HEDONICS FOR PERSONAL COMPUTERS:

A REEXAMINATION OF SELECTED ECONOMETRIC ISSUES

by

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I. INTRODUCTION

In its executive summary, the recently published U.S. National Research Council Panel Report on Conceptual, Measurement and Other Statistical Issues in Developing Cost-of-Living Indexes states that "Dealing with the ever-changing mix and quality of available goods and services poses the most numerous and difficult problems in constructing the CPI."[1] The NRC report goes on to say that:

> "Hedonic techniques currently offer the most promising approach for explicitly adjusting observed prices to account for changing product quality. But our analysis suggests that there are substantial unresolved econometric, data, and other measurement issues that need further attention....BLS should continue to expand its experimental development and testing of hedonic methods and its support of relevant outside research."[2]

In this paper (which represents research still in progress), we report on several assessments of the robustness of hedonic pricing methods.

II. BACKGROUND

It has long been known that only under a set of very restrictive conditions, estimated parameters on characteristics variables in hedonic price equations can be interpreted either as representing marginal costs or as reflecting consumers' marginal valuations.[3] In an important recent paper, however, Ariel Pakes [2002] builds on the literature in the economics of industrial organization, and argues that in the context of heterogeneous consumers, differentiated product competition and rapid technological change, in general one cannot interpret parameters in hedonic price equations as providing evidence on either demand or cost side interpretations; instead they merely trace out the shape of market budget constraints in characteristics space. Moreover, with differentiated product competition, producers' margins are endogenous, varying both across

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[1]National Research Council [2002], p. 6.
[2]Ibid.
[3]See, for example, Feenstra [1995], Griliches [1971,1988], Rosen [1974] and Triplett [1986,1990a,2000].

products and over time, creating wedges between marginal costs and prices.

Although data on products having characteristics that are vertically differentiated (all consumers view more of the characteristic as having the same sign impact on utility) can under certain conditions yield a priori expected signs on the estimated hedonic parameters, Pakes stresses that due to the dynamic nature of differentiated product competition there is no reason to expect stability in the estimated hedonic parameters over time. Furthermore, products having characteristics that are horizontally differentiated (some consumers receive positive utility with more of the characteristic, other consumers receive negative marginal utility) do not even generate unambiguous expectations regarding signs of the estimated hedonic parameters. According to this view, perhaps empirical researchers have devoted too much attention to whether their hedonic parameter estimates were plausible and consistent with engineering and market realities.

As a consequence, Pakes recommends that for personal computers the Bureau of Labor Statistics should experiment with producing quality-adjusted hedonic price indexes in real time, using data from Internet price quotes and automated econometric estimation and statistical procedures (with polynomials of characteristics variables).

In this paper we expand on some of Pakes' ideas and examine several related econometric issues, in the context of desktop and mobile personal computers in the US over the 1976-2001 time period. Specifically, in addition to assessing parameter equality between desktops and mobile PCs, and parameter stability over time, we devote attention to issues involving heteroskedasticity of residuals and implications for predicting quality-adjusted prices. Our research is still in progress.

Although it has long been known that when one estimates a regression equation with transformed ln y as the dependent variable and then predicts the *level* of that variable based on a retransformation of the estimated regression parameters, one must also incorporate into the retransformation an estimate of the residual variance. Specifically, if one estimates parameters in, for example, a cross-sectional regression equation at time t, having the form

$$\ln y_{it} = X_{it}\hat{\theta}_t + \epsilon_{it} \ , \qquad (1)$$

where $\in_{it} \sim N\left(0, \sigma_t^2\right)$, then a consistent predictor of y_{it} is $\hat{y}_{it}$, where

$$\hat{y}_{it} = \exp\left(x_{it}\hat{\beta}_t + .5 * \hat{\sigma}_t^2\right) \tag{2}$$

and $\hat{\sigma}_t^2$ is the mean squared error of the residuals from Eqn. (1). Traditionally this adjustment for residual variance in log-linear models has not received much attention in the empirical hedonic price index literature,[4] although the log-retransformation issue has been the focus of much attention in health econometrics.[5] One reason for this lack of attention in the hedonic literature is that if one computes predicted price *ratios* in two adjacent time periods (say, years), as is commonly done when constructing price indexes, and if one makes the additional assumption that disturbances are homoskedastic in the two adjacent years, then by assumption $\sigma_t^2 = \sigma_{t-1}^2$, and these disturbance variance terms cancel out. However, if disturbances are homoskedastic within years but heteroskedastic between years, then failure to incorporate disturbance variances when retransforming back from logs to levels results in a bias in the predicted price index ratio equal to 0.5 * (σ_t^2 - σ_{t-1}^2). Notice that if the disturbance variance is increasing (decreasing) over time, then failure to incorporate them results in a downward (upward) bias in the estimated price index ratio. Alternative measures of bias can be derived when the form of heteroskedasticity is more complicated. In the log-linear context, Duan [1983] has shown that a nonparametric adjustment, dubbed a "smearing" estimate, computed simply as the product of $\hat{y}_{it}$ and the sample mean of the exponentiated residuals from Eqn. (1), has very desirable statistical properties.

How one interprets the disturbance term in hedonic regression equations depends of course on details concerning model specification. One could plausibly postulate that when innovation is particularly dramatic, disturbance variance terms might increase, and that as markets settle down and accommodate to technological progress, disturbance variance terms might decrease. This would be consistent with the Pakes view of hedonics, because with dynamic differentiated products competition and changing costs, margins and heterogeneous consumers, there would be no <u>a priori</u> reason to expect homoskedastic disturbances. As an empirical strategy, therefore, it is reasonable to examine empirically whether the data reveal heteroskedasticity,

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[4]However, see Berndt [1991], fn. 16, p. 144, and Heien [1968].

[5]See, for example, Ai and Norton [2000], Duan [1983], Duan, Manning, Norris and Newhouse [1983], Manning [1998], Manning and Mullahy [2001], and Mullahy [1998].

rather than assume homoskedasticity. If heteroskedasticity appears to be present, then preferably one would allow for it in the econometric estimation process.

An alternative approach is to eschew the log-linear specification in the first place. Within the health econometrics literature cited earlier, for example, a common alternative to the log-linear specification in Eqn. (1) is the nonlinear exponential formulation,

$$y_{it} = \exp\left(x_{it}\hat{\alpha}\right) + \epsilon_{it}, \qquad (3)$$

whose parameters can be estimated by nonlinear least squares or other nonlinear methods. Even if ϵ_{it} is heteroskedastic, as long as its expected value is zero, a consistent predictor of y_{it} is simply

$$\hat{y}_{it} = \exp\left(x_{it}\hat{\alpha}\right) \qquad (4)$$

where $\hat{\alpha}$ is the vector of parameters estimated by nonlinear procedures.

In this paper we examine issues of parameter equality between desktops and mobile PCs, and parameter stability across years, where by parameters we include not only the β's in the regression equations, but also the estimated disturbance variances. We then assess the robustness of the estimated price indexes by estimating nonlinear exponential regression models. Finally, we examine robustness further by including in the set of regressors indicator variables for microprocessor type, and interaction variables involving processor type and clockspeed (this research is not yet completed). We begin with a discussion of data sources.

III. PC DATA SOURCES AND DATA TRENDS

Our data set covers a 28-year time period from 1976 through 2003, encompassing a remarkable period of technological changes in the PC industry[6]. Our basic unit of observation is a distinctly identified personal computer model available in the US. Over the 28 years, we have 13,714 observations, an average of about 490 models per year, with more recent years generally having a greater number of observations. Although undoubtedly an incomplete data set, we are reasonably confident that it is representative of trends in the overall market. Of these 13,714 observations, 8,832 are desktop models, and 4,482 are mobile PCs. Most of the data covers machines produced by the larger manufacturers, implying that the models in our data set constitute a large percentage of overall PC market sales.

[6] Our 2003 data only cover Q1.

