ECONOMETRIC ESTIMATES OF PRICES INDEXES FOR PERSONAL COMPUTERS IN THE 1990S

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ABSTRACT

In this paper we construct a number of quality-adjusted price indexes for personal computers in the U.S. marketplace over the 1989-92 time period. We generalize earlier work by incorporating simultaneously the time, age and vintage effects of computer models into a fully saturated parameterization, and then develop a corresponding specification test procedure.

While the simple arithmetic mean of prices of models by year reveals a price decline of about 11% per year, use of a matched model procedure similar to that commonly used by government statistical agencies generates a much larger rate of price decline -- about 20% per year. Since the matched model procedure holds quality constant, it ignores quality change embodied in new models. When data on new and surviving models are used in the estimation of hedonic price equations, a variety of quality-adjusted price indexes can be calculated, with varying interpretations. Although there are some differences, we find that on average these quality-adjusted price indexes decline at about 30% per year, with a particularly large price drop occurring in 1992. Parameters in hedonic price equations for desktop PC models differ from those for mobile PCs. Moreover, quality-adjusted prices fall at a slightly lower AAGR for mobile models (24%) than for desktops (32%). We conclude that taking quality changes into account has an enormous impact on the time pattern of price indexes for PCs.

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I. INTRODUCTION

The rapid technological and market developments in the personal computer

industry are widely known and have received a great deal of public attention.

One way of summarizing these major changes is to look at the unweighted

arithmetic mean of prices and technological characteristics of personal

computer models sold in the last four years in the US. Based on data as

reported by DATAPRO, in Table 1 we present such a summary.

As is seen in the final column of Table 1, mean nominal prices of PC

models dropped slightly between 1989 ($4616) and 1991 ($4320), but then fell

sharply in 1992 (to $3378), thereby changing at an overall average annual

Table 1

MEAN VALUES OF CHARACTERISTICS AND PRICES OF PC MODELS SOLD, 1989-1992

YEAR	NUMBER OF MODELS	PERCENT PROC8	PROC32	RAM IN KB	SPEED MHZ	HRDDSK IN MB	SIZE IN IN^3	WEIGHT IN LBS	AGE IN YEARS	NOMINAL PRICE
1989	225	8.0%	37.3%	1227	15.06	45.00	1695	30.43	0.467	4616
1990	575	4.7%	46.1%	1608	18.40	55.75	1638	30.71	0.473	4405
1991	453	3.1%	41.9%	1829	19.35	48.85	1469	27.87	0.547	4320
1992	630	0.0%	76.7%	4755	27.78	122.52	1419	25.60	0.654	3378
TOTAL/ MEAN	1883	3.1%	54.3%	2691	21.37	75.10	1531	28.28	0.551	4066
1982-88 MEAN	1265	32.0%	13.0%	561	8.35	17.20	NA	NA	0.600	2847

growth rate (AAGR) of -10.97%. Even as the unweighted mean of computer "box"

prices fell at about 11% per year, dramatic quality improvements occurred in the

performance and characteristics of PC models. For example in 1989, 37.3% of PC

models had a 32-bit processor, but by 1992 this had doubled to 76.7%; over the

same time period, the percent of PC models having an 8-bit processor fell from

8.0% to zero. As a result, random access memory (RAM) speed increased by about

80% from 15.06 to 27.78 MHZ. Furthermore, mean RAM capacity almost quadrupled

from 1227 kilobytes (KB) in 1989 to 4755 KB in 1992, while hard disk capacity

increased from 45 to 123 megabytes (MB). Even more dramatically, all this

increased capacity and performance was compressed into an ever smaller and

lighter "box"; while mean size fell from 1695 to 1419 cubic inches between 1989

and 1992, weight fell even more rapidly from 30.43 to 25.60 pounds.

The measurement of price indexes in the presence of such rapid

technological change is of course a difficult task. One measurement procedure

is to construct a "matched model" index, a procedure that is often used by

government statistical agencies. With a matched model index, prices of products

are compared only for models which survive unchanged from one year to the next.

Since the quality of the product is constrained to be unchanged, any price

variability represents a "pure" or "quality-constant" price change.

Using the same DATAPRO data, we have constructed a PC matched model index

for the 1989-1992 time period. A matched model is defined here in a rather

restrictive manner, having identical characteristics in the two adjacent time

periods (e.g., speed, memory, storage capacity) -- not just the same model name

and number. This limits the number óf matches. As is seen in Table 2, the

matched model index implies that prices of PC's declined at an AAGR of about

19%, with a rather steep decline of 44% in 1992.

Though the decline of this matched model price index is almost double that

of the unweighted mean price change for all models (-19.35% vs. -10.97%), it is

reasonable to expect that even the matched model procedure understates true

price declines. By examining only those models which survive unchanged year-to-

Table 2

PC PRICE INDEX BASED ON MATCHED MODEL PROCEDURE

YEAR	NUMBER OF VERSIONS SOLD	MATCHES FROM PREVIOUS YEAR	PERCENT MATCHES	MEAN PERCENT PRICE CHANGE
1989	225			
1990	575	18	8.00%	-7.860%
1991	453	185	32.17%	-0.874%
1992	630	14	3.09%	-42.561%

1992/1989 AAGR OF PRICE CHANGES: -19.348%

year, any incremental improvements in a particular model (e.g., a faster processor), disqualify it from the index. New models are ignored entirely, as of course are models that exited the market. Moreover, with the matched model procedure, since the percentage of matches in the data ranges from 3% in 1991-92 to 32% in 1990-91, a clear majority of observations are ignored. Finally, it is plausible to argue that the fewest matches occur precisely during the most active periods of product improvement -- and thus that the matched model procedure might substantially understate quality-adjusted price changes.

An increasingly useful alternative to the matched model procedure involves using econometric estimates of hedonic price equations to construct price indexes that hold quality fixed.[1] In an exploratory econometric study, Berndt-Griliches [1993] have estimated hedonic price equations based on computer magazine advertisement data for 1265 computers, beginning with the introduction of the IBM PC in 1982 and ending in 1988. The results they obtained were remarkably robust across alternative functional forms, and implied about a 28% annual decline in real quality-adjusted PC prices.[2,3]

In this paper we extend research on PC price indexes in a number of ways. First, because the data employed consists of an unbalanced panel, and because there is an identity relating the age of a model, its vintage, and the year in

which the model is observed, important issues emerge involving the interpreta-

tion of dummy variable coefficients. Some of these issues were addressed in

Berndt-Griliches. Here we extend that discussion by introducing a "saturated"

parameter model that permits a much richer parameterization, and also allows for

a more complete set of specification tests in the unbalanced panel context.

Second, in this paper we consider more completely the issue of parameter

stability and equality. This is particularly important in the PC context, where

new markets have emerged (notebooks vs. desktop models), technological change is

dramatic, and where equation specifications need to be sufficiently general to

capture possible nonlinearities. Third, this study incorporates data through

1992, and therefore updates results from previous studies. As we shall see,

1992 was a most remarkable year in terms of quality-adjusted price declines.

II.   THE DATA SETS

The data sets for this paper were obtained for the most part from Datapro

Information Services Group, a division of McGraw-Hill (DATAPRO).[4] As is seen in

Table 1, the data consists of an unbalanced panel of 1,883 observations on PC's

during the four-year 1989-1992 timeframe. To avoid the extreme end "toy" and

"file server" computer models, we have restricted the sample to models having at

least 512KB of RAM, a speed of at least 7.8MHZ (thereby still including the IBM-

XT), and having a hard disk capacity of no more than 500 MB.

In previous work, we used data from personal computer advertisements, such

as those for mail-order purchases, to obtain measures of a particular model's

characteristics and price. The advantage of that approach is that the resulting

price data more closely approximates actual transactions prices than does, say,

a list price. The disadvantage is that typical advertisements frequently

provide less than full information on the particular combination of attributes

"packaged together" in the model by the vendor. The DATAPRO data set has the

advantage of providing far more complete technical specifications than does the

typical magazine advertisement; DATAPRO provides technical information on

approximately 40 characteristics in a consistent form across manufacturers,

models and years. However, the price data from DATAPRO is for the list price of

the particular base model, rather than a transactions price. Although in our

previous study we found that "street" prices were frequently 35% lower than list

prices, whether this discount proportion has changed substantially over time is

not clear. Using 1970-86 data from the US mainframe computer market, Stephen

Oliner [1993] reports that rates of growth in list prices were very similar to

those of transactions prices. To the extent that percent discounts from list

price to transactions price have increased as competitive pressures in the PC

market have intensified in the past few years, the price indexes resulting from

our use of DATAPRO data might understate somewhat the true rate of price decline

of PC's. On the other hand, the share of sales accounted for by mail order

models has plausibly increased since 1989. Since in the mail order market there

is no apparent distinction between list and transactions prices, the use of list

price data for the entire 1989-92 time period might overstate the average

transaction price decline. Moreover, it is widely known that in 1992 major

brand manufacturers changed pricing strategies and brought list prices down

considerably to match others' transactions prices.   Which of these various

offsetting effects is dominant is, unfortunately, unknown at this point in time.

