A Fairytale Ending Or The Same Old Story?
The New Economy And Economic Growth in the United States

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Abstract

This paper uses a production function approach to put the recent contribution of the New Economy to US economic growth into perspective by undertaking an analysis of the sources of technical progress in the business sector over the post-war period. We model jointly a pair of factor demand equations derived consistently from an underlying CES production technology, and explicitly endogenise technical progress. Knowledge accumulation via R&D and education are found to be the main sources of technical progress, but there is evidence of significant structural change after 1995 which can be removed by allowing for externalities from investment in information processing equipment and software. Labour augmenting technical progress is estimated to be over 2 per cent per annum faster since 1995 than can otherwise be explained.

JEL Codes: O4, O3, E2

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I. Introduction and Summary

The performance of United States economy in the latter half of the 1990s was remarkable by the standards of the recent past. Annual output growth averaged more than 4 per cent per annum over 1995-2000, with the rate of labour productivity growth being more than double that observed over the previous 20 years. Inflationary pressures largely remained quiescent, even though the unemployment rate fell to its lowest level for more than 30 years. The confluence of these developments has led to suggestions that the US economy has enjoyed a fairytale ending to the twentieth century, entering into a new era of greater economic prosperity and possibility – the so-called 'Goldilocks' economy (Gordon, 1998; Driver *et al*, 2001) - helped by the significant reductions in the cost of computers, communications equipment and other new technologies in recent years (Department of Commerce, 2000). Several growth accounting studies have concluded that a combination of rapid productivity growth in producers of information technologies, along with rapid diffusion of those technologies to other industries via higher investment can account for many of the favourable developments observed in recent years (Oliner and Sichel, 2000; Jorgenson and Stiroh, 2000; Jorgenson, 2001). But others have argued that, in effect, the same old story is taking place, with much of the recent upturn in non-ICT producing industries primarily reflecting favourable cyclical influences stemming from a sequence of temporary supply shocks (Gordon, 1998 and 2000). This argument has gained credence as the marked slowdown in economic activity since the middle of 2000 has become apparent.

The purpose of this paper is to try and put the recent impact of information technologies on economic growth into a longer-run perspective by undertaking an econometric analysis of the sources of technical progress in the US business sector over the post-war period. There is no single agreed definition of what constitutes the 'New Economy'. We follow the majority of recent empirical studies by examining the impact of investments in information processing equipment and software (IPES). Computers and communications equipment are examples of general purpose technologies, with new ways of processing and managing information having many applications throughout the economy and becoming more valuable as a result of network externalities as the number of users expands.

We adopt a parametric approach, modelling jointly a pair of factor demand equations derived consistently from an underlying production technology with endogenous technical progress and non-linear cross-equation restrictions, using the modelling methodology utilised by Barrell and Pain (1997 and 1999). This permits direct tests of the elasticity of substitution, returns to scale and whether technical change is factor-biased, none of which can be tested in the alternative growth accounting approach. All of these matter, not just for decompositions of the sources of growth, but also because potentially invalid assumptions about the form of production technologies are widely employed in many other empirical exercises, such as the marginal cost based 'New Phillips Curves' derived by Gali and Gertler (1999).

The structure of this paper is as follows. Section II contains a detailed description of the modelling strategy adopted. Section III discusses data issues, including trends in the potential drivers of technical change – R&D, education and IPES investments. The main empirical results are in Section IV. Concluding comments are given in Section V.

II. The Model

There are many different ways of investigating the impact of absorbed technology on growth and of constructing estimates of potential output. De Masi (1997) and Hulten (2000) provide succinct overviews of the advantages and disadvantages of the competing approaches. Within the production function literature, a widely used approach is to look at the determinants of measures of total factor productivity (TFP), often derived from an assumed Cobb-Douglas production function after accounting for all observable factor inputs. The constructed measure of TFP can be regressed on a number of factors which are thought to determine it. Coe and Helpman (1995) provide a widely cited example of this approach, with national TFP levels related both to the domestic R&D stock and foreign R&D stocks embodied in trade.

TFP constructed in this fashion may be very different from underlying technical change (Hulten, 2000). A general problem is that the Cobb-Douglas function imposes an elasticity of substitution of unity and assumes that all technical change is Hicks-neutral. These are testable propositions. If they are invalid, then the constructed measures of TFP may be subject to bias (Nelson, 1966; Hulten, 2000). The TFP calculation also makes the assumption that firms are always on their production frontier. In practice firms face costs in adjusting their inputs, such as hiring and firing impediments and delays in ordering investment goods. In the short-term, demand fluctuations arising from the business cycle can be met by varying utilisation rates and hours worked, implying that factors such as productivity per head may well vary from time to time across industries and countries for reasons that have nothing to do with technological or organisational advances. Gordon (2000) illustrates the importance of accounting for cyclical effects when assessing the impact of new technologies over short time spans. Some studies are forced to apply employ ad-hoc adjustments to input levels to remove cyclical influences before calculating potential output from calibrated production functions (for example, CBO (1995)).

An alternative is to seek to identify endogenous determinants of technical progress within estimated dynamic factor demand equations consistent with a particular underlying production structure. This ensures consistency between the estimates of the cyclical effects and the long-run structure. In this paper we employ the methodology used by Barrell and Pain

(1997, 1999) and use factor demand equations consistent with an underlying CES production function of the form:

$$Q = \gamma\left[s\left(Ke^{\kappa t}\right)^{-\rho} + (1-s)\left(Le^{\lambda t}\right)^{-\rho}\right]^{-(v/\rho)} \qquad [1]$$

Here v denotes returns to scale, γ and s are production function scale and distribution parameters, and the elasticity of substitution (σ) is given by 1/(1+ρ). If σ=1 (ρ=0), then production is Cobb-Douglas. K and L denote the net capital stock and labour input measured in terms of employee hours. The production function allows for the possibility of both labour and capital augmenting disembodied technical progress at rates λ and κ respectively. Restrictions can be imposed to yield Hicks-neutral technical progress.