Our data-gathering project has consisted of three separate efforts over the last fifteen years. For the 1976-88 time period, our data come primarily from mail order magazine and newspaper advertisements listing a vendor, model, and computer attributes such as speed, memory, and storage capacity.[7] Although these data sources provided useful information on actual transactions (not just list) prices, the amount of technical information available was limited. Over the 1982-88 time period, proprietary sales data from the International Data Corporation was available for about 950 of the 1,265 models in the estimation sample. The 1976-81 and 1982-88 data are described in greater detail in Berndt and Griliches [1993].

From 1989 through 2003, we rely on a more standardized source of information, namely, Datapro and the Gartner Group. Details on the 1989-92 data are given in Berndt, Griliches and Rappaport [1995]. Briefly, the 1989-92 data come from the Data Information Services Group, then a division of McGraw-Hill. The price data are list prices of the particular base model, rather than transactions prices. However, a considerable amount of data on technical characteristics is available. For a subset of the models, we have proprietary quantity sales data available from the International Data Corporation.[8] From 1993 through 2003, we employ data from the Gartner Group. These data are comparable with the 1989-92 Datapro series. We have not yet been able to gain access to 1993-03 proprietary sales data at the level of individual models, but we are currently investigating the feasibility of employing sales data at the model-processor type level of aggregation.

Although the coverage spans the entire 28-year history of PCs, our data set has a number of limitations. First, the 1982-88 observations capture transactions rather than list prices, while 1989 onward represent list prices. The rapid growth of mail order in the 1990s suggests that the gap between list and transactions prices has likely closed. To the extent this has occurred, it is likely that our price series will overstate price declines since 1989. However, it is widely known that in 1992 major brand manufacturers changed pricing strategies to bring list prices down considerably to match others' transactions prices.[9] Thus, while price declines might be overstated for 1989-92, since then any bias is likely to be much smaller. Moreover, since the principal switch from transactions to list prices occurs between 1988 and 1989, it is likely

[7]Specifically, prices and models were taken from PC Magazine and New York Times, both in June of that year. Additional technical information came from Byte Magazine.
[8]Unfortunately, Datapro data are not available prior to 1989.
[9]See Chwelos [1999] for further discussion.

that we understate any price decline occurring between 1988 and 1989.

A second limitation of our data is that for 1976-81 and since 1993, we do not have access to quantity sales data. Thus in the price index computations we report here, we implicitly weight all models equally. To the extent the demand curve for PCs is downward sloping, we expect that competitive forces would favor PCs with lower quality-adjusted prices, and that leading-selling models would be those experiencing the largest quality-adjusted price declines compared to the previous year' s models. To the extent this has occurred, one might expect use of non-weighted model data to result in understated price declines. It is worth noting, however, that in Berndt, Griliches and Rappaport [1995], price indexes based on non-weighted adjacent year regressions yielded almost identical 1982-88 average annual growth rates (AAGRs) as those based on Divisia indexes employing quantity sales data: the AAGRs were -23.22% and -23.90%, respectively, for mobile PCs, and -31.15% and -31.93% for desktop PCs.

Implementation of hedonic regression methods requires making a decision on what types of variables one should include as regressors to measure the most salient price-determining characteristics. In our previous work (e.g., Berndt-Rappaport [2001]), to be parsimonious we limited the number of characteristic regressors to a small number, but allowed parameter estimates on these characteristics to vary annually, thereby permitting a relatively unrestricted impact of the estimated changes in characteristics on price. In this previous research the characteristics we included were hard disk memory (in megabytes -- MB), processor speed (in megahertz -- MHZ), the amount of RAM memory (MB), indicator variables for whether the model included a CD-ROM, and for the 1982-88 time period, whether the price quote was a discounted "street" price or a list price. To accommodate possible brand effects in a parsimonious manner, we included indicator variables for computers made by Apple, and another combining many of the "name-brand" IBM-compatible computers (Compaq, Dell, Gateway, HP, IBM, NEC, Toshiba and Zenith). All other brands were therefore in the referenced intercept term.

In this paper we examine robustness of the estimated price indexes by adding as regressors up to 24 different microprocessor type indicator variables observed between 1989 and 2003, and up to 24 interaction terms involving microprocessor type and clock speed (ln mhz). The rationale for distinct microprocessor dummy variables is that they proxy for many sources of user value that are otherwise unmeasured, such as,

e.g., speeds of the buses for communication between the CPU, cache and main memory, and enhancements to instruction sets and motherboards. The rationale underlying inclusion of processor-clock speed interaction terms is that it provides a measure of overall processing power, analogous to an auto' s horsepower that reflects the combination of the number of engine cylinders ($\cong$ processor type) and the revolutions per minute the engine operates ($\cong$ mhz).

Currently our data set does not incorporate processor type variables prior to 1989. Thus in these models estimated over the fifteen year 1989-2003 time period the number of observations is smaller – 7,607 desktop models, 4,677 mobile PCs, and 12,279 overall.

There are many ways in which one could describe price and quality change in the PC market over the 28-year history. In Appendix Table 1 to this paper, we list sample means, by year, for prices and a number of characteristics, separately for mobile and desktop models.[10] From 1976 through 2003, the average PC price for desktops fell by 2.4% yearly. While average prices increased until 1984, then oscillated for several years and hit a peak of $4,634 in 1989, from 1989 through 2003 the average desktop price decreased by about 10.6% per year, with the 2003 price of $964 being about a fifth of the $4,634 price in 1989.[11] For mobile models, the average price increased slightly from $1961 in 1983 (the first year in which we observe mobile PCs) to $2,070 in 2003, an average annual growth rate (AAGR) of 0.3%. The average 2003 price of $2,070 was slightly less than half that of the $4616 average mobile price in 1989.

While prices for mobile and desktop PCs generally fell, particularly after 1989, measures of various technical characteristics increased dramatically. For desktop PCs, hard disk capacity increased by a factor of 87,657 between 1979 (the first year in which hard disks appeared) and 2003, with AAGRs of 60.7% between 1979 and 2003, and a virtually identical 60.2% between 1989 and 2003. For mobile PCs, the 1983-2003 hard disk elasticity increased by a factor of 11,200, implying 1983-03 and 1989-03 AAGRs of 59.39% and 65.5%, respectively. Thus hard disk capacity increased at very similar rates for desktop and mobile PCs.

In terms of random access memory (RAM, in MB), for desktops between 1976 and 2003 the average

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[10]These numbers are based on the larger data set that includes observations for which microprocessor data are not available.

[11]The formula used throughout this paper to compute average annual growth rates of the price index PI between years t-τ and t is $\exp\{[\ln(PI_t/PI_{t-\tau})]/\tau\} - 1$.

model increased by a factor of 22,468, reflecting again virtually identical AAGRs of 44.9% and 44.5% between 1976-2003 and 1989-2003, respectively. For mobile PCs, since 1983 average memory increased by a factor of 2,968 to 2003, implying AAGRs of 49.1% between 1983-2003 and the same 49.1% between 1989 and 2003. Thus, as with hard disk capacity, average RAM memory grew at very similar rates for desktop and mobile PCs, although at a slower pace than did hard disk capacity.

Finally, average processor clock speed has also grown considerably over time, although not at quite the breakneck speeds of hard disk capacity and random access memory. From 1976 to 2003, mean MHZ for desktop models grew by a factor of 1,693, implying an AAGR of merely 31.7%, but accelerating to 42.7% between 1989 and 2003. For mobile PCs, mean MHZ increased by a factor of 438 between 1983 and 2003, an AAGR of 35.5%. Between 1989 and 2003, this growth accelerated to an AAGR of 42.1%, very slightly less than the 42.7% for desktops over the same time period.

IV. ISSUES IN SPECIFYING AND ESTIMATING PC HEDONIC PRICE EQUATIONS

We now briefly discussion various specification and estimation issues, first for the log linear model, and then for the nonlinear exponential specification. Let the price of PC model i in year t be P_{it}, let the value of the jth characteristic for that model be X_{ijt}, and let T_t be a time dummy variable having the value of unity in year t, else being zero.

A. PARAMETER EQUALITY ACROSS DESKTOP AND MOBILE PC MODELS

To examine parameter equality across classes of models, we distinguish desktop PCs from mobile PCs, designating desktops with the first subscript D, and for mobile PCs, M, i.e., the value of the jth characteristic for desktop (mobile) model i in year t is $X_{Dijt}\left(X_{Mijt}\right)$.