It would of course be desirable to link the 1989-92 data of this study

with the Berndt-Griliches 1982-88 data. Because the 1989-92 data employs list

price, there is no feasible link between the 1988 prices based on magazine

advertisements with the 1989 list price data from DATAPRO. A preliminary

attempt to merge the data sets resulted, not surprisingly, in very large and

unrealistic price increases from 1988 to 1989. Unfortunately, DATAPRO data

series prior to 1989 are not currently available. Consequently, the results of

this study are not directly comparable to those presented in Berndt-Griliches.

The technical characteristics data we employ from the DATAPRO data base include measures of the following variables: PRICE (in current dollars); RAM (installed random access memory capacity, in kilobytes); MAXRAM (maximum amount of addressable RAM capacity possible); MHZ (speed of processor chip, in megahertz); HRDDSK (hard disk capacity, in megabytes); SIZE (in cubic inches); WGT (in pounds); DENSITY (WGT/SIZE); AGE (age of the model, i.e. a model introduced in year t and observed in year t is 0 years old, a model introduced in year t-1 and observed in year t is 1 year old, etc). The sample is confined to models three years of age and younger. Dummy variables AGE0, AGE1, AGE2 and AGE3 are defined accordingly; DFLP23 is a dummy variable equal to one if the model has two or more floppy drives, and DNOHDDSK is a dummy equal to one if the model does not have a hard disk. Other dummy variables include DPROC (a series of dummy variables for whether the processor chip is 8-bit -- DPROC8, 16-bit -- DPROC16, or 32-bit -- DPROC32); DMOBILE (a dummy variable equal to one if the computer is defined by 'DATAPRO as notebook, laptop, portable or transportable, as distinct from desktop or floorstanding; a series of brand dummy variables equal to one if the computer brand is ALR (DALR), Apple (DAPPLE), AST (DAST), Compaq (DCMPQ), CompuAdd (DCOMPUAD), Dell (DDELL), Digital Equipment Corporation (DDEC), Hewlett Packard (DHP), IBM (DIBM), NEC (DNEC), Toshiba (DTOSH), Zenith (DZEN), or all other brands (DOTHER).[5] Finally, we have constructed dummy variables T89, T90, T91 and T92 that equal one if the computer model is being observed in that particular year, and seven vintage dummy variables (V86 through V92) corresponding to the vintage of the particular model (the year it was first introduced on the market).[6] This data is summarized in Table 3 below.

As is seen in Table 3 from means of the dummy variables, the data set is dominated by new models (AGE0 - 0.62) that are desktop or floorstanding (not laptops or notebooks, for DMOBILE - 0.23), having either zero or only one floppy disk drive (DFLP23 - 0.22) and a hard disk drive (DNOHDDSK - 0.25, implying 75%

of models have a hard disk), and by brands other than those from the nine

leading vendors (OTHER - 0.44, but with Dell, IBM, Compaq, DEC and NEC having

the largest number of models).

Table 3

SUMMARY OF DATAPRO PERSONAL COMPUTER DATA, 1989-1992

VARIABLE	MEAN	MINIMUM VALUE	MAXIMUM VALUE
PRICE	4066.31	525.00	22655.00
RAM	2690.92	512.00	32000.00
MAXRAM	21207.61	512.00	256000.00
MHZ	21.37	7.83	66.00
HRDDSK	75.10	0.00	440.00
SIZE	1530.82	34.20	7020.00
WGT	28.28	2.80	110.00

SAMPLE MEANS OF DUMMY VARIABLES

PROC8	0.0313	V89	0.1540
PROC16	0.4259	V90	0.3404
PROC32	0.5428	V91	0.2693
DFLP23	0.2103	V92	0.1620
DMOBILE	0.2268	DALR	0.0313
AGE0	0.6150	DAPPLE	0.0409
AGE1	0.2560	DAST	0.0308
AGE2	0.0924	DCOMPAQ	0.0531
AGE3	0.0366	DCOMPUAD	0.0361
T89	0.1194	DDEC	0.0531
T90	0.3054	DDELL	0.0935
T91	0.2406	DHP	0.0483
T92	0.3346	DIBM	0.0669
V86	0.0016	DNEC	0.0526
V87	0.0478	DTOSH	0.0292
V88	0.0250	DZEN	0.0228
		DOTHER	0.4414

Sample Size: 1883

III. ECONOMETRIC ISSUES

Our data set comes in the form of an unbalanced panel, in that the number

of observations by age, and by vintage, varies by year. Let TIME be a variable

equal to the year in which a model was observed (say, TIME - 1991), let VINTAGE

be a variable equal to the year in which the model was first introduced into the

market (e.g., VINTAGE - 1989), and let AGE be a variable equal to the number of

years that a particular model has been in the market (e.g., AGE - 2). For any

model-observation, we therefore have the identity

$$\text{TIME} = \text{VINTAGE} + \text{AGE}. \tag{1}$$

Because of this identity, one cannot include continuous linear versions of

all three of the TIME, VINTAGE and AGE variables into a regression equation,

else exact collinearity would occur; only two of the three could be included

directly, and indirect estimates for the impact of the third variable could be

computed using Eq. (1). Note that regardless of which two of the three

variables were included in the estimated regression equation, direct and

implicit parameter estimates would be identical, as would be the equation R^2.

Following Fienberg and Mason [1985], one could specify a model in which

two of the three variables were included linearly, and nonlinear transformations

of all three of the variables were also included; for example, all three squared

variables could be included, in addition to linear forms of two of the three

variables.[7] It is worth noting, incidentally, that this type of identity is not

unique to the hedonic price literature; a similar identity involving cohort, age

and time measures occurs in, for example, economic and sociological studies of

wage rates in the labor market.[8]

An alternative procedure for capturing nonlinearities is to specify the

time, age and vintage variables in dummy variable form. Recall that we have

constructed four time dummy variables (T89 through T92), seven vintage dummy

variables (V86 through V92), and four age dummy variables (AGE0 through AGE3).

As was noted by Berndt-Griliches, when the coefficients on these variables are

held constant over time and/or vintage, then the simple adding-up conditions

implied by Eq. (1) no longer hold. This raises significant issues concerning

the interpretation of dummy variable coefficients, as well as the maximal

parameterization possible that avoids exact collinearity.

Consider a traditional hedonic price equation relating $\ln P_{ivat}$, the

natural logarithm of the price of the ith model of vintage v, age a and time t,

to a constant term, a vector of time dummies T (T - T90, T91, T92, with T89

excluded), a vector of age dummies A (A - AGE1, AGE2 and AGE3, with AGE0

excluded), and a number of model i-specific characteristic variables (called

X's), as well as a random disturbance term u_{ivat}:

$$\ln P_{ivat} - \alpha + T'\alpha_t + A'\alpha_a + X'\beta + u_{ivat}. \qquad (2)$$

As Berndt-Griliches noted, in this T-A specification (with the V vintage dummy

variables excluded), one can interpret estimates of the α_t as changes in the

quality-adjusted price index relative to 1989, holding age fixed; hence the α_t

parameters represent the general movement in average PC prices, given the

average rate at which selectivity occurs in the sample. Similarly, estimates of

the α_a can be interpreted as the effects of age (relative to an AGE0 model) on

price, holding time fixed. The expected sign of the α_a coefficients is not

clear, however. One line of reasoning suggests that new models of superior

quality should command a premium price, and if market competition forces the

valuations of all characteristics of incumbent models to obsolesce at the same

rate, then prices of surviving vintages would decline appropriately, and

therefore since the time dummies capture the full price effects, one would

expect estimates of the α_a age dummy parameters to be approximately zero.

But a different line of reasoning might yield positive estimates of these

parameters. Specifically, if selectivity in the marketplace reflects survival

of the fittest, then the age coefficients may to some extent reflect the effects

of unobserved positive quality differentials among the survivors as vintages

progress. Such unmeasured quality characteristics could include, for example,

compatible software or hardware, or differential service policies or warranties.

Suppose that instead of the T-A specification, one included only the

vintage V and age A dummy variables, and excluded the T time dummy variables:

$$\ln P_{ivat} = \delta + V'\delta_v + A'\delta_a + X'\beta + u_{ivat}, \tag{3}$$

where, say, the vintage V89 dummy variable was omitted. To the extent the technical characteristics variables captured all of the quality changes embodied in models, the δ_v vintage coefficients in this V-A model would essentially be capturing the decline in prices by vintage. Since the δ_a coefficients now condition on vintage V rather than time T, the δ_a coefficients represent the average price decline of surviving models over the sample years, providing another measure of the average rate of improvement in the quality of new computers, forcing the price of incumbent models to decline.

As Berndt-Griliches noted, except for the special case when models of only one age are included (e.g., only AGE0 models), in spite of the identity in Eq. (1), estimates of, for example, α_a in the T-A specification from Eq. (2) will differ from estimates of δ_a in the V-A model of Eq. (3)! Estimates of the β's in the two equations will also differ, as will the equation R^2's.

A T-V model which excluded age dummies and only had time and vintage dummies would provide yet a third set of distinct parameter estimates and equation statistics. An easy way to see this is to note that the total number of T, A and V dummy variable coefficients estimated directly varies by model specification -- six in the T-A specification of Eq. (2), nine in the V-A model of Eq. (3), and nine in an analogous T-V specification. Also, since the dummy variables do not follow a strict linear dependence analogous to Eq. (1), parameter estimates for models that condition on time and age dummies will differ from those conditioning on vintage and age dummies.

