The Rise of Goods-Market Competition and The Fall of Nominal Wage Contracting:

Endogenous Wage Contracting in a Multisector Economy

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THE RISE OF GOODS-MARKET COMPETITION AND

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ENDOGENOUS WAGE CONTRACTING IN A MULTISECTOR ECONOMY*

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Abstract

This paper shows how heterogeneity in wage-setting and a link between nominal wage flexibility and goods-market competition arise in a multisector economy that is affected by aggregate and sector-specific shocks. Aggregate volatility increases the variance of real contract wages, whereas sectoral volatility increases the relative variance of real Walrasian wages. Given this tradeoff, the prevalence of nominal wage contracting reflects both the relative volatility of aggregate versus sectoral disturbances and the overall degree of goods-market market competition. We find that these variables help explain the decline in unionization (a proxy for contracting) in the United States.

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I. Introduction

Spurred by the seminal papers by Fischer (1977a) and Gray (1976), a voluminous literature has explored the implications of nominal wage contracts established by rational workers and firms. Yet relatively few contributions have resolved the fundamental question first debated by Barro (1977) and Fischer (1977b): Why would workers and firms in some portion of the economy adopt contracts that fix nominal wages and thereby move these agents away from potentially mutually advantageous opportunities to improve their welfare? Karni (1983), Duca and VanHoose (1991), and others have relied on theoretical models in which wages are indexed to output or to profits to motivate the potential optimality of nominal wage contracting. Although forms of implicit or explicit indexation, such as profit sharing, undeniably are important features of many contracts, the rational contracting literature continues to lack a clear theoretical explanation for why nominal wage contracting occurs and why its incidence has declined so sharply since the early 1980s.

This paper offers one possible, and surprisingly simple, explanation for the observation of nominal wage contracting. Although it differs in its specific approach, this explanation relates to earlier contributions by Gray (1983) and Woglom (1990), which focused on industry-specific factors that influence the optimal indexation of wage contracts. We find that nominal contracts are likely to be preferable to relying on market-clearing outcomes if sectoral disturbances are sufficiently variable in relation to aggregate shocks. Agreeing to fix the nominal wage over an interval exposes workers to real-wage volatility arising from aggregate price fluctuations that stem from economywide demand and supply disturbances. Nonetheless, such an agreement insulates workers from real-wage variability resulting from sectoral shocks.

This intuitive explanation provides the foundation for a theory for the determination of the equilibrium portion of sectors that choose to adopt nominal wage contracts. In the context of a stylized multisector model, we are able to examine key parameters that influence the extent to which workers and firms choose to contract. These include not only the variances of aggregate shocks, but also the variance of sectoral shocks and the degree of goods-market competition. Using measures of the extent of contracting and the explanatory variables implied by the theory, we find evidence supporting the theory's implication that sectoral volatility and goods-market competition play key roles in determining the degree of contracting in the U.S. economy and in explaining why contracting has fallen since the mid-1950s. Note that most of the drop in unionization since the early 1980s is due to declines in unionization rates within industries, rather than to shifts in employment across industries (see Table 1).

Table 1 Goes Here

The next section of the paper develops our model and uses it to demonstrate the conditions under which nominal wage contracts can be optimal. Section III extends this model to construct our theory of the determination of the equilibrium portion of the economy that opts to contract. Section IV provides the empirical analysis. Section V offers some concluding thoughts about our model's implications.

II. The Model and Its Solution

The model parallels the multisector framework developed by Duca and VanHoose (1997, 1998). It extends Ball's (1988) framework by incorporating some features from the Duca-VanHoose (1991) multisector model. The latter actually is a two-sector extension of Duca (1987), whereas the present model truly allows for multiple sectors.[1] Specifically, in the present framework there is a continuum of sectors, indexed j, that are distributed uniformly across a unit interval. Each sector is occupied by large numbers of representative firms and workers. All behavioral relationships are expressed in terms of logarithms, and most constants are suppressed in the exposition to simplify the presentation of the model and its key implications.