It is possible to obtain estimates of σ, λ and κ using the factor demand equations implied by the marginal productivity conditions that the marginal product of each input should equal its (mark-up adjusted) real price. Thus:

$$\frac{\beta w}{p} = \frac{\partial Q}{\partial L} = v(1-s)(\gamma)^{-\rho/v} Q^{(v+\rho)/v} L^{-(1+\rho)} (e^{\lambda})^{-\rho} \qquad [2a]$$

$$\frac{\beta c}{p} = \frac{\partial Q}{\partial K} = vs(\gamma)^{-\rho/v} Q^{(v+\rho)/v} K^{-(1+\rho)} (e^{\kappa})^{-\rho} \qquad [2b]$$

where w denotes labour compensation per employee hour, c is the nominal user cost of capital and p is the price of value-added at factor cost. These first-order conditions can be used to derive log-linear ‘desired’ labour and capital demand equations of the form:

$$\ln(L^*) = \frac{1+\sigma(v-1)}{v}\ln(Q) - \sigma\ln(w/p) - (1-\sigma)\lambda t + \omega_L \qquad [3a]$$

$$\ln(K^*) = \frac{1+\sigma(v-1)}{v}\ln(Q) - \sigma\ln(c/p) - (1-\sigma)\kappa t + \omega_K \qquad [3b]$$

where ω_L and ω_K are constants.

If it was assumed either that capital augmenting technical change is zero (κ=0), or that all technical change is Hicks-neutral (κ=λ=η, so that γ is replaced by $e^{\eta t}$ in [1]), then it would be possible to calibrate most of the production function by using the labour demand equation alone. But estimating the labour and capital demand equations jointly, with appropriate long-run cross-equation restrictions being imposed, allows these hypotheses to be tested. Earlier examples of papers which have sought to estimate the factor demand equations consistent with a CES production structure include Coen and Hickman (1970), Bergström and Melander (1979), and Barrell and Pain (1999). The latter two studies allow for endogenous technical change.[1]

The common coefficient on the real producer wage in [3a] and the real cost of capital in [3b] provides a direct point estimate of the elasticity of substitution, allowing the technical progress parameter(s) and returns to scale to be identified.[2] The equations also allow the assumption of an aggregate Cobb-Douglas representation used by the IMF, the OECD and the Congressional Budget Office (see, for example De Masi et al, 1999; Giorno et al, 1995; CBO, 1995 and 1997) to be tested. Four restrictions are required on the system of equations [3a] and [3b] to yield a relationship consistent with a Cobb-Douglas production function; a unit elasticity on output and real factor prices and a zero coefficient on the technical progress terms. Imposing these would give constant long-run factor shares.

Endogenous technical change is typically investigated either by introducing specific variables in the production function itself or by endogenising technical progress. Keller (1989) provides a comprehensive overview of the empirical literature. Here we assume that technical progress is primarily driven by the accumulation of knowledge, as captured by the stock of research and development (R) and a measure of human capital (H), two factors found to be important in many empirical studies of the growth process (Ahn and Hemmings, 2000). Thus:

$$\lambda t = \lambda_R \ln(R)_{t-i} + \lambda_H \ln(H)_{t-i} \qquad [4a]$$

$$\kappa t = \kappa_R \ln(R)_{t-i} + \kappa_H \ln(H)_{t-i} \qquad [4b]$$

This functional form implies that technical progress will grow at a constant rate if the real stock of R&D and the level of human capital grow at a constant rate. We arbitrarily assume that these variables enter with a one year lag (*i*=1). We subsequently test whether any additional factors are required in order to explain developments in the 1990s and then investigate whether they are consistent with the pattern of IPES expenditures.

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[1] Some authors have argued that price regularly exceeds marginal cost in many US industries (Hall, 1988). In this case allowance needs to be made for the potential endogeneity of the mark-up as well. This raises an interesting problem for the production function approach, since it is common to use the approach to derive an output gap for subsequent use in equations that seek to estimate the mark-up. Constant returns to scale are found in the empirical results below, providing some weak support for the hypothesis of competitive markets.

[2] Setting the ratio of the marginal products equal to the ratio of factor prices yields a further relationship which can be used to recover the distribution parameter s given the estimates of σ, κ and λ from the factor demand equations.