In an unrestricted single cross-section log-linear specification, separate hedonic regression equations are specified for desktop and mobile PCs:

$$\ln P_{Dijt} = \alpha_{Dt} + \sum_j \beta_{Djt} X_{Dijt} + \epsilon_{Dijt} \text{ for desktop PCs, and} \tag{5}$$

$$\ln P_{Mijt} = \alpha_{Mt} + \sum_j \beta_{Mjt} X_{Mijt} + \epsilon_{Mijt} \text{ for mobile PCs,} \tag{6}$$

where ln denotes natural logarithm, the ϵ's are random disturbance terms, and the α's and β's are parameters to be estimated. Note that $\beta_{Djt}\left(\beta_{Mjt}\right)$ represents the marginal price effect of a small change in desktop (mobile) characteristic j in time period t, other things equal. When j is a continuous variable, β_{Djt} and β_{Mjt} are

partial elasticities of price with respect to characteristic j.

In a partially restricted cross-section log-linear specification, all "slope" coefficients are equal across desktop and mobile PCs, but the intercept term is allowed to differ between desktop and mobile PCs, i.e., while $\alpha_{Dt} \neq \alpha_{Mt}$, it is assumed that $\beta_{Djt} = \beta_{Mjt} = \beta_{jt}$ for all j=1,...,J characteristics. Here the marginal effects of a change in each characteristic are the same for desktop and mobile PCs. In this partially restricted yearly log-linear specification, the hedonic price equation to be estimated takes the form

$$\ln P_{ijt} = \alpha_{Dt} D_{it} + \alpha_{Mt} M_{it} + \sum_j \beta_{jt} X_{ijt} + \epsilon_{ijt} \qquad (7)$$

where P_{ijt} and X_{ijt} are "stacked" vectors of desktop and mobile PC observations, ϵ_{ijt} is a stacked vector of random disturances for desktop and mobile PCs, D_{it} is a dummy variable equal to one only if model i is a desktop model, else it is zero, and M_{it} is a dummy variable equal to one only if model i is a mobile model, else it is zero.

In a fully restricted cross-section specification, in addition to the constraint that laptop and mobile slope coefficients are equal, it is assumed that intercept terms are equal, i.e., $\beta_{Djt} = \beta_{Mjt} = \beta_{jt}$ for all j=1,...,J characteristics, *and* $\alpha_{Dt} = \alpha_{Mt} = \alpha_t$. In this specification, quality-adjusted price indexes are identical for desktop and mobile PCs. The corresponding fully restricted yearly hedonic log-linear price equation takes the form

$$\ln P_{ijt} = \alpha_t + \Sigma_j \beta_{jt} X_{ijt} + \epsilon_{ijt}. \qquad (8)$$

For the nonlinear exponential specifications, Eqns. (5) and (6) are replaced with

$$P_{Dijt} = \exp(\alpha_{Dt} + \Sigma_j \beta_{Djt} X_{Dijt}) + \epsilon_{Dijt} \text{ for desktop PCs, and} \qquad (9)$$

$$P_{Mijt} = \exp(\alpha_{Mt} + \Sigma_j \beta_{Mjt} X_{Mijt}) + \epsilon_{Mijt} \text{ for mobile PCs.} \qquad (10)$$

Similarly, the partially restricted cross-section log-linear specification of Eqn. (7) is replaced with

$$P_{ijt} = \exp(\alpha_{Dt} D_{it} + \alpha_{Mt} M_{it} + \Sigma_j \beta_{jt} X_{ijt}) + \epsilon_{ijt}, \qquad (11)$$

while the fully restricted cross-section nonlinear exponential model is

$$P_{ijt} = \exp(\alpha_t + \Sigma_j \beta_{jt} X_{ijt}) + \epsilon_{ijt}. \qquad (12)$$

B. PARAMETER STABILITY OVER TIME

An alternative dimension of parameter equality involves stability over time. Again we begin with log-linear models. To abstract from issues involving lack of parameter equality across classes of models, assume

for the moment that based on results of hypothesis tests, we are dealing with a set of models over which the null hypothesis of either partial or full parameter equality has not been rejected, or, if partial parameter equality has been rejected, we are separately estimating cross-section log-linear hedonic price equations for desktop and mobile PCs. To keep our notation manageable, however, we now omit the D and M initial subscripts.

In the unrestricted adjacent time period log-linear regressions for time periods t and t-1, separate equations are specified for each time period:

$$\ln P_{ij,t-1} = \alpha_{t-1} + \Sigma_j \beta_{j,t-1} X_{ij,t-1} + \in_{ij,t-1}, \text{ and} \tag{13}$$

$$\ln P_{ij,t} = \alpha_t + \Sigma_j \beta_{j,t} X_{ij,t} + \in_{ij,t}. \tag{14}$$

In the common and widely used partially restricted adjacent time period log-linear regressions, it is assumed that each slope coefficient is equal (stable) in adjacent years, but that the intercept terms differ, i.e., $\beta_{j,t} = \beta_{j,t-1} = \beta_j$ for all j =1,...,J, but $\alpha_t \neq \alpha_{t-1}$. These parameter restrictions are frequently imposed by estimating adjacent time period hedonic price regressions having the form

$$\ln P_{ijt} = \alpha_{t,t-1} + a_t T_t + \Sigma_j \beta_j X_{ijt} + \in_{ijtt}, \tag{15}$$

where P_{ijt}, X_{ijt} and $\in_{ijt}$ are stacked vectors of observations for time periods t-1 and t, and since T_t is a dummy variable taking on the value of one if the observation is in time period t (else it is zero), a_t is the differential intercept for time period t relative to time period t-1, i.e., $a_t \equiv \alpha_t - \alpha_{t-1}$.

With this partially restricted adjacent year regression specification, frequently the quality-adjusted price index for year t relative to year t-1 is computed simply by exponentiating a_t. When slope coefficients of more than two adjacent time periods are constrained to be identical while intercept terms are permitted to differ, quality-adjusted price indexes relative to the base time period are often computed using the sequence of exponentiated time period-specific intercept differentials, i.e., the sequence of $\exp(a_t)$, $\exp(a_{t+1})$, $\exp(a_{t+2})$, etc.[12] As discussed earlier, for these procedures to produce consistent estimates of the quality-adjusted price index relative to the base time period, one must either assume homoskedasticity of disturbances or make an appropriate disturbance variance adjustment, as in Eqn. (2). We discuss tests for disturbance homoskedasticity

[12]There are of course other more complex ways in which quality-adjusted price indexes can be computed based on estimated hedonic price equations. See Armknecht, Lane and Stewart [1997], Berndt [1991, Chapter 4], Berndt and Griliches [1993], Berndt, Griliches and Rappaport [1995], Fixler, Fortuna, Greenlees and Lane [1999], and National Research Council [2002] for further discussion.

below.

Finally, in the fully temporally restricted hedonic price specification, it is assumed that not only are all slope coefficients in the adjacent time period regressions involving Eqns. (13) and (14) identical, but so too are the intercept terms, i.e., $ß_{j,t} = ß_{j,t-1} = ß_j$ for all j = 1,…,J, *and* $\alpha_t = \alpha_{t-1} = \alpha$, which implies that $a_t = 0$. The corresponding fully temporally restricted hedonic price equation is of the form

$$\ln P_{ijt} = \alpha + \Sigma_j ß_j X_{ijt} + \in_{ijt}. \qquad (16)$$

When the constraints in Eqn. (16) hold, the quality-adjusted hedonic price index for time period t is the same as that for time period t-1.

One way of interpreting parameter inequality over time is in terms of neutral and non-neutral technological progress, as in the theory of cost and production. Specifically, in exponentiated (non-logarithmic) form, one can re-write Eqn. (14), allowing for time-varying elasticities, as

$$P_{ijt} = \alpha_t \bullet Ð_j X_{ijt}^{\hat{a}_{jt}} \in_{ijt}. \qquad (17)$$

If the relationship between characteristics and price is time-invariant, i.e., if the $ß_{jt}$ elasticities are constant, or if technological change is characteristic-neutral, then one can rewrite Eqn. (17) as

$$P_{ijt} = \alpha_t \bullet Ð_j X_{ijt}^{\hat{a}_j} \in_{ijt}, \qquad (18)$$

which is analogous to the familiar Cobb-Douglas functional form.

