence that capacity output adjusts endogenously which makes the principle of effective demand and the stationarity of the utilization rate consistent with each other.

Another implication is that there does not seem to exist a natural growth rate of capacity which can be perceived as a long-run attractor for output growth. We find evidence for the contrary: Capacity follows realized output such that utilization behaves in a mean-reverting way.

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