Since the V-A and T-A models yield varying estimates, choice among them is not clear. There is ample precedent, of course, in the hedonic price literature for simply using a T-A type specification, for in this case the directly estimated time coefficients have a clear interpretation -- they can be used to

compute quality-adjusted price indexes in a very simple way. But precedent is
not a sufficient cause for determining functional form.

One way to proceed is to examine whether it is really necessary to discard
entirely the information on one of the three sets of dummy variables. Can one
not employ a specification that efficiently combines information simultaneously
from the T, A *and* V dummy variables, yet avoids exact collinearity?

In fact, as Hall [1971] has shown in a somewhat different context (a
balanced panel data set for second-hand trucks), this can be done. In the
current context, the maximal parameterization consistent with avoiding exact
collinearity is one in which five of the seven vintage dummy variables are added
to the T-A specification of Eq. (2). We write such a T-A-V specification as

$$\ln P_{ivat} = \gamma + T'\gamma_t + A'\gamma_a + V'\gamma_v + X'\beta + u_{ivat}, \qquad (4)$$

where the T vector excludes, say, T89, the A vector excludes AGE0, and the V
vector now excludes two vintages -- say, V86 and V92. To the extent that
quality improvements are not captured by the included X's but are embodied in
successive vintages, one might expect that the γ_v coefficients capture effects
of omitted variables. However, as Hall noted, estimates of the γ_v coefficients
should be interpreted as *differences* from the average rate of growth of
technical progress embodied but unobserved in pairwise comparisons of vintages.
Deleting the V86 and V92 dummy variables allows us to interpret the remaining
coefficients as measuring the acceleration or deceleration of technological
progress from the average rate between 1986 and 1992. In the current context
with V86 and V92 omitted, the coefficient on, say, the V90 dummy variable should
therefore be interpreted as the *difference* between the average 1990 effect and
the mean of the average vintage effects over the entire 1986-92 set of vintages.

This type of reasoning led Berndt-Griliches to a specification test: A
necessary condition for an hedonic price equation such as Eq. (2) to be satis-
factory empirically is that the portion of quality change not captured by the

characteristics variables should be unrelated to vintages, i.e., in a desirable

specification, the γ_v in Eq. (4) should be approximately zero. Based on this

specification test, Berndt-Griliches then proceeded to choose a preferred model.

Although this specification test has obvious utility and interpretation,

we now show that an alternative model can be developed, one that is richer in

parameters and therefore permits greater flexibility in hypothesis tests and

choice of specifications. It also permits us to interpret more clearly why the

T-A, V-A and T-V specifications generate different parameter estimates.

Specifically, since unbalanced panel data are available on seven vintages,

four age dummies, and four time dummies, one can envisage a fully saturated

hedonic price equation specification in which there is no intercept, but there

are separate coefficients for each of the sixteen time-vintage "interaction"

variables: V86T89, V87T89, V88T89, V89T89, V87T90, V88T90, V89T90, V90T90,

V88T91, V89T91, V90T91, V91T91, V89T92, V90T92, V91T92 and V92T92, where each of

these interaction dummy variables is an element-by-element product of Vv·Tt, for

v=86 through 92 and t = 89 through 92.

With this fully saturated parameter model as the most general

specification, it can easily be shown that for the T-A model to be valid, the

following restrictions must be placed on coefficients of the dummy variables:[9]

```
  T-A:   V90T90 - V89T89 = V89T90 - V88T89 = V88T90 - V87T89 = V87T90 - V86T89,
         V91T91 - V89T89 = V90T91 - V88T89 = V89T91 - V87T89 = V88T91 - V86T89,
    and  V92T92 - V89T89 = V91T92 - V88T89 = V90T92 - V87T89 = V89T92 - V86T89.
```

Note that with these nine independent restrictions, the number of time and age

dummy variable coefficients, plus the intercept term, is seven. Incidentally,

the interpretation of these restrictions is that in this model, the time effect

is the same for all ages in 1990, in 1991, and in 1992.[10]

By contrast, the V-A model of Eq. (3) is a special case of the more

general fully saturated model when six linearly independent restrictions are

placed on the interactive vintage-time dummy variables. The first three of

these restrictions result from the constraint that in the V-A model, vintage

effects are the same for all ages, while the last three reflect the constraint

that age effects are the same for all vintages. Together, the following six

restrictions are placed on coefficients of the dummy variables:

V-A: V87T88 - V90T91 = V87T89 - V90T92, V88T89 - V90T91 = V89T91 - V90T92,
       V89T90 - V90T91 = V89T91 - V90T92, as well as
       V89T90 - V89T89 = V90T91 - V90T90 = V91T92 - V91T91, and
       V89T91 - V89T89 = V90T92 - V90T90.

Finally, for the T-V model to be empirically valid, six linearly

independent restrictions must be placed on the fully saturated model, thereby

leaving it with ten coefficients involving the intercept, time dummies and

vintage dummies. The first three restrictions imply that with the T-V model,

the vintage effect is the same for all time periods, while the second three

restrictions result from the time effect being the same for all vintages. Hence

interactive dummy variable constraints must be as follows:

T-V: V89T90 - V90T90 = V89T91 - V90T91 = V89T92 - V90T92, and
       V91T91 - V90T91 = V91T92 - V90T92, as well as
       V87T90 - V87T89 = V88T90 - V88T89 = V89T90 - V89T89, and
       V88T91 - V88T89 = V89T91 - V89T91 - V89T89.

Hence, roughly speaking, the saturated model places no restrictions on the

acceleration or deceleration of vintage and age effects over time, the T-A-V

model of Hall places parameter restrictions on this acceleration or

deceleration, while the T-A model constrains this deceleration or acceleration

of vintage effects to be zero. Specifically, relative to the fully saturated

model, the T-A-V model involves two constraints on the age effects,

T-A-V: (V90T92 - V91T92) - (V90T91 - V91T91) =
       (V88T90 - V89T90) - (V88T90 - V89T89) = α_2 - $2\alpha_1$, and
       (V89T92 - V90T92) - (V89T91 - V90T91) =
       (V87T90 - V88T90) - (V87T89 - V88T89) = α_3 - $2\alpha_2$ + α_1

where the α's are coefficients on age dummies, as well as two restrictions on

the vintage effects over time,

       (V91T92 - V90T91) - (V90T92 - V89T91) =
       (V91T91 - V90T90) - (V90T91 - V89T90) = α_{91} - $2\alpha_{90}$ + α_{89}, and

$$(V90T92 - V89T91) - (V89T92 - V88T91) -$$
$$(V90T91 - V89T90) - (V89T91 - V88T90) = \alpha_{90} - 2\alpha_{89} + \alpha_{88},$$

where the α's are coefficients on vintage dummies.

The existence of this more general specification suggests a somewhat different specification test procedure than was used in Berndt-Griliches. First, we undertake year-by-year regressions; since these regressions involve only one time period, the age and vintage dummies are perfectly colinear, and thus we can estimate these equations with either (but not both) the age or vintage dummies. Second, we test for parameter stability by pooling the 1989-92 data and estimating the fully saturated time-vintage interaction specification described above. If these parameter stability restrictions are satisfied empirically, then we will estimate and test for restrictions incorporated in the various T-A-V, T-A, V-A and T-V specifications. As a preferred specification, we will choose that model consistent with the data but most parsimonious in parameterization; we will also examine whether there is empirical support for the traditional T-A model. Third, if the above test for 1989-92 parameter stability is rejected, we will test for a weaker type of parameter stability over time by undertaking adjacent (two-) year regressions, and comparing them with the year-by-year regressions. If the parameter constraints from adjacent year regressions are not rejected, we will compare the corresponding fully saturated, T-A-V, T-A, V-A and T-V models and choose a preferred specification. If, however, the parameter constraints of these adjacent year models are rejected, we will choose as the preferred model the year-by-year regressions.

The above tests can be carried out using the standard F-test methodology. Among others, however, Arrow [1960] and Ohta-Griliches [1976] have noted that when samples are large and standard test procedures are employed, one is likely to reject most simplifying parameter restrictions on purely statistical grounds, even though they may still serve as adequate approximations for the purpose at hand. There are several ways one can deal with this problem.

First, to accommodate the larger sample size, we can compensate by choosing very tight significance levels for the standard F-tests. We do that here by choosing 0.01 significance levels. Second, since in our hedonic regressions the dependent variable is the natural logarithm of price, the root mean squared error (RMSE) measures the unexplained variation in prices in, roughly, percentage units. A reasonable criterion is to employ the difference in the RMSE of the constrained and unconstrained regressions as a relevant measure of the explanatory power of a particular model. As an alternative test criterion, therefore, we will reject the null hypothesis when the RMSE under the alternative results in a reduction of more than 5% in the RMSE (the standard deviation of the unexplained variation in log prices). With a typical RMSE of around 0.40, this RMSE criterion implies that we are looking for a movement of at least about 0.02, say, from 0.40 to 0.38, before we will "give up" on the more parsimonious parameterization corresponding to the null hypothesis.