The parameters of the technical progress function can be estimated jointly with those of the factor demand schedules by substituting [4a] into [3a] and [4b] into [3b]. Allowing for adjustment lags, we have a pair of non-linear dynamic equations for labour and capital with the form:

$$\Delta \ln(L_t) = \beta_{L0} + \beta_{L1}\Delta \ln(Q_t) + \beta_{L2}\Delta \ln(w_t / p_t)$$
$$+ \beta_{L4}\left[\ln(L_{t-1}) - \left[\theta \ln(Q_{t-1}) - \sigma \ln(w_{t-1}/p_{t-1}) - \theta_{HL} \ln(H_{t-1}) - \theta_{RL} \ln(R_{t-1})\right]\right]$$

[5]

$$\Delta \ln(K_t) = \beta_{K0} + \beta_{K1}\Delta \ln(Q_t) + \beta_{K2}\Delta \ln(c_t / p_t) + \beta_{K3}\Delta \ln(K_{t-1})$$
$$+ \beta_{K4}\left[\ln(K_{t-1}) - \left[\theta \ln(Q_{t-1}) - \sigma \ln(c_{t-1}/p_{t-1}) - \theta_{HK} \ln(H_{t-1}) - \theta_{RK} \ln(R_{t-1})\right]\right]$$

where $\theta_{HL}=(1-\sigma)\lambda_H$, $\theta_{RL}=(1-\sigma)\lambda_R$, $\theta_{HK}=(1-\sigma)\kappa_H$ and $\theta_{RK}=(1-\sigma)\kappa_R$. Failure to allow for any cyclical effects would imply the strong assumption that companies always use the minimum inputs necessary to produce a given level of output.

Estimation is on a consistent, annual time series data set over the period 1950-2000. In all cases current values of output and real factor prices are treated as endogenous, with parameter estimates being obtained using three-stage least squares (3SLS) with appropriate cross-equation restrictions imposed. Terms in lagged real wage inflation, the cost of capital, unemployment, output and capacity utilisation were used as additional instruments.

Some care is needed in testing for externalities from measures of human capital. It could be argued that the measure of labour used in the production function and in the real wage should be measured in efficiency units. In this case education should only enter the technical progress function if there are external economies from the existence of the stock of human capital over and above the implicit improvement in labour quality. There are two ways in which the hypothesis of externalities from human capital can be tested. In the model set out above, in which there is no direct allowance for improvements in labour quality, we can write desired labour demand as:

$$\ln(L^*) = -\sigma \ln(w/p) - (1-\sigma)\lambda_H \ln(H) + \ldots\ldots \quad [6a]$$

Alternatively, allowing for quality-adjusted labour:

$$\ln(LH^*) = -\sigma \ln(w/pH) - (1-\sigma)\lambda_{H1} \ln(H) + \ldots\ldots \quad [6b]$$

which can be rewritten as:

$$\ln(L^{*}) = -\sigma \ln(w/p) - (1-\sigma)[1+\lambda_{H1}]\ln(H) + \ldots\ldots \qquad [6c]$$

So, if λ_H=1 in [6a] we can reject the hypothesis that there are significant externalities from human capital (λ_{H1}=0) , although improvements in human capital would still of course raise output by raising the quality of labour. Alternatively if λ_H is significantly greater than unity in [8a], then this would constitute evidence of externalities from human capital.

III. Data Issues

The primary focus of this paper is on developments in the non-housing business sector, using an annual data set covering 1947-2000. The business sector is defined as GDP less the gross product of general government and gross housing product. It makes little sense to include government output or the flow of services from owner-occupied housing in any analysis of productivity trends, or indeed measures of the output gap. Government output is measured on the basis of inputs, so that there is in effect no productivity growth, whilst housing services are measured by applying a rate of return to the housing stock.[3]

Labour input is measured in terms of annual hours worked, with hours worked per employee derived from Bureau of Labor Statistics (BLS) data. Over our sample period labour productivity, as shown in Chart 1, reflects a familiar pattern of rapid growth in the so-called 'Golden Age' up to the early 1970s, followed by a marked slowdown before some improvements in the 1990s. The strong upturn in productivity growth since the mid-1990s is readily apparent in both the annual data and the moving 5 year trend, but is much less marked in the 10 year trend, indicating the relatively short interval over which the upturn has occurred. However the growth of recent years is not unprecedented, with the long-term trends revealing the sustained period of productivity growth prior to 1973.

We initially measure capital inputs using the real net stock of private sector non-residential fixed assets, as published by the Bureau of Economic Analysis (BEA). Whilst capital stocks are typically used in production functions, it should be noted that there are alternative ways of measuring capital inputs. The BEA estimates are a 'wealth measure', with each component of the aggregate stock weighted by its current asset price. Growth accounting studies, such as Oliner and Sichel (2000) and Jorgenson and Stiroh (2000) employ a 'productive measure' of capital services, in which each asset is weighted by its marginal revenue product, or rental price. If the flows of capital services were always proportional to the capital stock then this distinction would not matter all that much (Hulten, 2000). But at times of rapid change in the

[3] We use a 'product side' measure of output. Alternatively, as Oliner and Sichel (2000) note, output could be measured from the 'income side'. A gap between the two has emerged in recent years, and productivity growth measured from the income side would be a little faster than the series used here. However there is no means of judging which, if either, of the two measures is correct.

asset composition of the total stock it can be more important.[4] We report results using both measures of capital inputs.

The Cost of Capital

The cost of capital used in the empirical work is based on the Hall-Jorgenson rental rate formula for the cost of a unit of capital services. Full details of the data sources used to construct a measure for the business sector are given in Pain (2001). It is shown there that the average level of the user cost of capital in the 1990s is around 1 percentage point lower than the average over the previous forty years. Other things being equal, this is likely to have acted to stimulate the strong growth of fixed investment seen during the last decade. A long-term downward trend in the tax wedge, in conjunction with the fall in the relative price of capital goods are the two factors responsible for the decline in the constructed user cost, more than offsetting the recent acceleration in the rate of increase of capital consumption.