C. ESTIMATION

For the linear regressions, parameters are estimated both by ordinary least squares (OLS) and generalized least squares (GLS), where the latter allows for arbitrary heteroskedasticity of disturbances. GLS is implemented using the RREG command in STATA, which employs Huber and bi-weight iterations until convergence is attained.[13] For the nonlinear exponential regressions, we will utilize the NL command in STATA. Although we intend to obtain heteroskedasticity-consistent standard errors in the nonlinear least squares estimations, we have not yet completed that research.

D. HYPOTHESIS TESTING AND THE COMPUTATION OF QUALITY-ADJUSTED PRICE INDEXES

[13] STATA Corp. [1999].

For the log-linear models, we test hypotheses using the TEST commands in STATA in both the OLS and GLS estimations. The hypotheses tested in the nonlinear exponential models are the same joint linear hypotheses as in the log-linear formulation, and therefore they are also implemented using the TEST commands along with STATA's NL estimation procedure.

In the log-linear framework, consistent with the hedonic aggregation approach outlined in Triplett [1990b], we implement the sequence of hypothesis tests by first running yearly regressions separately for desktop and mobile PCs as in Eqns. (5) and (6), and testing whether the restrictions in Eqn. (7) are valid. If these restrictions are not rejected, we impose them, and continue on with various temporal stability tests. If the restrictions in Eqn. (7) are rejected, we do not impose them, and instead continue on with various temporal stability tests, separately for desktop and mobile PCs.

When annual estimation takes place over the entire 1983-2003 time period (recall that the mobile PC data begins in 1983), account should be taken of the simultaneous nature of these hypothesis tests. Using the Bonferroni test procedure, we assign an overall p-value of 0.05 to the sequence of 21 annual tests, and thus assign a p-value of 0.0024 (0.05/21) to each individual annual test for parameter equality. For the regressions based on 1989-2003 data, the overall p-value of 0.05 implies individual year p-values of 0.0033 (0.05/15). Analogous procedures are followed for the nonlinear estimation.

To assess parameter stability over time for the desktop and mobile PCs, we first estimate cross-sectional hedonic price equations for each pair of adjacent time periods, as in Eqns. (13) and (14), and then test whether the restrictions in Eqn. (15) are valid. Since this sequence of tests for parameter stability is undertaken for each pair of adjacent years, for the overall sequence of 27 tests to have a p-value of 0.05, each adjacent year test is assigned a p-value of 0.0018 (0.05/27). For the adjacent year mobile models estimated over the 1983-84 to 2002-03 time period, the 0.05 overall p-value is consistent with p-values of 0.0025 for each adjacent year regression.

If the parameter stability restrictions are not rejected, we compute quality adjusted price indexes by chaining the sequence of estimated exponentiated a_t's, as discussed above. If, however, the parameter restrictions in Eqn. (7) are rejected, we compute various quality adjusted price indexes for successive time periods by employing parameter estimates from the separate adjacent time period cross-sectional regressions

and computing predicted values based on alternatively the mean values of characteristics from time period t-1 (analogous to the base period quantity weights used in the Laspeyres index), from time period t (analogous to the current time period quantity weights used in the Paasche index), and then compute the Fisher Ideal price index as the geometric mean of the these "Laspeyres" and "Paasche" price indexes. Analogous hypothesis testing procedures are followed for price indexes based on the nonlinear estimation.

Finally, to examine empirically the possibility of heteroskedasticity, we take residuals from temporally pooled regressions, square them, and then run an auxiliary regression of the squared residual on a constant and a set of annual time dummies, thereby yielding yearly estimates of the disturbance variances. A joint test of homoskedasticity across all years is based on an F-statistic from this auxiliary regression. Analogous tests are done for adjacent year regressions, both in the log-linear and nonlinear exponential regressions.

V. ECONOMETRIC FINDINGS ON PARAMETER EQUALITY AND PARAMETER STABILITY

We now report on econometric findings obtained from estimation of a variety of hedonic regression equations. Since the results are numerous, here we focus on the most salient findings. We begin with the relatively simple log-linear results.

A. PARAMETER EQUALITY BETWEEN MOBILE AND DESKTOP MODELS

Mobile models have evolved considerably since their introduction in the early 1980s. Early mobile models were essentially suitcase-sized versions of desktop machines with a smaller screen replacing the desktop monitor, and as their "luggable" nickname implied, they could be lugged from place to place but only with some difficulty. With the advent of laptop computers in the late 1980s and notebooks in the 1990s, true computing portability was achieved. But this portability came with a cost. In exchange for lighter weight and smaller footprint came slower processor speeds, less memory, and less storage capacity, all at typically a premium price relative to a desktop with otherwise similar characteristics.

Nine luggable models first appear in our data set in 1983, increasing to 16 mobile models in 1984, 22 in 1985, 32 in 1986 and 73 in 1987; in 2003 Q1, we observe 200 distinct mobile models.

We estimate parameters in yearly regressions for 1983-03, first allowing for common slope coefficients for mobile and desktop models but a different intercept, as in Eqn. (7), and then allowing both for

different intercepts and different slope coefficients, as in Eqns. (5) and (6). From 1987 onward, for both OLS and GLS estimation, the coefficient estimate on the DPORT dummy variable in Eqn. (7) is positive and statistically significant. Interestingly, the magnitude of the OLS (GLS) estimated DPORT increased from 0.040 (.088) in 1983 to 1.086 (1.124) in 1998, and then began falling to 0.932 (0.938), 0.822 (0.829) and 0.642 (0.692) in 1998-2001, respectively, and then leveled off in 2002-3. This implies that, when one constrains the characteristic coefficients (speed, hard disk capacity, memory, CD-ROM, discount, Apple, and IBM-COM) to be the same across mobile and desktop PCs, in 1998 (2003) the price of a notebook computer was about 3.0 (1.9) times that of desktop having the same technical characteristics. The time trend of estimated coefficients also suggests that the mobile price premium hit a peak in 1998, with the subsequent decline reflecting incentives for and the effects of greater entry and price competition in the mobile PC market segment.

For each year from 1983 onward, we then test the null hypothesis that the slope coefficients are equal across mobile and desktop models, against the alternative hypothesis that both slope and intercept terms differ. Using an individual year p-value of 0.0238 (=0.05/21 tests), beginning in 1989 and for each year thereafter, based on the GLS estimations we reject the null hypothesis of slope parameter equality between mobile and desktop PCs. Although the OLS-based results are qualitatively similar, parameter equality cannot be rejected in 1983-1987, and in 1997 (p value of 0.0362). The general pattern, therefore, is that beginning about 1989 and continuing thereafter, the effects of changes in characteristics on prices differ between desktop and mobile models.

It is also of interest to examine which characteristic coefficients appear to differ between mobile and desktop models. To conserve space we do not report detailed results, but instead offer a brief summary in the top panels of Table 1. In the first column of Table 1, we present parameter (elasticity) estimates based on the pooled over all years (1976-2003 for desktops, 1983-2003 for mobile) regressions, in which slope coefficients are constrained to be constant. When pooled over all years, the mobile hard disk capacity elasticity is 67% (0.15 vs. 0.09) larger than that for desktops. Subsequent columns reveal elasticity estimates over sub-periods. We find that from 1983 through 1994, the estimated elasticities of price with respect to hard disk capacity were similar for desktops and mobile PCs, but beginning in around 1995 this elasticity became much larger for mobile than desktop PCs.

By contrast, the estimated elasticities of price with respect to clock speed for desktops is always larger than for mobile PCs, and this gap increases sharply in 1999-2003; the temporally pooled clock speed elasticities are 0.42 (desktop) and 0.16 (mobile).

Finally, in terms of RAM, the relative size of elasticities flip during the 1980s and 1990s; the temporally pooled elasticity estimates for RAM are 0.32 (desktop) and 0.23 (mobile), but tend to coverge by the end of the sample time period.