Two other comments concerning hypothesis testing are appropriate here: (a) Since heteroskedasticity may be present, we will compute standard errors and variance-covariance matrices of the coefficients using the White [1980] heteroskedasticity-robust procedure. Traditional F-tests involving joint hypotheses, not just t-tests, will therefore be adjusted to accommodate the heteroskedasticity-consistent variance-covariance matrix; and (b) as an additional check on the various model specifications, we will assess the sensitivity of computed price indexes to various parameter restrictions, including some of those rejected by the above test criteria. If growth in the price indexes is robust to alternative models, the empirical significance of rejections of particular models will, for our purposes, be diminished.

IV.   FUNCTIONAL FORMS, STABILITY AND RESULTS OF SPECIFICATION TESTS

We now turn to a discussion of empirical findings. Results from preliminary regressions suggested that parameters differed significantly for

mobile and desktop models, and thus we proceeded by disaggregating PC's into

these two groups.[11] We now discuss the two sets of findings separately.

IV.1 RESULTS FOR DESKTOP MODELS

We began by estimating yearly regression equations using an expanded

version of the T-A functional form specified in Berndt-Griliches, where squared

values of LRAM, LMAXRAM, LMHZ, LHDDSK, LRAM, LSIZE and LWGT (variables

identified by the suffix SQ), as well as interaction variables LRAMHARD,

LMHZSIZE, and LWGTSIZE were included as regressors.[12] We also added a dummy

variable DNOHDDSK for models having no hard disk. With our basic model refined

in this manner, we then examined parameter stability issues further.

Our first result is that for desktop models from 1989 to 1992, there is

little support for parameter stability. Comparing the year-by-year regressions

with the 1989-92 pooled fully saturated model, we reject the hypothesis of

parameter stability using both the traditional F-test (test statistic of 4.11,

.01 critical value of 1.39) and the 5% ΔRMSE criteria (ΔRMSE is 9.70%).

Moreover, this rejection of parameter stability largely remains even when one

compares year-by-year with adjacent- (two) year regressions. Specifically, for

1989-90 and 1991-92, the F-test statistics of 4.35 and 3.22 exceed the .01

critical values of 1.71 and 1.79, while for 1990-91 there is marginal support

for parameter stability (test statistic of 1.52 compared to 1.70 critical

value). Similar conclusions emerge when using the ΔRMSE criterion (the 1989-90,

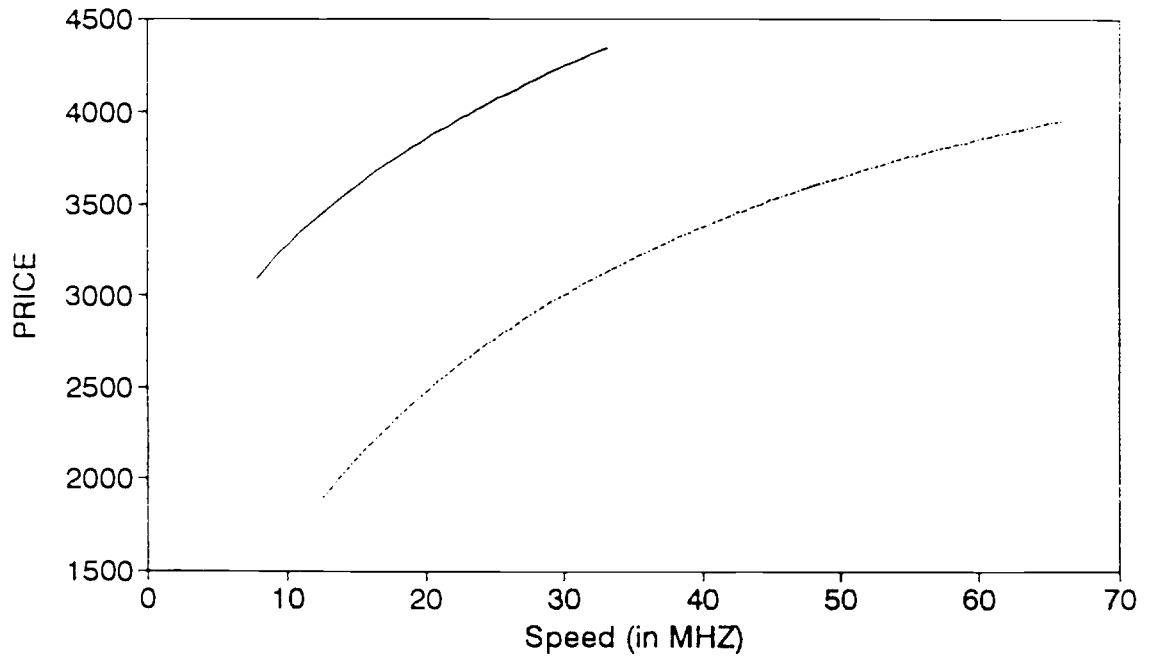90-91 and 91-92 ΔRMSE are 7.67%, 1.35%, and 5.10%, respectively).

Given this rejection of parameter stability for the desktop models and our

testing methodology, we terminate our testing procedure for further functional

form restrictions, such as those involving the T-A and other models. However,

several other results are worth noting. First, with each yearly regression,

results from Box-Cox regressions strongly supported the log-log specification.[13]

Second, as is seen in Table 4 where results from year-by-year regressions are

TABLE 4

PARAMETER ESTIMATES OF YEAR-BY-YEAR REGRESSIONS FOR DESKTOP COMPUTERS
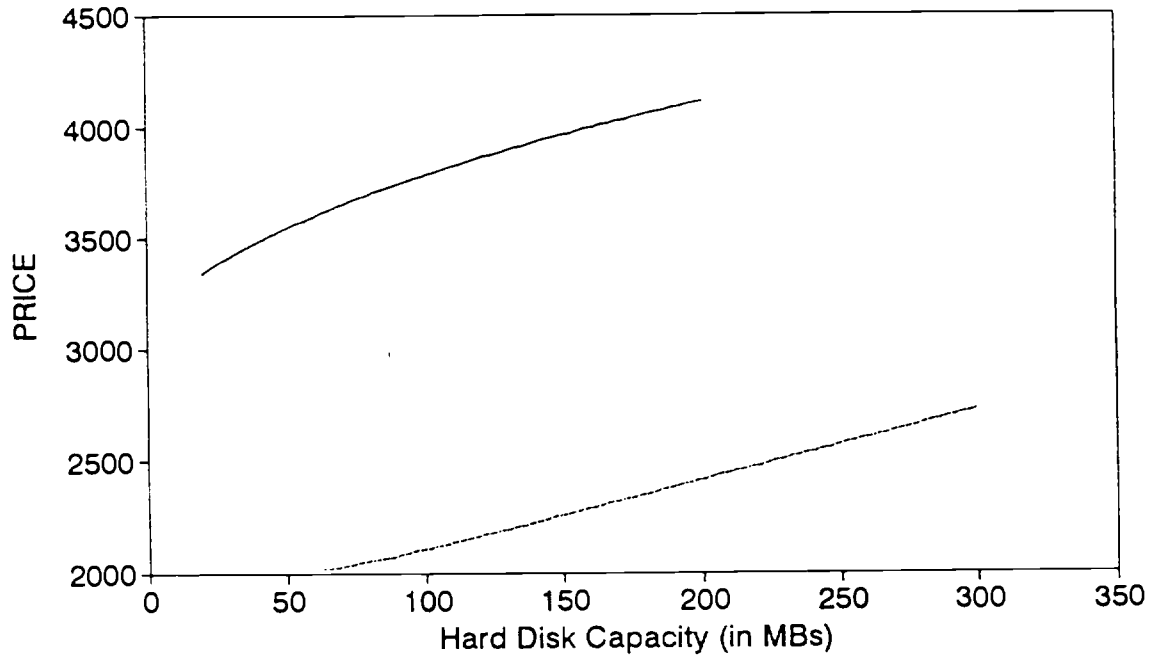(t-statistics from heteroskedasticity-robust standard errors in parentheses)