The production function for a firm in sector j is

(1) $y_j = \alpha l_j + \theta,$

where y_j is the log of firm j's output, l_j is the log of employment at firm j, and θ is a common, zero-mean productivity shock with finite variance σ_θ^2. The demand for firm j's output as a share of aggregate output is given by

(2) $y_j - y = -\varepsilon(p_j - p) + \delta_j,$

where $y \equiv \int_0^1 y_j\, dj$ is aggregate output and $p \equiv \int_0^1 p_j\, dj$ is the aggregate price level; δ_j is a firm-specific zero-mean demand shock with a finite variance (discussed in more detail in section III); and $\varepsilon > 1$ is the elasticity of demand for the output of firm j. For the sake of expositional clarity, δ_j is the single sectoral disturbance faced by a firm. It is straightforward to show that including sectoral supply shock greatly complicates expressions for worker losses that we consider below without fundamentally changing comparisons of the relative effects of sectoral versus aggregate shocks.

Aggregate demand for output is characterized by the quantity equation,

(3) $y = m + v - p,$

where v is a zero-mean aggregate demand disturbance with finite variance variance σ_v^2. All three types of shocks (θ, δ_j, and v) are assumed to be i.i.d.

Converting (1), (2), and (3) into levels (denoted by upper-case letters) and combining with the profit function $\pi_j = P_jY_j - W_jL_j$ yields the following labor demand function (with the intercept term suppressed):

$$(4) \qquad l_j^d = \frac{-\varepsilon\,(w_j - p) + (m + v + \delta_j - p) + (\varepsilon - 1)\theta}{\alpha + \varepsilon - \alpha\varepsilon},$$

where w_j is the log of the nominal wage at a firm in sector j. Each firm acts as a perfect competitor in its labor market, in which it faces a pool of sectorally immobile workers whose labor supply schedule is given by

$$(5) \qquad l_j^s = \lambda(w_j - p),$$

where $\lambda > 0$. As in Duca (1987) and Duca and VanHoose (1991) [also see Carlson and Findlay (1992)], the crucial aspect of this stylized structure is that firms value the real wages they pay in terms of the prices of the products that they produce, while workers value the real wages that they earn in terms of the aggregate price level.[2]

A fraction, Ω, of the sectors have workers and firms that use nominal wage contracts. The remaining share, 1-Ω, do not. For both groups of sectors, the full-information, market-clearing wage equates (4) and (5) and is given by

$$(6) \qquad w_j^* - p = \frac{(m + v + \delta_j - p) + (\varepsilon - 1)\theta}{\lambda\,(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon}.$$

This is the wage actually paid by a firm in sector j if its sector is among the share 1-Ω of sectors that do not use contracts. For such a firm, denoted "*nc*," substitution of (6) into either (4) or (5) yields the employment level,

$$(7) \qquad l_j^{nc} = \frac{\lambda\,[(m + v + \delta_j - p) + (\varepsilon - 1)\theta\,]}{\lambda\,(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon}.$$

At a contracting firm, however, nominal wage contracts are set to satisfy $w_j = Ew_j$

= 0. Using this in (4) yields the employment level at a firm in a contracting sector:

$$(8)\qquad l_j^c = \frac{\varepsilon p + (m + v + \delta_j - p) + (\varepsilon - 1)\theta}{\alpha + \varepsilon - \alpha\varepsilon}.$$

Using (8) in (1) implies that output at such a contracting firm is equal to

$$(9)\qquad y_j^c = \frac{\alpha\varepsilon p + \alpha(m + v + \delta_j - p) + \varepsilon\theta}{\alpha + \varepsilon - \alpha\varepsilon}.$$

Substituting (9) into (2) yields an expression for the jth contracting sector's price in terms of the quantity of money, the aggregate price level, and the shocks:

$$(10)\qquad p_j^c = \frac{(1-\alpha)(m + v) - \theta + [(\alpha + \varepsilon - \alpha\varepsilon) - 1]p}{\alpha + \varepsilon - \alpha\varepsilon}.$$

For a noncontracting firm, substituting (7) into (1) yields the output expression,

$$(11)\qquad y_j^{nc} = \frac{\alpha\lambda(m + v + \delta_j - p) + \varepsilon(1+\lambda)\theta}{\lambda(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon}.$$