R&D and Education

To model technical progress, measures of the stock of R&D and human capital are required. Although annual data on the level of R&D expenditure in the US are readily available, there are no estimates of the stock published regularly. A benchmark stock estimate at 1996 prices in 1959 was constructed by multiplying the current estimate of GDP at 1996 prices in that year by the stock-GDP ratio at 1987 prices from Carson *et al* (1994) This was then updated using a perpetual inventory formula, using annual data on the economy-wide flow of expenditure deflated using the implicit price deflator for GDP from the NIPA and an annual depreciation rate of 11 per cent, again following Carson *et al* (1994). The constructed R&D stock is estimated to have risen by an average 4.6 per cent per annum over 1950-2000, compared to growth of 3½ per cent per annum in output. The rate of growth of the stock was especially rapid in the late 1950s and in the 1960s, before subsequently slowing during the 1970s. It has accelerated consistently each year since 1994, but remains well below the levels seen in the 1960s (Pain, 2001).

The business sector component of total expenditure on R&D will already be included in the data on capital and labour inputs, but the public sector component will not be. If the social rates of return from R&D are equal to the private rates of return, there will not be a separate effect of R&D on innovation (λ_R and κ_R would be insignificant). Hence significant R&D effects would provide evidence of externalities from innovation.

[4] Assets such as computers and other IT equipment with high rates of depreciation and declining nominal prices will have high rental prices because of their high marginal products. A measure using rental price weights will give higher weight to these assets than a measure using asset price weights. Jorgenson and Stiroh (2000) provide an extended discussion.

The empirical literature on the sources of economic growth at the macroeconomic level typically seeks to include terms in the stock of human capital. However there is little agreement on how this should be measured. Two widely used approaches are to use either indicators of educational attainment based on years of completed schooling or wage-based estimates of the relative skill levels of employees. Both sets of measures are explored in Pain (2001). In this paper we concentrate on the measure of educational attainment, partly to keep matters to a manageable length, but also because use of the alternative human capital index does not change the basic conclusions, although the point estimates of the long-run parameters do differ slightly.

Data on the educational attainment of males and females in the private business sector between 1948-98 were taken from BLS (1999 and 2000), and extended up to 1999 using information on the trends in the educational attainment of the total adult population in the *Digest of Education Statistics*. These provide clear evidence of an improvement in educational attainment over time, with the fastest rate of growth in the median number of years completed occurring between 1950 and the late 1960s. The constructed measures shows mean years of education per person rising by an average 0.58 per cent per annum between 1948 and 1998. As with the R&D stock, the rate of growth in educational attainment was slower during the 1990s than in the post-war period as a whole.

IPES Investment

There is no single agreed way of measuring the New Economy, although computers, software and communications equipment all clearly matter. Further complications arise from the need to find a series that is available throughout the post-war period. The measure used here is defined as the ratio of gross fixed investment in information processing equipment and software (IPES) to business sector output. IPES incorporates four types of asset: computers and related equipment, software, communications equipment and other IPES equipment. The latter includes scientific equipment as well as photocopying, office and accounting equipment. Expenditure on computers and software is included in the NIPA only after 1959. However the small level of expenditure at that time means that their omission prior to then from the IPES data is unlikely to have much effect.[5] In 1999 expenditure on computers and software represented approximately 75 per cent of IPES.

Other Measurement Issues

One further issue which affects the precise form of the estimation results lies in the adjustments necessary to accommodate discontinuities in the measurement of real output over

[5] In 1960 the nominal dollar value of expenditure on computers and software represented approximately 6 per cent of IPES and 0.06 per cent of GDP.

the sample period resulting from changes in the way in which real outputs and prices are measured in the US National Income and Product Accounts (NIPA). This is discussed in detail in Pain (2001). Tests reported there establish that there is a statistically significant structural break between measured output and measured labour and capital inputs from 1978. The most likely explanation for this marked structural break lies in the method of calculation of the private consumption deflator since the comprehensive benchmark revisions made to the NIPA in October 1999.[6] In the empirical work we allow for a potential change in returns to scale and the elasticity of substitution after 1978, and include a full sample deterministic time trend and a time trend which begins in 1978 in both the labour and the capital demand equations. The imposition of equal and opposite coefficients on the two trend measures could not be rejected using a quasi-likelihood ratio test [QLR(2)=0.52]. Imposition of this restriction means that there is a trend effect up until 1978, but not after that point.

The restricted trend terms are included in all the regressions reported below. Their principal effect is to raise the growth of capital productivity after 1978 (see Pain, 2001). Although this obviously affects the other coefficients, it does not appear to alter the conclusions drawn from any of the reported statistical tests. The output elasticity is allowed to shift as from 1978, but the elasticity of substitution is not, as a separate parameter for 1978-2000 was insignificant [QLR(1)=0.93].

IV. Empirical Results

Is There A Structural Break In The 1990s?

The hypothesis of a structural break in the 1990s was tested by estimating jointly the pair of factor demand equations [5] augmented by the additional terms to reflect the data change in 1978 (denoted TT-TT78 in Table 2 below) and a systems analogue of the Salkever (1976) dummy variable test. An unrestricted model was estimated over a sample from 1950 to year t, including separate (1,0) dummies for each year from 1990 to t in both factor demand equations and the instrument set. A quasi-likelihood ratio test was then used to test the joint significance of the dummies.

The results are summarised in Table 1. There was no evidence of significant instability over 1990-94, a breakpoint chosen because 1995 (or 1996) appear to be widely viewed as the year in which the New Economy started to exert an important influence. But for 1990-2000, there is significant evidence of joint instability, and as expected this is particularly marked when

[6] The price index is constructed as a geometric mean of its components as from 1978, and in turn this resulted in large upward revisions to the estimated rate of growth in the volume of consumers' expenditure and GDP relative to employment and investment in fixed capital after that date. Oliner and Sichel (2000) note a similar problem.

the joint significance of the dummies for 1995-2000 is tested. There was a marked distinction between the labour and capital demand equations, with the dummies in the capital demand equation being jointly insignificant, but the dummies in the labour demand equation being jointly significant. This suggests that there may have been some factor bias in whatever lies behind the reported instability.