Variable	1989	1990	1991	1992	Variable	1989	1990	1991	1992
Constant	7.565 (0.73)	10.49 (4.11)	17.36 (3.72)	14.00 (4.14)	DAGE1	-0.128 (1.30)	-0.028 (0.44)	0.132 (2.37)	0.199 (5.02)
LRAM	3.832 (2.74)	-0.369 (0.71)	-0.297 (0.38)	-2.173 (4.12)	DAGE2	0.030 (0.38)	0.057 (0.82)	0.293 (2.13)	0.344 (7.62)
LRAMSQ	-0.231 (2.32)	0.051 (1.44)	0.042 (0.78)	0.157 (4.65)	DAGE3	0.043 (0.36)	0.122 (2.51)	0.489 (6.16)	0.214 (2.94)
LMAXRAM	-0.395 (1.14)	0.071 (0.26)	-1.446 (3.55)	1.145 (2.13)	DPROC8	-0.342 (2.10)	-0.019 (0.13)	-0.265 (0.98)	na
LMAXRAMSQ	0.018 (0.85)	0.003 (0.20)	0.084 (3.89)	-0.042 (1.62)	DPROC32	0.303 (2.62)	0.222 (4.16)	0.228 (2.73)	0.090 (1.94)
LMHZ	2.556 (1.18)	0.175 (0.23)	1.018 (0.94)	1.140 (1.29)	DALR	na	na	na	-0.285 (3.12)
LMHZSQ	0.001 (0.01)	0.114 (0.95)	-0.528 (3.10)	-0.108 (1.26)	DAPPLE	0.514 (3.84)	0.627 (8.25)	0.446 (3.21)	0.279 (2.53)
LHDDSK	0.017 (0.04)	0.205 (0.82)	-0.130 (0.33)	-0.185 (0.69)	DAST	0.551 (4.10)	-0.455 (4.04)	-0.230 (1.41)	0.021 (0.26)
LHDDSKSQ	0.019 (0.36)	0.051 (2.00)	0.060 (1.25)	0.092 (3.90)	DCMPQ	0.589 (5.33)	0.465 (7.72)	0.592 (6.05)	0.114 (1.17)
LRAMHARD	-0.012 (0.26)	-0.064 (4.76)	-0.030 (1.42)	-0.062 (4.32)	DCOMPUAD	na	na	na	-0.295 (3.81)
LSIZE	-5.790 (1.44)	-1.181 (2.05)	-1.663 (1.47)	-0.945 (1.98)	DDEC	na	na	na	0.142 (2.50)
LSIZESQ	0.503 (1.30)	0.072 (1.26)	0.009 (0.08)	-0.120 (2.87)	DDELL	0.761 (6.46)	0.310 (4.11)	0.385 (4.88)	0.033 (0.40)
LWGT	1.896 (0.72)	-0.516 (0.78)	0.152 (0.16)	-2.747 (4.42)	DHP	0.349 (3.10)	0.395 (6.97)	0.521 (7.92)	0.282 (4.00)
LWGTSQ	-0.166 (0.66)	-0.148 (1.10)	-0.305 (1.40)	-0.444 (2.24)	DIBM	0.344 (3.21)	0.475 (8.60)	0.509 (6.28)	0.021 (0.24)
LMHZSIZE	-0.311 (1.24)	-0.059 (0.63)	0.351 (2.20)	0.004 (0.04)	DNEC	0.464 (5.67)	0.086 (0.85)	-0.061 (0.35)	-0.012 (0.19)
LWGTSIZE	-0.118 (0.22)	0.191 (1.34)	0.227 (0.95)	0.808 (3.43)	DTOSH	na	na	na	1.042 (7.82)

Figure 1
Desktop Computer Price vs Speed

Figure 2
Desktop Computer Price vs Hard Disk

DFLP23	0.037	-0.315	-0.242	-0.040	DZEN	na	0.259	0.350	-0.215
	(0.36)	(5.17)	(3.77)	(0.53)			(3.27)	(4.62)	(3.61)
DNOHDDSK	-0.643	-0.471	-0.498	-1.373	R^2	0.832	0.830	0.769	0.824
	(0.71)	(0.94)	(0.61)	(2.42)	RMSE	0.341	0.306	0.370	0.248
					N	173	457	321	504

presented, a number of reasons underly the rejection of parameter stability. In 1989 and 1992, for example, the coefficients on LRAM and LRAMSQ are statistic-ally significant, but are reversed in sign; similar significant sign reversals occur in 1991 and 1992 for LMAXRAM and LMAXRAMSQ. Trends in other coefficients are also evident, such as those for LHDDSKSQ, LRAMHARD, LWGT and LWGTSQ. Of special interest are coefficients on several premium brand name dummy variables. For IBM, DELL, CMPQ, APPLE and ZEN, the 1992 estimates indicate sharp declines from earlier years, corroborating the highly publicized aggressive changes in pricing strategy adopted by these companies in 1992 to meet competitive market pressures. In fact, by summing the company and company-1992 coefficients, we find that in 1992, quality-adjusted list prices for IBM, DELL, and CMPQ were not significantly different from the "OTHER" brand names, ZEN had become somewhat cheaper, and while APPLE still commanded a premium price, this premium had been approximately halved. Third, even though parameters vary considerably by year, the yearly estimates are plausible. In Figure 1, for example, we plot the fitted price of desktop PC models as a function of speed (in MHZ) for 1989 and for 1992, where all other variables are measured at their yearly sample means and MHZ is plotted over the observed range of values for that year. As is seen in Figure 1, the price frontier is upward sloping, but has fallen sharply between 1989 and 1992. A similar pattern is displayed in Figure 2, where desktop computer price is plotted as a function of hard disk capacity.

IV.2 RESULTS FOR MOBILE MODELS

The results we obtained from regressions with mobile (laptop, notebook, and transportable) models differed considerably from those of the desktop models. First, the considerable nonlinearities among variables such as MHZ, RAM, HDDSK, and MAXRAM were not significant in the mobile model regressions. This led us to stay with a more parsimonious and linear specification. Second, somewhat surprisingly, there was some support for parameter stability over time in the mobile model regressions. Specifically, when the pooled fully saturated regression (with 1992 dummy intercepts for premium brands) was compared with the year-by-year regressions, the parameter restrictions of the former did not reduce goodness of fit significantly; the traditional F-statistic was 0.97 (with a .01 critical value of 1.60), while the ΔRMSE was 1.00%. However, this non-rejection of parameter stability over the entire time period masks somewhat the possibility that 1992 may be different. In particular, based on a comparison of year-by-year with adjacent-year regressions, the null hypothesis of parameter stability (except for time dummies) is clearly not rejected in 1989-90 and 1990-91 (F-test statistics of 0.46 and 0.49, .01 critical values of 2.01 and 2.07, and ΔRMSE of -2.72% and -2.04%), but is rather marginal for 1991-92 (F-statistic of 2.07, .01 critical value of 2.18, and ΔRMSE of 4.70% being very close to our 5% threshhold). We consider this 1992 result further in Section V below.

Based on the pooled model, when the more restrictive T-A and T-V models were compared to the fully saturated model, again the loss of fit was insignificant; the F-statistic for the nine restrictions of the T-A model was 1.60 (.01 critical value of 2.41), and for the four restrictions of the T-V model it was 1.02 (.01 critical value of 3.32); the respective ΔRMSE's were 0.82% and 0.14%. However, the two restrictions of the V-A model were decisively rejected (F-statistic of 8.85 compared to a .01 critical value of 4.61, and ΔRMSE of 8.94%). Hence the T-A and T-V models empirically dominate the V-A

specification. In terms of choice between the T-A and T-V models, we can

compare their RMSE (recall they differ in number of parameters), and on this

criterion, there is a modest preference for the traditional T-A hedonic price

equation -- 0.2650 for T-A vs. a slightly larger 0.2664 for T-V. Moreover, with

this T-A model, the specification test outlined in Section III yields supportive

results; T-A as a special case of the fully saturated model is not rejected.[14]

Finally, with T-A, a Box-Cox transformation yields further support for the log-

log model; the estimate of λ is 0.1854 and its 99% confidence interval ranges

from -0.1958 to 0.4094, thereby including the λ = 0 (log-log) specification.

Parameter estimates from the pooled T-A model are given in Table 5. As

expected, coefficient estimates on LRAM, LMHZ, LMAXRAM and LHDDSK are positive

and significant, with that on LHDDSK being particularly large (1.27). Since the

LDENSE variable is defined as log(WGT/SIZE), the total effect of LWGT on LPRICE

is the sum of the coefficients on LWGT and LDENSE -- about -0.20. The marginal

effect of LSIZE on LPRICE is the negative of the LDENSE parameter estimate --

therefore about -0.19. Hence increases in weight and in size, ceteris paribus,

reduce prices of mobile models, but there is also a premium for density. The

estimate of the DCOLOR parameter is positive, as is that on DNOHARD, and both

are substantial (the no hard disk models tended to be very small, powerful and

expensive models). Although the DAGE dummy variables have positive parameter

estimates consistent with the notion of "survival of the fittest", the number of

AGE3 models is very small, and thus the insignificance of this parameter is not

surprising. In terms of brand effects, the APPLE, CMPQ and ZEN premia are

substantial until 1992, but the large negative estimates of DAPPLE92, DCMPQ92,

DIBM92 and DZEN92 reveal a sharp change in pricing strategies by these companies

to meet heightened competition. Interestingly, the slightly positive estimate

of DTOSH92 suggests that Toshiba did not follow the other major brands in

TABLE 5

PARAMETER ESTIMATES OF POOLED T-A REGRESSION FOR MOBILE COMPUTERS
(t-statistics from heteroskedasticity-robust standard errors in parentheses)

Variable	Estimate	Variable	Estimate	Variable	Estimate	Variable	Estimate
CONSTANT	1.2085 (1.80)	DNOHARD	0.8221 (5.63)	DAST	-0.2093 (2.71)	DNEC	0.2295 (4.08)
LRAM	0.7769 (10.38)	DPROC8	-0.1399 (2.06)	DCMPQ	0.5279 (11.49)	DTOSH	0.1152 (2.30)
LMHZ	0.1862 (1.75)	DPROC32	0.3143 (7.51)	DCMPQ92	-0.6416 (6.22)	DTOSH92	0.0925 (1.12)
LMAXRAM	0.0779 (2.26)	DAGE1	0.0645 (1.63)	DCOMPUAD	-0.3924 (4.34)	DZEN	0.4413 (7.12)
LHDDSK	1.2739 (7.98)	DAGE2	0.2446 (3.16)	DDELL	-0.1109 (2.34)	DZEN92	-0.5852 (6.52)
LWGT	-0.3927 (3.18)	DAGE3	0.0483 (0.73)	DDELL92	-0.1249 (1.51)	D90	-0.1393 (2.79)
LDENSE	0.1913 (3.22)	DAPPLE	0.5175 (8.06)	DHP	0.3525 (3.76)	D91	-0.2595 (4.09)
LRAMHARD	-0.1504 (7.51)	DAPPLE92	-0.6547 (5.89)	DIBM	0.0975 (1.28)	D92	-0.7634 (7.27)
DCOLOR	0.4601 (9.99)	DALR	-0.0263 (0.30)	DIBM92	-0.4315 (4.15)	R^2 RMSE N	0.6890 0.2660 427

reducing 1992 prices. Finally, the very large and negative estimate of the D92 dummy variable suggests that in 1992, quality-adjusted prices of mobile computer models dropped significantly. This raises issues of price index measurement, to which we now turn our attention.