In summary, for the log-linear specification, coefficient estimates on characteristics differ significantly between mobile and desktop models, generally beginning in 1989 and continuing annually thereafter. The proportional price premium charged for mobile vs. desktop models increased steadily from 1983 to 1998, and then began falling. Mobile models have increasingly had a larger elasticity of price with respect to hard disk capacity than have desktop PCs. In contrast, while elasticities of price with respect to processor speed have always been larger for desktop than mobile PCs, the difference between them has increased over time.

[Non-linear exponentiated model – research still underway]

B. PARAMETER STABILITY OVER TIME

The null hypothesis of stability of all slope coefficients over time is decisively rejected, both for desktop and mobile PC models (p-value <0.0001). Although it is possible to examine parameter stability over various lengths of time, separately for mobile and desktop models, we simplify the analysis here by assessing stability in successive pairs of adjacent year regressions.

As discussed above, because estimated slope coefficients for desktop and mobile PCs do not generally differ significantly until 1989, we first examine parameter stability for the pooled sample (but with a distinct mobile intercept) for years up to 1989. We find that up through and including 1987-88 (OLS estimates) and 1986-87 (GLS), we cannot reject the null hypothesis of all slope coefficients being equal in adjacent years; from 1988-89 (OLS) and 1987-88 (GLS) onward through 1988-89, however, adjacent year parameter stability is rejected.

Since the GLS results indicate that mobile and desktop coefficients differ beginning in 1989, from 1989 onward we assess parameter stability separately for mobile and desktop PCs. For desktop PCs,

beginning in 1992-93, for both OLS and GLS estimates, the null hypothesis of slope coefficients being equal in adjacent years (but allowing for different intercepts) is rejected in all pairs of adjacent years.

For mobile PCs, adjacent year parameter stability is rejected in 1991-92 and then for all subsequent pairs of years beginning in 1994-95, except for 1997-98, where the OLS p-value is 0.2308 while that for GLS is 0.0359. Thus there is some evidence suggesting that the mobile PC market generally exhibited parameter stability for several years beyond that for desktops, 1994 vs. 1992.

We interpret these findings as suggesting that in the desktop PC segment, particularly since the early 1990s, there have been dramatic non-neutral annual changes in the relationships among technical characteristics and price; for mobile PCs (except for an outlier in 1997-98), these non-neutral annual changes apparently began several years later.

Since presentation of all parameter estimates would take an enormous amount of space, in the top three panels of Table 1 we briefly summarize patterns of parameter change over time, separately for desktop and mobile PCs, over selected time intervals. As noted earlier, in the first column of Table 1, we present parameter (elasticity) estimates based on the pooled over all years (1976-2003 for desktops, 1983-2003 for mobile) regressions, in which slope coefficients are constrained to be constant. The remaining columns in the top three panels of Table 1 are arithmetic means of parameter (elasticity) estimates from the adjacent year log-linear regressions over selected time intervals.

As seen in the top rows of Table 1, the elasticity of price with respect to hard disk capacity was very high at about 0.39 when hard disks initially appeared on desktop PCs in the early 1980s, this elasticity dropped in half to about 0.19 in 1984-89, it dropped even further to about 0.05 in 1990-93, but particularly for mobile PCs it increased sharply between 1994-97, only to fall again between 1998-2003. Recall that in the 1980s the early mobile PCs were luggable, in the later 1980s and early 1990s smaller but still rather heavy and cumbersome portables and laptops appeared, and finally the very thin and light notebooks have come to predominate in the late 1990s. In recent years space for hard disk capacity has apparently commanded a hefty but declining price premium in mobile PCs -- the 1994-97 mobile elasticity of 0.37 was more than four times that for desktops at 0.09, but by 1998-2003 these elasticities become 0.21 and 0.07.

Trends in processor speed are quite different, as is seen in the second top panel in Table 1. Since 1980

for desktop PCs, the elasticity of price with respect to speed has been much larger than that for hard disks, and by 1998-2003 the desktop elasticity of 0.33 is many times that for mobile PCs at 0.02.

Finally, in terms of RAM, as is seen in the third panel in Table 1, there is no clear monotonic pattern over time and between mobile and desktop PCs in the estimated elasticity of price with respect to RAM. Intertemporal variability here is clearly considerable.

[Nonlinear exponential model research still in progress]

C. HOMOSKEDASTICITY OVER TIME

Tests for homoskedasticity can be done in a number of ways. One way is to take the squared residuals from the temporally pooled combined mobile-desktop, desktop only, and mobile only log-linear regressions, and regress the squared residuals on a constant term and a series of yearly dummies (excluding a base year, such as 1992). The estimated yearly residual variances from these intertemporally pooled regressions are plotted in Figure 1. The desktop and pooled desktop-mobile residual variances generally increase in the 1970s and early 1980s, then fall later in the 1980s, with a particularly large decrease in 1984. After about 1992, while the pooled desktop-mobile residual variances continue to decrease in the 1990s and level off in 1999-2003, the desktop residual variances increase considerably in 1999-2001, only to fall again in 2002-3. In contrast, although there is an initial positive spike in 1986, for mobile PCs the residual variances show a generally declining time trend. An F-test for joint equality of the residual variances across all years is decisively rejected for all three temporally pooled log-linear regressions (p-values < 0.001). The ratio of largest to smallest residual variance is 11.1 for the pooled models, 3.9 for desktop PCs, and 5.7 for mobile PCs.

The rejection of disturbance homoskedasticity in the above log-linear regressions could reflect the inappropriate assumption of intertemporal stability in the underlying parameters on the characteristics variables, and/or the invalid pooling of desktop and mobile PCs. Additional research will examine patterns of heteroskedasticity in the residuals.

However, of even more basic interest is the variation over time in the distribution of prices. In Figure 2 we display the annual standard deviation of observed prices, for pooled, desktop only, and mobile only PCs. As is seen in Figure 2, for the pooled desktop-mobile sample, the time pattern of standard deviations of

observed prices is quite similar to that of the residual variances, and in particular, for both there is a noticeable decline in the last ten years of the sample. Standard deviations of PC prices for desktop-only PCs display even more dramatic trends: after oscillating up and down until 1984, they steadily rise to a peak of $3,768 in 1991, and then generally decline, reaching a minimum of $2,112 in 2003. Interestingly, while variability in mobile PC prices is generally less than that for desktop PCs through 1992, from 1993 through 1996 the price variability is similar, but from 1997 on variability in mobile PC prices is greater than that for desktop PCs.

VI. ALTERNATIVE QUALITY-ADJUSTED PRICE INDEX CALCULATIONS

The econometric findings reported above for parameter equality between mobile and desktop PCs, and parameter stability over time, have implications for the construction and interpretation of quality-adjusted price indexes. We have constructed alternative price indexes that vary in the way they accommodate parameter inequality across desktop and mobile PCs, and instability in parameter estimates over time. Alternative results are presented in Tables 2 and 3. Table 2 presents various price indexes based on the parsimonious log-linear model, while Table 3 presents a small set of price indexes for the expanded log-linear model with microprocessor type and microprocessor-clock speed interactions included as regressors. Our research is still underway, and particularly for the nonlinear and expanded variable specifications, not all indexes have been computed. We begin with the parsimonious log-linear specification in Table 2.

The simplest and most restrictive procedure is simply to pool observations across all years, restrict all slope coefficients to be constant over time, allow for a time-invariant dummy variable for mobile models, and specify yearly time dummy variables. Results from this simplest way to adjust PC prices for quality change are presented in Column 2 of Table 2. Under these assumptions, adjusted for quality changes, PC prices have fallen by a factor of 1,816 from 1976 through 2003, implying an AAGR of -24.27%. Holding quality fixed, in the first seven years of our sample (1976-83), PC prices fell on average about 13% per year, and for the next decade between 1983 and 1993 the price decline accelerated to about –20% annually. The rate of price decline then almost doubled to –38% between 1993 and 1998, but in the last five years this price decline decelerated to a still very respectable –31% per year.