The coefficients on the individual dummies in the labour demand equation were all found to be negative, with the absolute value increasing over time. A test indicated that the restrictions required to parameterise the dummies as a single time trend could not be rejected [QLR(5)=7.39]. This is consistent with the New Economy hypothesis, in that it suggests that the level of labour augmenting technical progress has become ever higher than might otherwise have been expected since the mid-1990s given the rate of accumulation of knowledge through R&D and education.

Sources of Technical Change

Having found evidence of structural change the original labour-augmenting technical progress function [4a] was re-specified to include an additional term λ_TT95 where T95 denotes a time trend beginning in 1995. The initial unrestricted empirical results are shown as equation 1 of Table 2. The parameters on the two equilibrium-correction terms (β_{L4} and β_{K4} in [5]) are both well-determined, providing some evidence that statistically valid long-run relationships have been found. As Gordon (2000) notes, actual productivity is clearly pro-cyclical, other things being equal, with the growth in hours worked and the capital stock lagging behind the growth in output.[7] Constant returns to scale could be validly imposed after 1978 [QLR(1)=0.29], but there was marginal evidence of increasing returns to scale prior to then. There are significant effects from the stock of R&D on both labour and capital augmenting technical progress. Improvements in educational attainment appear to raise the level of output that can be produced by a given labour input, confirming that education does indeed raise labour quality, but do not appear to offer significant externalities either for labour or capital, although the point estimate of the long-run parameter (λ_H) is above unity. There is clear evidence that the rate of growth of labour-augmenting technical progress was significantly faster from 1995-2000, with the long-run parameter (λ_T) implying that on average technical progress rose by 2.3 per cent per annum between 1995-2000 over and above the rate implied by the rate of knowledge accumulation as captured by the R&D stock and educational attainment.

[7] For example, in the first year a 1 per cent rise in output is associated with a 0.26 per cent rise in output per employee hour, other things being equal.

A variety of possible restrictions on the technical progress parameters can be tested. The hypothesis that all technical progress is solely labour-augmenting was rejected [p-value 0.000], as was the joint hypothesis of neutral effects from the R&D stock and education [p-value 0.028]. Breaking down the later finding indicated that the hypothesis that R&D has a neutral effect could not be rejected [p-value 0.680], but the hypothesis of common effects from education could be [p-value 0.018]. As might be expected given the results reported in equation 1, the hypothesis that education improvements do not affect capital-augmenting technical progress could not be rejected [p-value 0.999]. Imposing the joint restrictions of common long-run R&D parameters (λ_R=κ_R) and the absence of externalities from educational attainment (λ_H=1 and κ_H=0) gave equation 2 in Table 2. These restrictions on technical progress were readily accepted by the data [p-value 0.917]. The impact of them is to raise the point estimate of the impact of R&D and the extent of exogenous labour-augmenting technical progress since 1995. This is now put at 2.5 per cent per annum.

A further point of interest about all the results reported in Table 2 is that they decisively reject the hypothesis that the aggregate business sector in the US can readily be represented by a Cobb-Douglas production function. The point estimate of the elasticity of substitution is just under 0.3, close to that found for West Germany in a related model (Barrell and Pain, 1999), and is clearly significantly different from unity. This suggests that that use of a Cobb-Douglas function may be unduly restrictive and generate biased estimates of TFP and marginal costs.

Tests for serial correlation were also undertaken in order to check the statistical adequacy of the reported factor demand equations. There was no evidence of significant first-order autocorrelation in any of the estimated equations. For example, the test statistic for equation 2 in Table 2 was [QLR(4)=4.66].[8]

Equation 3 in Table 2 re-estimates equation 2 using an index of capital services in the business sector produced by the BLS instead of the fixed capital stock. The index is obtained by Tornqvist aggregation of capital stocks using estimated rental prices for each asset type. The model is now estimated only up to 1998, reflecting the unavailability of the BLS series beyond this point at the time of estimation.

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[8] Letting **Y** and **X** denote vectors of dependent and explanatory variables, the presence of serial correlation of order φ in a simultaneous model implies that $\mathbf{AY} + \mathbf{BZ} + \mathbf{R}_\varphi\mathbf{U}_{-\varphi}=\mathbf{E}$, where $\mathbf{U}_{-\varphi}$ denotes the matrix of lagged residuals from the system estimates and **A**,**B** and **R** are matrices of parameters. **E** is a covariance matrix with zero off-diagonal elements, see Godfrey (1988) for further details. The lagged residuals are included in the instrument set. A quasi-likelihood ratio test of R=0 is a valid test for the absence of serial correlation.

There are some interesting differences and similarities. Changing the definition of capital inputs does not cause the previous estimates to fall apart. There are still well-determined coefficients on the equilibrium-correction terms in both factor demand equations. The R&D stock remains significant in the labour demand equation and there continues to be evidence of a significant structural break associated with an acceleration of labour-augmenting technical progress from 1995. However the R&D stock is now no longer significant in the capital demand equation. Thus an alternative explanation for its previous significance was not that technical change from R&D is Hicks-neutral, but simply that it was picking up an underlying improvement in capital quality now reflected in the capital services index.