V.   PRICE INDEX RESULTS

Price indexes based on the various hedonic price equations can be constructed in a variety of ways, often reflecting different assumptions concerning what is being held fixed over time. In this section we construct and

comment on several price indexes, based for the most part on our hedonic

regression equations, but varying in their interpretation and in their use of

parameter estimates and quantity weights.

V.1 PRICE INDEXES FOR MOBILE COMPUTER MODELS

Recall from the introduction that a crude measure of price increase is

simply the change over time in the unweighted sample means of models. For

mobile computer models, as is shown in the first row of Table 6, for 1989-92

this yields an AAGR of -10.6%. Since no matched models occur between 1989 and

1990, one cannot replicate the BLS-like matched model procedure for the mobile

PC models.[15] Based on the various estimated hedonic price equations, however,

one can construct a number of quality-adjusted price indexes.

One simple procedure for constructing price indexes is to use the dummy

variable coefficient on the last-year time dummy in adjacent year regressions,

and simply exponentiate it. For example, letting the price index for 1989 be

1.0, exponentiate the T90 time dummy from the 1989-90 adjacent year regressions,

and set this equal to the 1990 price index. For 1991, exponentiate the T91 time

dummy from the 1990-91 adjacent year regression, and multiply this by the 1990

price index; do an analogous calculation for the 1992 price index, based on the

1991-92 adjacent year regression. The results of such calculations are given in

the second row of Table 6.[16] There it is seen that the quality-adjusted price

index falls about 13% in 1990, another 12% in 1991, and then drops more

dramatically by about 41% in 1992. The average annual growth rate (AAGR) over

the three-year time period is about -23%. This is more than twice as large a

price decline as that based on simple arithmetic means.

An alternative price index calculation can be undertaken using the

saturated dummy variable model, estimated over the 1989-92 time period.

Analogous to the matched model procedure used by the BLS, one can envisage a

matched vintage price index procedure. Specifically, note that from 1989 to

Table 6

ALTERNATIVE QUALITY-ADJUSTED PRICE INDEXES FOR MOBILE PERSONAL COMPUTERS
(Percent Change from Previous Year in Parentheses)

Procedure	1989	1990	1991	1992	1989-92 AAGR
Arithmetic Means	1.0000	0.9323 (-6.8%)	0.8723 (-6.4%)	0.7144 (-18.1%)	-10.60%
Adjacent Year Regressions	1.0000	0.8741 (-12.6%)	0.7666 (-12.3%)	0.4526 (-41.0%)	-23.22%
Saturated Model	1.0000	0.8832 (-11.7%)	0.8271 (-6.3%)	0.3968 (-52.0%)	-26.52%
Pooled T-A Model	1.0000	0.8699 (-13.0%)	0.7714 (-11.3%)	0.4422 (-42.7%)	-23.81%
Laspeyres - Pooled TA Parameters, Fixed 1989 Quality	1.0000	0.9975 (-0.2%)	0.8911 (-10.7%)	0.5305 (-40.5%)	-19.05%
Laspeyres - Yearly Parameters, Fixed 1989 Quality	1.0000	0.8627 (-13.7%)	0.7339 (-14.9%)	0.5696 (-22.4%)	-17.11%
Paasche - Yearly Parameters, Updated Quality	1.0000	0.9350 (-6.5%)	0.9198 (-1.6%)	0.4887 (-46.9%)	-21.23%
Divisia - Yearly Parameters	1.0000	1.0005 (+0.5%)	0.9762 (-2.4%)	0.4408 (-54.8%)	-23.90%

1990, three vintages overlap -- 1987, 1988 and 1989. Using overall 1989-92 sample means for all other right-hand variables, we construct a fitted price for 1989 and for 1990, where the vintage-time interaction coefficients on the 1987-89 overlapping vintages are weighted by the sample means of the vintage-time dummy variables in the two years, where these weights sum to unity. Then we construct a 1990 price index as the 1990 fitted price divided by the 1989 fitted price. We do similar calculations involving three overlapping vintages for successive years 1991 and 1992. The results of this calculation are given in

the third row of Table 6, where it is seen that the matched vintage price index

based on the saturated model yields price indexes that fall at a slightly larger

rate than the hedonic price index from the adjacent year regressions (an AAGR of

-26.5% vs. -23.2%).

Recall from Section IV.2 that there was some empirical support for the

1989-92 pooled T-A specification, although there also was evidence pointing to

the 1992 parameters being different. In the pooled model, if one exponentiates

the 1990, 1991, and 1992 time dummy coefficients (with 1992 modified to account

for the brand-specific dummies as noted in the previous footnote), one obtains

another set of price indexes, which we present in row four of Table 6. The T-A

model yields AAGR's slightly smaller than those based on the saturated model

(-23.8% vs. -26.5%), but very similar to those based on adjacent year regres-

sions (-23.8% vs. -23.2%). This similarity in AAGR's of quality-adjusted price

indexes is not surprising, since the adjacent year, fully saturated and T-A

models were not rejected as valid special cases of the yearly regression models.

Moreover, the price index calculation for the saturated model is based on a

matched vintage notion, which is not quite the same as that based on time

dummies, and thus the slight difference among them is not surprising.

Each of the above price indexes is based on a regression equation that

implicitly weights each observation equally. An alternative price index concept

is one we call a fixed-weight Laspeyres index. In this index, one employs the

parameter estimates (which reflect implicit characteristic quality prices) and

fixed base-period quantity weights for the characteristics (sample means of the

regressors in the base year, say 1989) to compute fitted prices for 1989, 1990,

1991 and 1992. The ratio of the various yearly fitted values to the 1989 value

is then the Laspeyres quality fixed (1989) price index. Such a calculation

could be done with yearly, or pooled parameters. Since the evidence presented

above is somewhat marginal concerning 1992, we employ and compare results based

on both the year-by-year and pooled coefficients. With pooled coefficients, as

seen in Table 6, the resulting Laspeyres index declines at a smaller rate than

the pooled T-A index (-19.0 vs. -23.8%), even though both are based on identical

parameter estimates; when yearly coefficients are employed, the AAGR becomes

-17.1%. Both these values are considerably smaller (in absolute value) than

that based on the pooled TA model. The reason for this difference is that the

use of 1989 fixed quality weights does not capture adequately the significant

change in quality that occurred for mobile models between 1989 and 1992.

To shed more light on this weighting issue, we also compute a Paasche-type

quality updated price index. For 1990, this Paasche index computes a fitted

price using the pooled and 1990-specific parameters along with the 1990 quality

(mean characteristics) weights, and for 1989 the fitted price is computed using

the same 1990 quality weights, but the pooled and 1989-specific parameters. The

1990 Paasche index is then calculated as the ratio of 1990 to 1989 fitted

prices. For 1991, the quality weights are updated to 1991, and fitted prices

are calculated for 1991 using pooled and 1991-specific parameters, while for

1989 the pooled and 1989-specific parameters are employed with 1991 quality

weights. The 1991 price index is then computed as the ratio of these 1991 and

1989 fitted prices. A similar quality updating occurs for 1992. Hence, unlike

the case with the Laspeyres fixed 1989 quality price index, with this Paasche

updated quality price index the changes over time in quality are explicitly

incorporated as changing weights. As shown in Table 6, with the Paasche index

the AAGR declines slightly more rapidly than the Laspeyres indexes (-21.2% vs

-17.1 to -19.1%).

Yet another set of possible weights to employ in calculating an aggregate

price index is based on model-specific revenue shares that vary over time.

Following procedures detailed in Berndt-Griliches [1993], we calculate a Divisia

price index, using the year-by-year coefficients and models for which we were

able to obtain quantity sales data.[17] It is worth noting here that with the

mobile models, there was very little overlap of models between years -- indeed

in 1992, all mobile models for which sales data were available were new. This

implies that the Berndt-Griliches Divisia index procedure, a procedure that

compares the fitted price of a model in, say, 1992 with a fitted price of an

unobserved model having 1992 characteristics but fictitiously existing in 1991,

is not likely to be as reliable when there is no overlap of models between

years. Somewhat surprisingly, therefore, as is shown in the bottom panel of

Table 6, with this share-weighted Divisia procedure, the decline in the price

index is -23.9% per year, very similar to price index trends based on the

adjacent year and pooled T-A model regressions.[18].