A slightly less restrictive set of price index calculations involves maintaining the assumption of stability of all slope coefficient parameters over time, but separately estimating models for desktop and mobile

PCs. Results from these calculations are given in Columns 3 and 4 of Table 2.

For desktop PCs, the 1976/2003 ratio of quality-adjusted prices is 2,702:1, and over the shorter 1983-2003 and 1989-2003 time periods it is 687:1 and 226:1. Over the entire 1976-2003 time period, the AAGR is -26.03%, but this masks considerable intertemporal heterogeneity. Between 1976 and 1983 the AAGR is about -14%, increasing to about -22% between 1983 and 1993. It then accelerates dramatically to about –41% from 1993 to 1998, and finally decelerates modestly to -33% from 1998 to 2003.

Quality-adjusted price declines are considerably smaller in magnitude for mobile than desktop PCs, as one sees by comparing Columns (3) and (4) in Table 2. The 1983/2003 ratio of quality-adjusted prices for mobile PCs is 90:1 (about one-eighth that for desktops), reflecting an AAGR of -20.14%. Although mobile PC price declines were rather modest until 1993 (around -12%), they more than doubled between 1993 and 1998 (-28%), and then declined slightly to –27% between 1998 and 2003.

While provocative and intriguing, these price indexes adjusted for quality change depend on strong assumptions regarding parameter stability over time. Recall from our earlier econometric findings discussion that for slope coefficients in yearly regressions, significant differences between desktops and laptops generally began to appear in 1989, and continued thereafter.

For purposes of comparison, in columns 5 and 7 of Table 2 we present quality-adjusted price indexes separately for desktop and mobile models, respectively, based on adjacent year regressions. For desktops, over the entire 1976-2003 time frame and for all sub-periods except between 1998 and 2003, AAGRs based on adjacent year regressions decline more rapidly (sometimes considerably) than those based on intertemporally pooled desktop regressions. Over the entire 1976-2003 period the AAGR difference is about two percentage points (-28.19% vs. -26.03%), for 1976-83 the difference is just under seven percentage points (-20.2% vs. -13.6%), but between 1983 and 1988 the AAGRs are very similar at –22.8% and –22.6%. This understatement of price decline when pooling desktop models over time increases to about 4% between 1988-93 (-25.8% vs.. –21.9%), and then falls to just more than one percentage point during the torrid 1993-1998 era (-42.2% vs. –40.8%). In the last five years of the sample, however, price indexes based on adjacent year regressions fall less rapidly than those based on intertemporally pooled hedonic regressions; the 1998-2003 AAGRs are –25.31% vs. –32.55%.

For mobile PCs, comparison of columns 7 and 4 indicates that in all time sub-periods except 1998-2003, AAGRs based on adjacent year regressions decline considerably more rapidly than those based on the intertemporally pooled regression. This difference is particularly notable between 1993 and 1998, when the AAGRs are –42.21% vs. –28.35%. Over the entire 1983-2003 time frame, the difference is about five percentage points (-25.28% vs. -20.14%). However, between 1998 and 2003 the inequality is reversed, with adjacent year regressions for mobile PCs having an AAGR of –23.0%, while those based on the intertemporally pooled regression having an AAGR of –26.8%.

Recall from our hypothesis test results that for desktops, yearly differences in coefficient estimates became significant beginning in 1992, and for mobiles, in 1994. In constructing a Laspeyres-type quality-adjusted price index between adjacent years t and t+1, we use parameter estimates from the separate year t and t+1 desktop regressions, we employ as "weights" the means of the desktop regressors in the "base" year t, and then take the ratio of predicted desktop prices in years t+1 and t. The analogous Paasche-type quality adjusted price index employs as quantity weights instead the means of the desktop regressors from the "current" year t+1. The Fisher Ideal is the geometric mean of the "Laspeyres" and "Paasche" price indexes.

[Fisher Ideal quality-adjusted research computations still underway]

To summarize, based on the log-linear models, in general the intertemporally pooled models reveal smaller price declines than those based on adjacent year regressions, although this inequality does not hold for all subperiods.

Recall from our earlier discussion that if intertemporal disturbance heteroskedasticity is present, then merely exponentiating the time dummies to transform from log-linear to price level results in biased price index ratios. Ocular least squares estimates of the time trend in the estimated residual variance (see Figure 1) suggest that for desktop models, residual variance increased to 1983, then oscillated but leveled off between 1986 and 1995, increased between 1995 and 2001, falling again in 2002-3. The general upward trend in residual variance for desktops in the final decade of the sample suggests that the high rates of price decline during these years, presented in Table 2, may be overstated. By contrast, the substantial decline in the estimated residual variances of the mobile PCs (see Figure 1) over the latter portion of our time series implies that for mobile PCs, the AAGRs in Table 2 will understate price declines. We will assess the magnitude of

this heteroskedasticity-induced bias, using alternatively the parametric adjustment as in Eqn. (2), and the non-parametric smearing factor procedure developed by Duan [1983].

We now return in Table 3 to the log-linear specification, but add as regressors the up to 24 microprocessor type dummy variables, as well as up to 24 microprocessor - ln mhz interaction terms. Recall that our time series now covers only 1989-2003, for reasons of data availability. The price indexes reported do not adjust for residual heteroskedasticity. Research involving computation of all pooled, desktop pooled, mobile pooled and the yearly Fisher price indexes for desktop and mobile models is not yet complete. Several results are worth noting, however, particularly regarding the adjacent year regression-based price indexes in columns (4) and (6).

First, the comparison of column 4 AAGRs in Table 3 with entries from column 5 of Table 2 based on the more parsimonious log-linear specification indicates that for all time periods, price declines in the expanded log-linear are smaller than those in the more parsimonious log-linear model. Over the entire 1989-2003 time period, the desktop AAGRs are –29.48% (expanded) vs. –33.29% (parsimonious). While differences are relatively modest in 1989-1993 (-28.42% vs. –30.78%) and 1993-98 when price declines are largest (-39.62% vs. –42.15%), in the final five years of the sample, 1998-2003, the difference becomes much larger (-18.63% vs. –25.31%).

Second, for mobile PCs, as seen in column 6 of Table 3, the smaller price decline in price indexes based on the expanded vs. parsimonious follows the same general pattern as with desktop PCs (-25.34% vs. –29.00%), but now in both the 1989-1993 (-11.78% vs. –16.99%) and 1998-2003 (-18.44% vs. –23.01%) years, the difference is approximately the same at around five percentage points. Notably, as with desktops, the difference is smallest (-40.19% vs. –42.21%) during the torrid 1993-98 years when price declines were the largest.

In research currently underway, we are examining whether the smaller price declines obtained from the expanded vs. parsimonious adjacent year specifications continue to emerge based on intertemporally pooled and yearly regressions, and whether any systematic differences in AAGRs become larger or smaller with temporal disaggregation.

VII. CONCLUSIONS AND SUGGESTIONS FOR FURTHER RESEARCH

Our research is ongoing and what we have reported here is not yet complete. What we have found to date is that significant heterogeneity in parameter estimates of hedonic price models has existed between desktop and laptop computers at least since 1989, and that annual heterogeneity in parameter estimates began for desktops in about 1992, and for laptops in 1994. These results are robust across OLS and GLS estimation. Estimates of disturbance variances also display substantial heteroskedasticity, and while these variances appear in general to be smaller in the last half of the sample, there is no clear monotonic trend with time.

Not surprisingly, quality-adjusted price indexes based on these various estimated models display considerable variability as well. Several patterns have emerged from our analysis to date. First, although the difference in AAGRs has declined in the most recent five years (1998-2003), AAGRs of mobile PCs are in most cases less than those of desktop PCs; over the 1983-2003 time span, these AAGRs are –20.14% and –29.93%, respectively, based on intertemporally pooled regressions, and –25.28% vs. –30.78% based on adjacent year regressions.

Second, for both mobile and desktop PCs, price declines were considerably larger in the 1993-2003 time period than in the ten years preceeding 1993. Quality-adjusted price declines were particularly dramatic between 1993 and 1998, when, based on adjacent year regressions, AAGRs were –42.15% for desktops and –42.21% for mobile PCs. Since that remarkable time period, between 1998 and 2003, price declines have decelerated, with still very respectable AAGRs of –25.31% (desktop) and –23.01% (mobile).