The findings from the production function approach appear to be consistent with, and offer empirical support to, the assumptions employed in detailed growth accounting studies such as Jorgenson and Stiroh (2000). Both approaches suggest that a considerable proportion of what is typically included in TFP can be explained by improvements in factor quality. However it is described, accumulation of knowledge is raising output for a given level of physical inputs.

A further point of interest concerns the change in the coefficient on the contemporaneous dynamic term in the cost of capital when capital services are used. The significant negative coefficient found when using the capital stock becomes a significant positive coefficient when using capital services. The most likely explanation for this is that the current user cost, or rental prices, has been used to construct the index of capital services.

Has IPES Expenditure Affected Technical Progress?

The results to date indicate that there is evidence of a significant break in the sources of technical progress dated around 1995. To test whether measures of the New Economy are important determinants of technical progress, the technical progress functions [4a] and [4b] were respecified to include additional terms λ_{IPES}(IPES/Y) and κ_{IPES}(IPES/Y) respectively, where IPES/Y denotes the ratio of IPES investment to business sector output.[9]

It continued to be possible to impose the joint restrictions that R&D has a neutral effect on technical progress and that there are no significant externalities from education on either labour or capital. The resulting coefficients are reported in equation 4 of Table 2, using the capital stock data. IPES investment appears to have a significant effect on labour-augmenting technical progress, but does not appear to have any effect on capital augmenting technical

[9] Although it may seem like double counting to add investment back into the capital demand equation, this has to be done if the hypothesis of neutral technical progress is to be tested. It is frequently argued that computers have helped to generate general improvements in the organisational efficiency of many businesses (Brynjolfsson and Hitt, 2000), and if this is the case, then there may be common effects on labour and capital augmenting technical progress.

progress [QLR(1)=0.02]. The cross-equation restriction required to yield a neutral effect is strongly rejected by the data [QLR(1)=18.93].

Imposing the restriction that IPES investment affects only labour-augmenting technical progress yields equation 5 of Table 2. A rise of 1 percentage point in the investment-output ratio is estimated to raise the level of labour-augmenting technical progress by 2.54 per cent (standard error 0.46 per cent). A rise of 1 per cent in the stock of R&D is estimated to raise both capital and labour-augmenting technical progress by 0.26 per cent, a little below the coefficients obtained from equation 2.

To test whether the inclusion of IPES investments can account for the structural instability initially found for the period 1995-2000, the model was re-estimated with separate dummies for each year from 1995 to 2000 included in both factor demand equations. In contrast to the original findings, it was now possible to reject the hypothesis of structural instability over this period, as the dummies were jointly insignificant [QLR(12)=16.72 *[p-value 0.1601]*]. Thus the evidence from the production function approach appears consistent with the hypothesis that New Economy measures have an important role to play in any explanation of the surge in labour productivity in the latter half of the 1990s. If IPES investment is not included, then there is evidence of structural instability. If it is included, this evidence becomes insignificant. A further point of interest is that it appears that the impact of information technologies has had a marked factor bias; it would be misleading to model it simply as neutral technical progress.

Between 1989 and 1999 the share of IPES expenditure in business output rose by approximately 5 percentage points. If maintained at this level, the results imply that this will eventually raise the level of labour-augmenting technical progress by 12¾ per cent and hence, ceteris paribus, the level of output by close to 8¼ per cent. These contributions appear comparable to the estimates from growth accounting studies. For example Jorgenson (2001, Table 6) suggests that information technology contributed 0.57 per cent per annum to GDP growth over 1990-95 and 1.18 per cent per annum over 1995-99, implying a cumulative effect on the level of GDP of just under 8 per cent.

Technical Progress, Productivity and Trend Output

In order to help draw out the implications of the production function estimates it is useful to compare actual labour and multi-factor productivity with technical progress, defined as in equation 5 of Table 2 to include quality improvements from educational attainment and R&D, the impact of IPES and the split time trend for the structural break in 1978. Under CES technology multi-factor technical change (MFTC) can be approximated by a weighted average of labour and capital augmenting technical progress:

$$\Delta \ln(MFTC_t) \;=\; \alpha_{Lt}\Delta \ln(e^{\lambda}) + (1-\alpha_{Lt})\Delta \ln(e^{\kappa}) = \;\alpha_{Lt}\Delta\lambda_t + (1-\alpha_{Lt})\Delta\kappa_t \qquad [7]$$

where α_{Lt} denotes the average share of labour compensation in business sector value added at times t and t-1.

The annual rates of growth of MFTC and labour-augmenting technical progress are shown in Chart 2. As might be hoped, technical progress always rises from one year to the next, although there was a marked slowdown in the rate of growth in the mid-1970s, and is much less volatile than actual productivity. The underlying rate of growth of multi-factor technical change is consistently lower than that of labour-augmenting change, but the pattern over time is very similar. It is interesting to note that the rate of growth at the end of the sample period (an estimated 3.3 per cent in 2000) was higher than at any other time.

These results have important policy implications. A key question in recent years has been whether favourable supply-side developments in the US economy over the second half of the 1990s are at all sustainable. The constructed estimate of technical progress suggests that it may well be, despite the economic downturn in 2001. In the 1960s and early 1970s it is clear that labour productivity was rising more rapidly than technical progress. Such a situation could not be sustained indefinitely. In contrast the upturn in labour productivity in recent years has been accompanied by an upturn in the rate of growth of technical progress. Indeed, the model including IPES investment actually implies that by 2000 technical progress was rising at a faster rate than at any time since the mid 1950s.