V.2   PRICE INDEXES FOR DESKTOP COMPUTER MODELS

We now turn to desktop models, which dominate mobile models in terms of

numbers (1456 vs. 427), and on the basis of revenue shares (the yearly revenue

shares of desktops in 1989-92 are 83%, 88%, 83% and 76%). Again, we construct a

variety of price indexes, reported in Table 7 below. A price index calculated

as the simple arithmetic means of prices of all models by year, relative to

1989, declines at an annual rate of 9.7%, while one based on the matched model

procedure declines about twice as rapidly, -19.32%. Both these measures

understate considerably the quality-adjusted decline in prices, however.

As seen in the third row of Table 7, when quality-adjusted price indexes

are computed using the saturated model specification with pooled 1989-92

parameters, the index declines at a much larger rate of -31.23%. This rate of

decline is reasonably robust to alternative model specifications, being -32.13%

for the pooled T-A model, and -31.15% based on the adjacent year regressions.

It is even robust to the use of year-by-year coefficients and the Divisia index,

which declines at an AAGR of -31.93%. However, if one employs the fixed 1989

quality weights and computes a Laspeyres index analogous to that discussed

Table 7

ALTERNATIVE QUALITY-ADJUSTED PRICE INDEXES FOR DESKTOP PERSONAL COMPUTERS
(Percent Change from Previous Year in Parentheses)

Procedure	1989	1990	1991	1992	1989-92 AAGR
Arithmetic Means	1.0000	0.9601 (-4.0%)	0.9620 (+0.2%)	0.7363 (-23.5%)	-9.70%
Matched Models	1.0000	0.9214 (-7.9%)	0.9134 (-0.9%)	0.5251 (-42.3%)	-19.32%
Saturated Model	1.0000	0.7428 (-25.7%)	0.6717 (-9.6%)	0.3253 (-51.6%)	-31.23%
Pooled T-A Model	1.0000	0.7496 (-25.0%)	0.6735 (-10.2%)	0.3126 (-53.5%)	-32.13%
Adjacent Year Regressions	1.0000	0.8093 (-19.1%)	0.7356 (-9.1%)	0.3264 (-55.6%)	-31.15%
Laspeyres - Yearly Parameters, Fixed 1989 Quality	1.0000	0.6732 (-32.7%)	0.6403 (-4.9%)	0.2550 (-60.2%)	-36.59%
Paasche - Yearly Parameters, Updated Quality	1.0000	0.9282 (-7.2%)	0.9304 (+0.2%)	0.4119 (-55.7%)	-25.60%
Divisia - Yearly Coefficients	1.0000	0.9462 (-5.4%)	0.9042 (-4.4%)	0.3154 (-65.1%)	-31.93%

above, the rate of decline increases to an AAGR of -36.59%. By contrast, if one

computes a Paasche price index, the decline is considerably smaller, an AAGR of

-25.6%. This suggests that with this data, the choice of weights appears to

make a much more critical difference to price index calculations than does the

choice of model specification. Price indexes based on (statistically rejected)

models with parameter stability imposed are very similar to those based on

yearly regressions and adjacent-year regressions.

It is also worth noting that quality-adjusted prices for desktop models

tended to decline more rapidly than for mobile models (the respective Divisia

indexes decline at -31.93% and -23.90%). This raises the issue of how one might

compute a price index for the combined desktop-mobile markets.

V.3   PRICE INDEXES FOR COMBINED DESKTOP AND MOBILE COMPUTER MODELS

One way of computing a composite PC price index is simply to pool the

data, ignore differences in parameters between mobile and desktop models, and

simply exponentiate the appropriate time dummy variable coefficients. Such a

simple approach has some support, for the evidence presented above suggests that

AAGR's of price indexes are reasonably robust, even when calculated using

hedonic models that are statistically rejected. Using a pooled T-A type hedonic

model, we obtain 1989-92 price indexes of 1.000, 0.771, 0.694 and 0.358,

yielding an AAGR of -28.97%.

Alternatively, one could employ the separate regression models for the

mobile and desktop models along with year-by-year coefficients, and then

calculate a combined mobile-desktop Divisia index. When this is done, we obtain

1989-92 price indexes equal to 1.000, 0.963, 0.914 and 0.349, implying an AAGR

of -29.62%. Hence even though the procedures and models differ considerably,

the resulting average growth trends are surprisingly similar -- a 1989-92 price

index for the total PC market declines on average between 29 and 30% per year.

VI.   SUMMARY REMARKS

Our purpose in this paper has been to construct a number of quality-

adjusted price indexes for personal computers in the US marketplace over the

1989-92 time period. To do this, we generalized earlier work on incorporating

simultaneously time, age and vintage effects into an unbalanced panel data set,

and we developed a corresponding specification test procedure.

Our results can be summarized as follows. Parameters of hedonic price

equations differed in the mobile and desktop PC markets. Although there was

very little evidence supporting parameter stability over time in the desktop

market, in the mobile market the evidence suggested overall parameter stability;

however, even in the mobile market there was a suggestion of parameters being

different in 1992. Given parameter stability in the mobile market, we also

found substantial support for the traditional hedonic price equation involving

time and age variables, in addition to the various measures of quality. In

particular, the traditional T-A model was not rejected as a special case of the

fully saturated dummy variable specification. For both the mobile and desktop

markets, the Box-Cox model provided support for the log-log formulation.

A simple arithmetic mean of prices of models by year reveals a price

decline of about 10% per year. The common government statistical agency

procedure of using only matched model prices generates a much larger rate of

price decline -- about 19% per year. But this matched model procedure is also

unreliable, for by construction it excludes new models. When data on new

models, as well as surviving vintages, are employed in the estimation of hedonic

price equations, a variety of quality-adjusted price indexes can be calculated,

with varying interpretations. Although there are some differences, on average

these quality-adjusted price indexes based on econometrically estimated hedonic

price equations decline at about 30% per year. The average decline in the

mobile market was about 24%, while that in the desktop market was larger at 32%

per year. Together these results imply that taking quality changes properly

into account has a very substantial impact on the time pattern of price indexes.

The research in this paper can be extended in a number of ways. First,

the DATAPRO data set we have employed is based on list price data. To the

extent it is feasible, it would be useful to gather "street" prices for as many

models as possible from advertisements in appropriate magazines, newspapers and

journals. To the extent that increasing competition has widened the gap between

list and transactions prices, the approximately 30% decline in quality-adjusted

price indexes for personal computers may understate somewhat the true rate of

price decline. On the other hand, the 1992 change by premium manufacturers in

list price policy that more closely matched the transactions prices of mail order firms could generate overstated price declines. Which effect dominates empirically is an important empirical issue worthy of further examination.

Second, it would be interesting to link the mobile and desktop markets, which are essentially separate in our analysis. One way of doing this would be to pool mobile and desktop models (with, perhaps, a separate intercept term for each) until mobile models became sufficiently important to generate a "regime shift", and thereafter to treat them separately. The link could be constructed by comparing fitted prices for the various models pre-regime shift (using the pooled parameters) and post-regime fit (based on separate regressions), and then weighting by using a Divisia index.

Finally, the prices and embodied technological change in new models of personal computers have an obsolescing impact on older vintages. A most interesting line of research would be to gather data on prices of used computers, and then to model and estimate the price movements of used models over time in response to the persistent and dramatic technological change embodied in new models.

FOOTNOTES

[1]For historical discussions and references to appropriate studies, see Berndt [1991, Chapter 4], Cole et al. [1986], Dulberger [1989], Fisher [1983], Gordon [1989,1990], Griliches [1961,1971,1988,1990], Hall [1971] and Triplett [1986,1989,1990].

[2]Year-to-year real price changes based on their preferred T-A pooled model were, from 1982-83 to 1987-88, respectively, -33%, -1%, -36%, -34%, -28%, and -35%, implying an AAGR of -28.7%.

[3]A recent unpublished paper by Nelson, Tanguay and Patterson [1992], using a sample of 1841 IBM-compatible PC models (but excluding laptops and notebooks) from 1984 through 1991 and similar estimation procedures yields an AAGR of -23.4%; for the common period of 1984-88, the Berndt-Griliches estimate of price change is an AAGR of -25.5%, while that for Nelson et al. is a comparable -24.5%.

[4]Datapro Information Services Group, Delvan, NJ 08075.

[5]Brand-specific dummy variables were introduced for all brands having a market share of 2.0% or more, where market share here means the number of model versions of a particular brand divided by the total number of all model versions (1,883) from 1989 to 1992.

[6]DATAPRO reports the date of the first sale of each computer model.