Third, although the evidence to date is somewhat limited, based on adjacent year regressions it appears that price declines based on expanded specifications (with processor and clockspeed-processor interaction variables included as regressors) are modestly smaller than those based on more parsimonious specifications. Over the 1989-2003 time period, the respective AAGRs were –29.48% vs. –33.29% for desktop PCs, and –25.34% vs. –29.00% for mobile PCs.

Finally, a comment regarding variability in hedonic price estimates. In order to provide empirical evidence to government statistical agencies in determining how to implement hedonic pricing methods, and to assess the robustness and credibility of hedonic-based price indexes (relative to matched model price indexes), we suggest that governments put together a group of applied economists, engineers and marketing specialists,

that this group choose a product and specify what price and characteristic data they want for this product, and then that members of the group each come up with their preferred estimated hedonic price model. Based on these various preferred models, a set of price indexes would then be constructed. An examination of the variability in these alternative quality-adjusted price indexes could provide valuable empirical evidence that could facilitate greater diffusion of hedonic pricing methods in the construction of official price indexes.

Table 1

Mean Values of Elasticity Estimates on Characteristics Variables Over Selected Time Intervals, Based on OLS Linear Regression Equations without Clockspeed - Processor Interactions

	Pooled Time	Means from Adjacent Year Regressions					
Characteristic	Regression	76-79	80-83	84-89	90-93	94-97	98-03
Hard disk memory (ln mb)							
Desktop	0.09	na	0.39	0.19	0.05	0.08	0.07
Mobile	0.15	na	na	0.18	0.05	0.37	0.21
Processor speed (ln mhz)							
Desktop	0.42	0.24	0.48	0.69	0.53	0.60	0.33
Mobile	0.16	na	na	0.50	0.39	0.42	0.02
Active Memory (ln mb)							
Desktop	0.32	0.45	0.58	0.25	0.34	0.31	0.18
Mobile	0.23	na	na	0.29	0.18	0.13	0.22

MEAN ELASTICITY ESTIMATES BASED ON NONLINEAR MODEL WITHOUT PROCESSORS

	Pooled Time	Means from Adjacent Year Regressions					
Characteristic	Regression	76-79	80-83	84-89	90-93	94-97	98-03
Hard disk memory (ln mb)							
Desktop		na	na	na			
Mobile		na	na	na			
Processor speed (ln mhz)							
Desktop		na	na	na			
Mobile		na	na	na			
Active Memory (ln mb)							
Desktop		na	na	na			
Mobile		na	na	na			

Notes: Based on authors' regression analysis. na is not applicable.

Table 2

Alternative Price Indexes Adjusted for Quality Change (1992 = 1.0000), based on Parsimonious OLS Linear Logarithmic Regression Equations

(1)	(2)	(3)	(4)	(5)	(6)	(7)	
YEAR	ALL POOLED	DESKTOP POOLED	MOBILE POOLED	DESKTOP ADJACENT YEARS	DESKTOP YEARLY FISHER	MOBILE ADJACENT YEARS	MOBILE YEARLY FISHER
1976	20.1598	24.8562	na	87.9878		na	na
1977	12.3905	14.9977	na	57.4363		na	na
1978	10.9771	13.0663	na	49.4985		na	na
1979	11.7569	13.7718	na	43.0944		na	na
1980	12.1483	14.2093	na	42.9269		na	na
1981	12.6142	14.4325	na	39.2055		na	na
1982	10.4603	11.8157	na	30.5174		na	na
1983	7.6340	8.8992	3.7427	18.0701		6.1110	
1984	7.3513	8.2526	4.0367	14.4211		7.5455	
1985	5.1198	5.7581	3.3295	9.7985		5.3717	
1986	3.6889	4.0526	2.6656	6.7183		3.4912	
1987	2.8816	3.1460	2.1934	4.9499		2.6666	
1988	2.3712	2.4673	1.9958	3.4056		2.1838	
1989	2.5051	2.6766	2.3252	3.3786		2.1740	
1990	2.2165	2.2701	2.0245	2.5606		1.9081	
1991	1.9793	2.0782	1.8836	2.2905		1.7720	
1992	1.0000	1.0000	1.0000	1.0000		1.0000	
1993	0.7813	0.7166	1.0467	0.7641		1.0323	
1994	0.4638	0.4088	0.7364	0.5121		0.8278	
1995	0.3448	0.2994	0.5678	0.3632		0.5244	
1996	0.2355	0.1881	0.4894	0.1984		0.3685	
1997	0.1619	0.1326	0.3536	0.1358		0.2498	
1998	0.0715	0.0519	0.1977	0.0495		0.0665	
1999	0.0510	0.0327	0.1563	0.0330		0.0527	
2000	0.0378	0.0251	0.1231	0.0252		0.0449	
2001	0.0228	0.0161	0.0750	0.0172		0.0292	
2002	0.0149	0.0098	0.0545	0.0136		0.0229	
2003	0.0111	0.0092	0.0417	0.0115		0.0180	
AAGRs							
76-03	-24.27	-26.03	na	-28.19		na	na
83-03	-27.87	-29.93	-20.14	-30.78		-25.28	
76-83	-12.95	-13.65	na	-20.24		na	na
83-88	-20.85	-22.63	-11.82	-22.82		-18.60	
88-93	-19.91	-21.91	-12.11	-25.84		-13.92	
93-98	-38.02	-40.85	-28.35	-42.15		-42.21	
98-03	-31.11	-32.55	-26.75	-25.31		-23.01	

Notes: Calculations based on authors' regression models. na is not applicable.

Table 3
Alternative Price Indexes Adjusted for Quality Change (1992 = 1.0000)
Based on Loglinear Regression Equations with Processor and Clock Speed-Processor Interactions Variables Included (Expanded Specifications)

YEAR	(1) ALL POOLED	(2) DESKTOP POOLED	(3) MOBILE POOLED	(4) DESKTOP ADJACENT YEARS	(5) DESKTOP YEARLY FISHER	(6) MOBILE ADJACENT YEARS	(7) MOBILE YEARLY FISHER
1989				2.9443		1.6800	
1990				2.2665		1.5724	
1991				2.0034		1.4693	
1992				1.0000		1.0000	
1993				0.7731		1.0176	
1994				0.5322		0.8812	
1995				0.3813		0.5722	
1996				0.2058		0.4100	
1997				0.1497		0.2657	
1998				0.0621		0.0779	
1999				0.0460		0.0601	
2000				0.0365		0.0517	
2001				0.0273		0.0384	
2002				0.0228		0.0351	
2003				0.0221		0.0281	
AAGRs							
89-03				-29.48		-25.34	
89-93				-28.42		-11.78	
93-98				-39.62		-40.19	
98-03				-18.63		-18.44	

COMPARABLE AAGR'S FROM TABLE 2

AAGRs							
89-03				-33.29		-29.00	
89-93				-30.78		-16.99	
93-98				-42.15		-42.21	
98-03				-25.31		-23.01	

Notes: Calculations based on authors' regression models. na is not applicable.