V. Conclusions

The rate of productivity growth observed in the United States during the late 1990s was not all that unusual by the standards of the 1950s and 1960s. What appears different is that the rapid pace of knowledge accumulation observed in those years was much less apparent during the 1990s. We model knowledge accumulation using the stock of R&D and a measure of educational attainment. Formal tests indicate the presence of a structural break in the factor demand relationships as from 1995 and we find that the rate of labour, but not capital, augmenting technical progress since then has been over 2 per cent per annum above that which could have been predicted given recent trends in R&D and educational attainment. When various indicators of investment in information processing equipment and software are included as proxies for New Economy type effects, we find that the structural instability disappears and that these measures have a significant positive effect on labour-augmenting technical progress. The resulting estimates suggest that the rate of growth of both labour-augmenting and multi-factor technical progress has been higher in recent years than at any time since the mid-1950s, suggesting that New Economy effects have played an important part in the favourable macroeconomic developments in the late 1990s.

The balance of the evidence does offer some support to the notion that there has been a fairytale ending, with the US economy crossing into a new period of higher, sustainable levels of economic activity and higher, sustainable productivity gains. But they are not the only factor and there is no guarantee that the rate of growth of potential output will continue to rise indefinitely at the rate observed in the late 1990s. A finding that there have been structural factors that have pushed the production possibility frontier outwards does not mean that the business cycle has been abolished. Ultimately what matters for future economic growth is the rate of growth of the factors that determine the rate of growth of technical progress. Evidence that indicates that particular factors have raised technical progress in the past obviously helps to support speculation that they will do so in the future, but any judgements must depend on the available models of the determinants of knowledge accumulation and forecasts about the future economic environment.

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Chart 1. Growth of Business Sector Labour Productivity 1948-2000 (per cent)

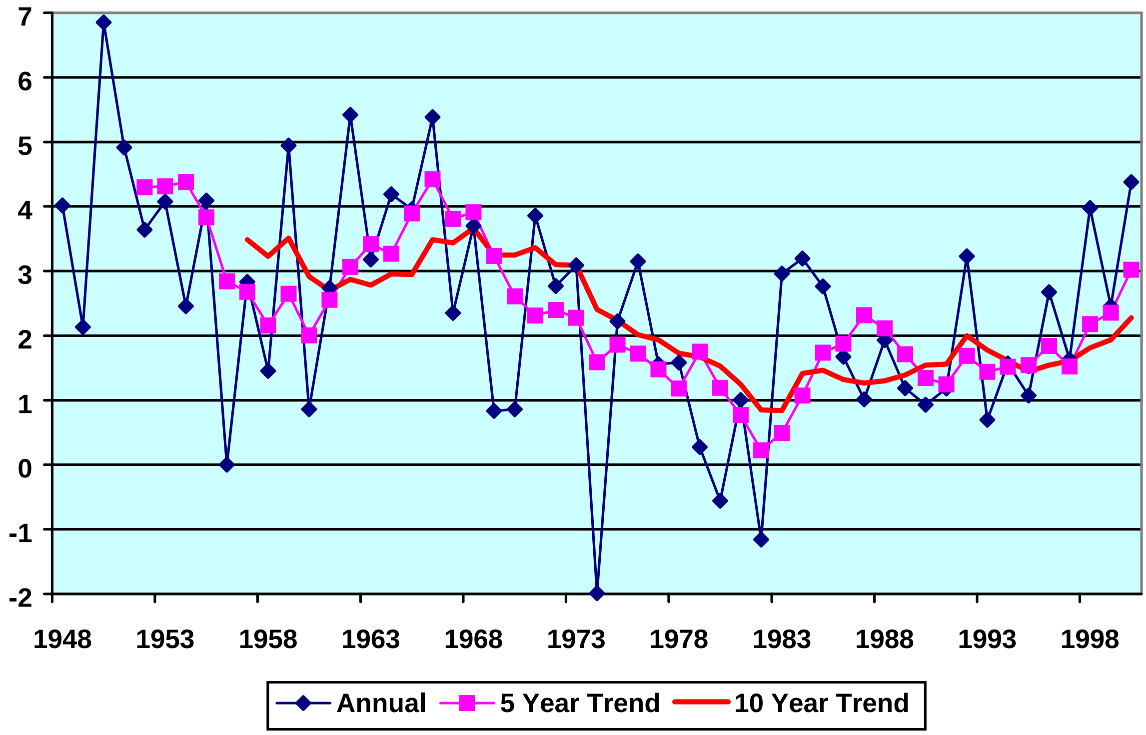

Chart 2. The Annual Growth of Labour-Augmenting and Multi-Factor Technical Progress (% log scale)