[7]In this case, one could identify the distinct nonlinear impacts. For example, if the model had the terms α_1TIME + α_2VINTAGE + α_3AGE + α_4TIME^2 + α_5VINTAGE^2 + α_6AGE^2, one could solve out for AGE = TIME - VINTAGE, substitute in, and thereby obtain the form $(\alpha_1 - \alpha_3)$TIME + $(\alpha_2 - \alpha_3)$VINTAGE + $(\alpha_4 + \alpha_6)$TIME^2 + $(\alpha_5 + \alpha_6)$VINTAGE^2 - 2$\alpha_6 \cdot$TIME$\cdot$VINTAGE, a form that permitted distinct identification of α_4, α_5 and α_6, but only the differences $(\alpha_1 - \alpha_3)$ and $(\alpha_2 - \alpha_3)$.

[8]See, for example, Heckman and Robb [1985].

[9]The restrictions are based on the assumption that V90, T89 and AGE0 are the omitted dummy variables in the T-A, V-A and T-V models having an intercept term.

[10]That the effect of age is the same for all time periods turns out to involve redundant restrictions.

[11]The F-test statistic for the null hypothesis of parameter equality across mobile and desktop models was 13.47, while the .01 critical value is 1.49; the ΔRMSE was 14.05%. The rejection was also decisive when a DMOBILE dummy variable was included.

[12]The L prefix denotes a natural logarithm transformation. Following Berndt-Griliches, we defined LHDDSK and LHDDSKSQ as log(MB + 1) and [log(MB + 1)]^2, respectively. Note that the interaction variables are calculated as the product of the logs, and not as the log of the product.

[13]Estimates of the Box-Cox λ parameter and values of the χ^2 test statistic for the null hypothesis that $\lambda = 0$ are, respectively, -0.05 and 0.35 for 1989, 0.02 and 0.27 for 1990, 0.06 and 1.56 for 1991, and -0.30 and 3.33 for 1992.

[14]The F-test statistic is 1.60, while the .01 critical value is 2.41; the ΔRMSE is 0.82%.

^{15}For 1990-91 and 1991-92, however, matched models do occur. Based on these matched models, price indexes for 1991 and 1992 are 0.9966 and 0.5673, respectively. Thus the 1990-92 AAGR for matched mobile models is -24.7%, which is about twice the comparable growth rate based on arithmetic means (-12.5%).

^{16}Since the 1991-92 adjacent year regressions have 1992-specific brand dummy variables (see Table 5), it is necessary to weight the 1992 brand coefficients by the 1991-92 mean proportion of that brand's models, and add this weighted sum to the 1992 time dummy before exponentiation.

^{17}Proprietary sales data were provided us by the International Data Corporation.

^{18}However, when we employed pooled coefficients, the 1990-92 Divisia price indexes were 0.9786, 1.0169 and 0.7935, implying an AAGR of but -7.42%.

REFERENCES

Arrow, Kenneth J. [1960], "Decision Theory and the Choice of a Level of
    Significance for the t-Test," in I. Olkin, W. Hoeffding, S. G. Gurye, W.
    G. Madow and H. B. Mann, eds., Contributions to Probability and
    Statistics, Stanford: Stanford University Press, 70-78.

Berndt, Ernst R. [1991], "The Measurement of Quality Change: Constructing an
    Hedonic Price Index for Computers using Multiple Regression Methods," ch.
    4 in Berndt, The Practice of Econometrics: Classic and Contemporary,
    Reading, MA: Addison-Wesley Press.

Berndt, Ernst R. and Zvi Griliches [1993], "Price Indexes for Microcomputers: An
    Exploratory Study," Chapter 2 in Murray F. Foss, Marilyn E. Manser and
    Allan H. Young, eds., Price Measurements and Their Uses, Studies in Income
    and Wealth, Vol. 57, Chicago: University of Chicago Press for the National
    Bureau of Economic Research, 63-93.

Chow, Gregory C. [1967], "Technological Change and the Demand for Computers,"
    American Economic Review, 57:5, December, 1117-1130.

Cole, Rosanne, Y. C. Chen, Joan A. Barquin-Stolleman, Ellen Dulberger, Nurhan
    Helvacian and James H. Hodge [1986], "Quality-Adjusted Price Indexes for
    Computer Processors and Selected Peripheral Equipment," Survey of Current
    Business, 66:1, January, 41-50.

Dulberger, Ellen R. [1989], "The Application of an Hedonic Model to a Quality
    Adjusted Price Index for Computer Processors", in Dale W. Jorgenson and
    Ralph Landau, eds., Technology and Capital Formation, Cambridge, MA: The
    MIT Press, 37-75.

Fienberg, Stephen E. and William M. Mason [1985], "Specification and
    Implementation of Age, Period and Cohort Models," in William M. Mason and
    Stephen E. Fienberg, eds., Cohort Analysis in Social Research, New York:
    Springer-Verlag, 45-88.

Fisher, Franklin M., John J. McGowan and Joen E. Greenwood [1983], Folded,
    Spindled and Mutilated: Economic Analysis and U.S. v. IBM, Cambridge, MA:
    The MIT Press.

Gordon, Robert J. [1990], The Measurement of Durable Goods Prices, Chicago:
    Univesity of Chicago Press for the National Bureau of Economic Research.

Gordon, Robert J. [1989], "The Postwar Evolution of Computer Prices," chapter 3
    in Dale W. Jorgenson and Ralph Landau, eds., Technology and Capital
    Formation, Cambridge, MA: MIT Press, 77-125.

Griliches, Zvi [1990], "Hedonic Price Indexes and the Measurement of Capital and
    Productivity: Some Historical Reflections," in Ernst R. Berndt and Jack E.
    Triplett, eds., Fifty Years of Economic Measurement, Chicago: University
    of Chicago Press for the National Bureau of Economic Research, 185-202.

Griliches, Zvi [1988], "Postscript on Hedonics," in Zvi Griliches, Technology,
    Education, and Productivity, New York: Basil Blackwell, Inc., 119-122.

Griliches, Zvi [1971], "Introduction: Hedonic Prices Revisited," in Zvi
    Griliches, ed., Price Indexes and Quality Change: Studies in New Methods
    of Measurement, Cambridge, MA: Harvard University Press, 3-15.

Griliches, Zvi [1961], "Hedonic Price Indexes for Automobiles: An Econometric
    Analysis of Quality Change," in The Price Statistics of the Federal
    Government, General Series No. 73, New York: Columbia University Press for
    the National Bureau of Economic Research, 137-196. Reprinted in Zvi
    Griliches, ed., Price Indexes and Quality Change: Studies in New Methods
    of Measurement, Cambridge, MA: Harvard University Press, 1971, 55-87.

Hall, Robert E. [1971], "The Measurement of Quality Change from Vintage Price
    Data," chapter 8 in Zvi Griliches, ed., Price Indexes and Quality Change,
    Cambridge, MA: Harvard University Press, 240-271.

Heckman, James J. and Richard Robb [1985], "Using Longitudinal Data to Model
    Age, Period and Cohort Effects in Earnings Equations," in William M. Mason
    and Stephen E. Fienberg, eds., Cohort Analysis in Social Research, New
    York: Springer-Verlag, 137-150.

Leamer, Edward E. [1978], Specification Searches: Ad Hoc Inference with
    Nonexperimental Data, New York: John Wiley & Sons.

Nelson, Randy A., Tim L. Tanguay, and Christopher D. Patterson [1992], "A
    Quality-Adjusted Price Index for Personal Computers," Waterville, Maine:
    Colby College, Department of Economics, unpublished manuscript.

Ohta, Makoto and Zvi Griliches [1976], "Automobile Prices Revisited: Extensions
    of the Hedonic Hypothesis," in Nestor E. Terleckyj, ed., Household
    Production and Consumption, Studies in Income and Wealth, Vol. 40, New
    York: Columbia University Press for the NBER, 325-390.

Oliner, Stephen D. [1993], "Constant-Quality Price Change, Depreciation, and
    Retirement of Mainframe Computers," Chapter 1 in Murray F. Foss, Marilyn
    E. Manser and Allan H. Young, eds., Price Measurements and Their Uses,
    Studies in Income and Wealth, Vol. 57, Chicago: University of Chicago
    Press for the National Bureau of Economic Research, 19-61.

Sinclair, James and Brian Catron [1990], "An Experimental Price Index for the
    Computer Industry," Monthly Labor Review, 113:10, October, 16-24.

Stoneman, Paul, Derek Bosworth, Denis Leech and David McCausland [1992],
    "Quality Adjusting the Producer Price Index for Computers in the UK," A
    Report to and Commissioned by the Central Statistical Office, London,
    November.

Triplett, Jack E. [1990], "Hedonic Methods in Statistical Agency Environments:
    An Intellectual Biopsy," in Ernst R. Berndt and Jack E. Triplett, eds.,
    Fifty Years of Economic Measurement, Chicago: University of Chicago Press
    for the National Bureau of Economic Research, 207-233.

Triplett, Jack E. [1989], "Price and Technological Change in a Capital Good: A
    Survey of Research on Computers," chapter 4 in Dale W. Jorgenson and Ralph
    Landau, eds., Technology and Capital Formation, Cambridge, MA: MIT Press,
    127-213.

Triplett, Jack E. [1986], "The Economic Interpretation of Hedonic Methods,"
       Survey of Current Business, 86:1, January, 36-40.

White, Halbert L. [1980], "A Heteroskedasticity-Consistent Covariance Matrix
       Estimator and a Direct Test for Heteroskedasticity," Econometrica, 48:4,
       May, 817-838.