Figure 1

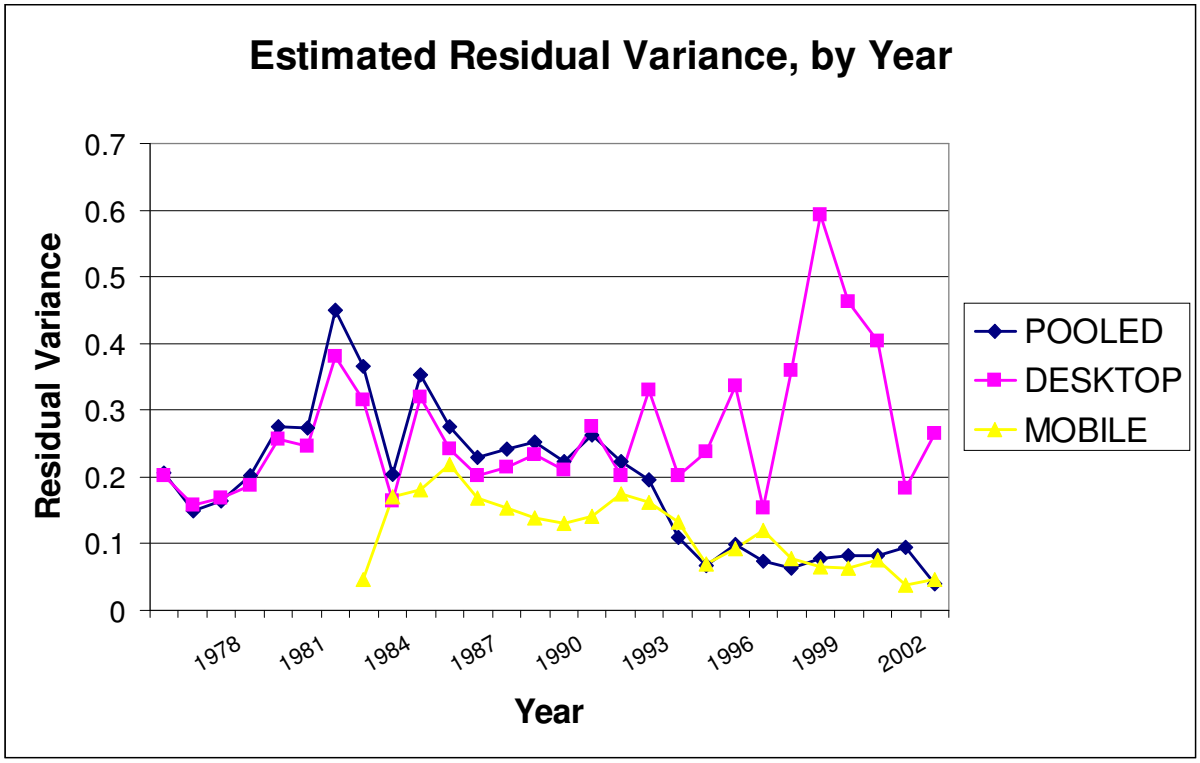
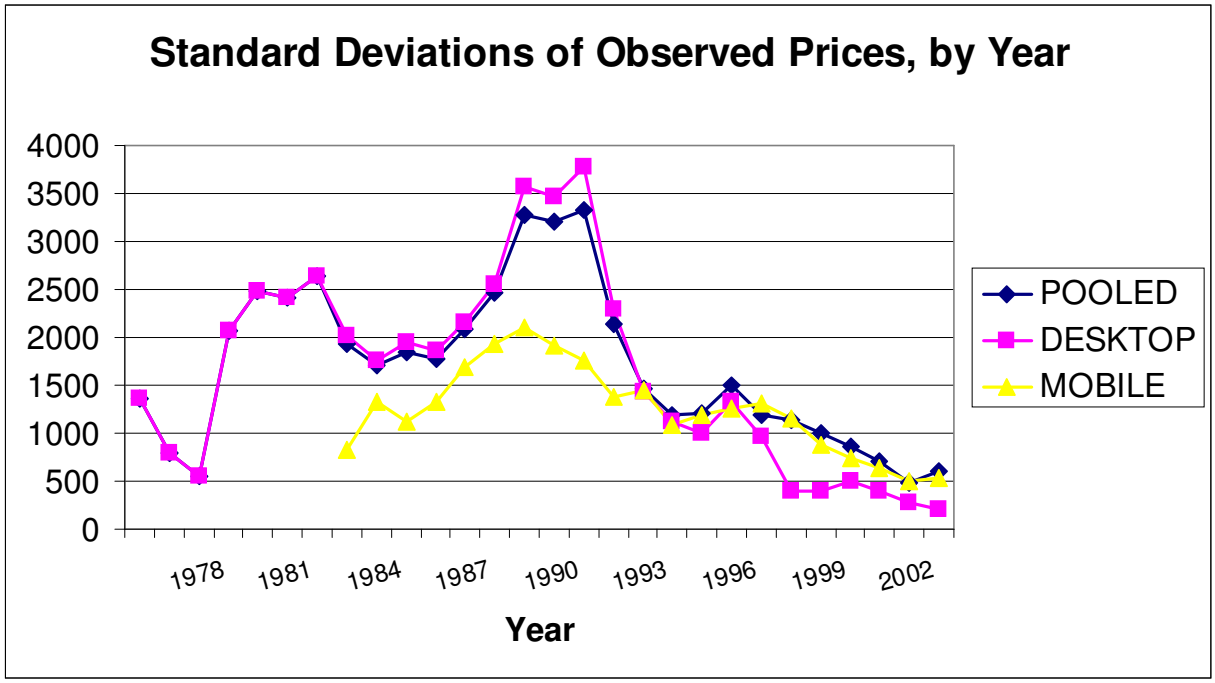


Figure 2



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APPENDIX TABLE 1

MEAN PRICES AND VALUES OF SELECTED TECHNCIAL CHARACTERISTICS VARIABLES ANNUALLY, 1976-2003, SEPARATELY FOR DESKTOP AND MOBILE MODELS

YEAR	NUMBER OF OBSERVATIONS	DOLLAR PRICE	HARD DISK CAPACITY-MB	RANDOM ACCESS MEMORY - MB	PROCESSOR SPEED - MHZ
			DESKTOP MODELS ONLY		
1976	10	1898	0	0.00975	1.35
1977	16	1364	0	0.01197	1.94
1978	14	1154	0	0.01043	2.14
1979	26	1957	0.42	0.02162	2.29
1980	36	2126	0.61	0.02656	1.94
1981	35	2372	0.63	0.03246	2.46
1982	44	2531	0.73	0.05973	2.62
1983	65	2821	1.65	0.10960	4.17
1984	111	3040	3.64	0.17845	5.19
1985	92	3034	6.40	0.31443	5.89
1986	156	2757	10.19	0.48821	7.21
1987	324	2916	21.13	0.71086	9.43
1988	296	2965	39.96	0.93603	12.64
1989	168	4634	50.39	1.26619	15.71
1990	448	4420	61.45	1.79232	19.42
1991	316	4440	55.05	2.07106	21.21
1992	501	3415	133.34	5.12176	29.09
1993	891	2489	168.97	4.00224	33.80
1994	308	1965	838.21	5.26623	51.92
1995	473	2155	476.98	8.60888	78.15
1996	199	2590	1465.45	18.55276	145.64
1997	359	2031	1942.84	21.27159	182.55
1998	209	1466	4954.07	51.13876	330.51
1999	603	1357	7724.54	80.98176	449.19
2000	1720	1384	12020.51	99.57209	651.40
2001	496	1200	20026.61	127.35480	986.48
2002	715	1078	33234.97	191.41820	2121.04
2003*	201	964	36815.92	219.06470	2285.62

* Q1 only

MOBILE MODELS ONLY

YEAR	NUMBER OF OBSERVATIONS	DOLLAR PRICE	HARD DISK CAPACITY-MB	RANDOM ACCESS MEMORY - MB	PROCESSOR SPEED - MHZ
1983	9	1961	3.00	0.10400	4.11
1984	16	2460	1.00	0.12950	3.70
1985	22	2159	1.86	0.16473	4.24
1986	32	2489	3.91	0.35125	5.62
1987	73	2442	7.84	0.44909	7.42
1988	58	2807	14.47	0.72166	9.87
1989	52	4616	29.10	1.14508	13.15
1990	117	4336	33.68	1.25494	14.40
1991	126	4083	33.49	1.25635	14.88
1992	125	3295	83.12	2.97312	22.75
1993	260	3338	86.59	2.56871	22.62
1994	67	3109	150.57	5.61194	32.39
1995	266	3314	364.53	5.86466	65.92
1996	122	4104	1009.67	11.60656	114.64
1997	79	3490	1383.92	15.64557	137.32
1998	131	3176	4445.04	45.06870	257.66
1999	653	2783	5556.70	56.53553	331.59
2000	1464	2585	8049.04	69.46448	534.23
2001	884	2070	14108.61	113.52040	807.06
2002	126	2073	25698.41	215.36510	1397.97
2003*	200	1861	33600.00	308.64000	1800.09

* Q1 only