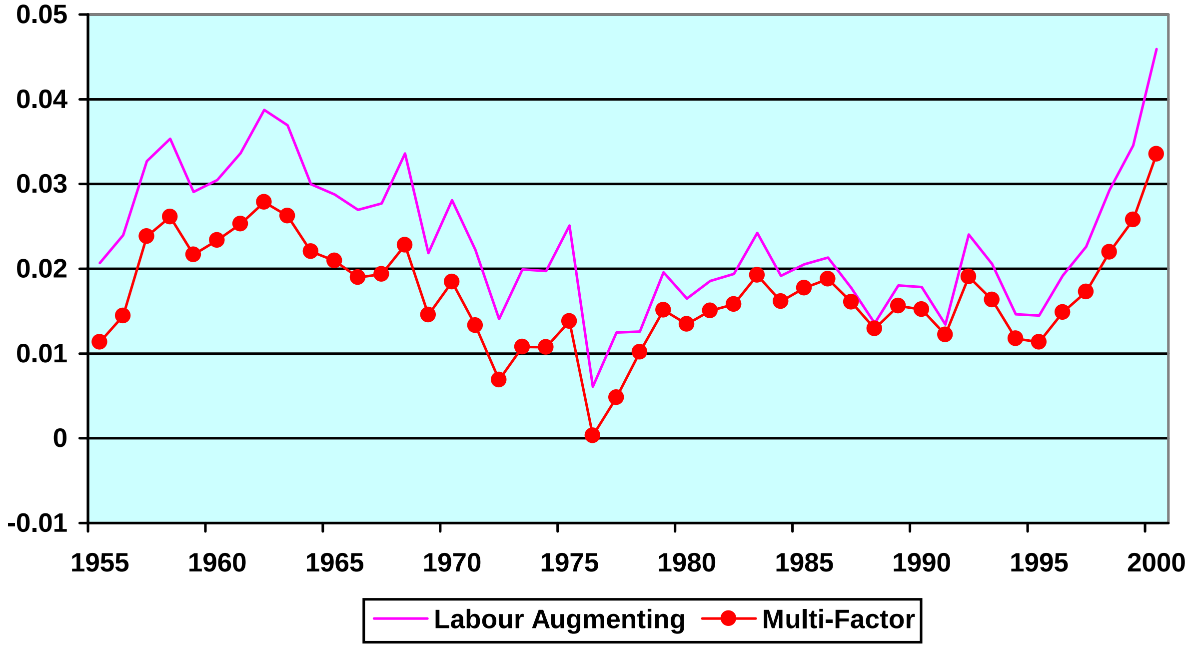

Table 1. Testing For Structural Instability In The 1990s

Estimation Period	Potential Break Period	QLR statistic p-value
1950-1994	1990-1994	0.8954
1950-2000	1990-2000	0.0440
1950-2000	1995-2000	0.0052
1950-2000	1995-2000 Labour only	0.0399
1950-2000	1995-2000 Capital only	0.4618

Note: when testing for parameter stability over 1995-2000, year dummies for 1990-94 are excluded.

Table 2. 3SLS Estimates of the Factor Demand Equations

Dependent variables: $\Delta\ln(L)$, $\Delta\ln(K)$ **Sample:** 1950-2000 except as noted in text.

		Equation 1.	Equation 2.	Equation 3	Equation 4	Equation 5
	σ	0.2473 (3.2)	0.2922 (9.4)	0.3378 (2.0)	0.2212 (4.3)	0.2223 (4.6)
	λ_R	0.3073 (5.3)	0.3136 (12.3)	0.3308 (7.2)	0.2664 (6.2)	0.2644 (7.3)
	λ_H	1.0699 (1.9)	1.0 (-)	1.0 (-)	1.0 (-)	1.0 (-)
	λ_T	0.0232 (4.6)	0.0248 (5.2)	0.0262 (3.4)		
	λ_{IPES}				0.0250 (4.3)	0.0254 (5.5)
	κ_R	0.2803 (3.6)	0.3136 (12.3)	0.0921 (0.6)	0.2664 (6.2)	0.2644 (7.3)
	κ_H	-0.012 (0.0)	-	-	-	-
	κ_{IPES}				-0.0006 (0.1)	-
	v 1978-2000 1947-1977	1.0 1.0043	1.0 1.0039	1.0 1.0034	1.0 1.0066	1.0 1.0067
	TT-TT78: Labour : Capital	0.0001 (0.1) 0.0126 (5.6)	-0.0001 (0.1) 0.0135 (11.4)	0.0009 (0.5) 0.0117 (1.8)	-0.0018 (0.8) 0.0119 (5.9)	-0.0019 (0.9) 0.0118 (7.6)
Labour Demand:	$\ln(L / L^*)_{-1}$	-0.3228 (5.5)	-0.3306 (5.4)	-0.3264 (4.3)	-0.3204 (5.0)	-0.3213 (5.1)
	$\Delta\ln(Q)$*	0.7391 (20.2)	0.7385 (20.8)	0.7433 (17.7)	0.7336 (19.3)	0.7340 (19.3)
	$\Delta\ln(w/p)$*	-0.2096 (1.6)	-0.2303 (1.9)	-0.1918 (1.2)	-0.2451 (2.1)	-0.2466 (2.1)
Capital Demand:	$\ln(K / K^*)_{-1}$	-0.0959 (8.7)	-0.0937 (8.1)	-0.0751 (4.7)	-0.0955 (7.7)	-0.0957 (7.8)
	$\Delta\ln(K)_{-1}$	0.7470 (13.7)	0.7413 (15.7)	0.5102 (5.0)	0.8037 (17.1)	0.8025 (17.0)
	$\Delta\ln(Q)$*	0.1763 (10.5)	0.1729 (11.5)	0.1171 (3.4)	0.1832 (12.3)	0.1831 (12.3)
	$\Delta\ln(c/p)$*	-0.0248 (3.7)	-0.0264 (4.1)	0.0646 (2.6)	-0.0259 (3.7)	-0.0257 (3.7)
	R^2: Labour Capital	0.868 0.928	0.869 0.927	0.870 0.747	0.867 0.927	0.868 0.927
	Standard Error: Labour : Capital	0.716% 0.234%	0.711% 0.235%	0.721% 0.514%	0.715% 0.238%	0.715% 0.238%

Notes: An * indicates a variable that is instrumented. T-statistics reported in parentheses. Constants also included but not reported here.