Substituting this result into (2) yields the sectoral price,

$$(12)\qquad p_j^{nc} = \frac{[\lambda(1-\alpha)+1](m + v + \delta_j) - (1+\lambda)\theta + \{\lambda(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon] + [\lambda(1-\alpha)+1]\}p}{\lambda(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon},$$

Our key assumption for aggregating output and prices is that the sector-specific shocks "wash out" when summing across firms within each group: $\int_0^{\Omega} \delta_j \, dj = 0$ and $\int_{\Omega}^{1} \delta_j \, dj = 0$ for all Ω. Under this assumption, the aggregate price level is given by $p = \Omega p_j^c|_{\delta_j=0} + (1-\Omega)p_j^{nc}|_{\delta_j=0}$. Using (10) and (12), this yields

$$(13)\qquad p = \frac{[\lambda(1-\alpha)+1](\alpha + \varepsilon - \alpha\varepsilon) - \alpha\Omega](m + v) - [(1+\lambda)(\alpha + \varepsilon - \alpha\varepsilon) + \alpha(\varepsilon - 1)\Omega]\theta}{\Omega[\lambda(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon] + (1-\Omega)(\alpha + \varepsilon - \alpha\varepsilon)[\lambda(1-\alpha)+1]}.$$

Equation (13) permits us to calculate the real wages faced by workers in individual sectors. For a worker in the jth contracting sector, the real wage is

$$(14)\quad w_j^c - p = \; - p = \frac{\{\alpha\Omega - (\alpha+\varepsilon-\alpha\varepsilon)[\lambda(1-\alpha)+1]\}(m+\nu) + [(1+\lambda)(\alpha+\varepsilon-\alpha\varepsilon) + \alpha(\varepsilon-1)\Omega]\theta}{\Omega[\lambda(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon] + (1-\Omega)(\alpha+\varepsilon-\alpha\varepsilon)[\lambda(1-\alpha)+1]},$$

while for a worker in the jth noncontracting sector, the real wage is

$$(15)\quad w_j^{nc} - p = \frac{\alpha\Omega(m+\nu) + [(\alpha+\varepsilon-\alpha\varepsilon) + \alpha(\varepsilon-1)\Omega]\theta}{\Omega[\lambda(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon] + (1-\Omega)(\alpha+\varepsilon-\alpha\varepsilon)[\lambda(1-\alpha)+1]} + \frac{\delta_j}{\lambda(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon}$$

An immediate implication of (14) and (15) is that workers who agree to use nominal wage contracts remove sector-specific shocks as a source of real-wage volatility. Only aggregate shocks influence the real, *ex post* contracting wage. In contrast, workers in noncontracting sectors expose themselves to real-wage volatility. On average, of course, the values of the aggregate and sector-specific shocks are equal to zero, given the simplified structure of the model. Consequently, a risk-averse worker choosing between contractual or Walrasian wage setting must assess the potential increase or reduction in real-wage volatility in light of exogenous sources of aggregate and sector-specific instability.[3]

Differentiating (14) and (15) with respect to $(m + \nu)$ yields

$$|\partial(w_j^{nc} - p)/\partial(m+\nu)| = |D^{-1}\alpha\Omega|$$

and

$$|\partial(w_j^{c} - p)/\partial(m+\nu)| = |D^{-1}\{\alpha\Omega - [\lambda(1-\alpha)+1](\alpha+\varepsilon+\alpha\varepsilon)\}|,$$

where D is the denominator in (14) and (15). Comparing these expressions indicates that a sufficient condition for $|\partial(w_j^{nc} - p)/\partial(m+\nu)| < |\partial(w_j^{c} - p)/\partial(m+\nu)|$ to hold unambiguously is $\Omega < 1/2$. This inequality can hold for a large range of parameter values for higher values of Ω, however. Hence, for many possible parameter values, the real wage earned by a worker at a firm with a nominal wage contract is more volatile in the face of exogenous monetary policy innovations and aggregate demand disturbances. Analogous comparisons indicate that $|\partial(w_j^{nc} - p)/\partial\theta| < |\partial(w_j^{c} - p)/\partial\theta|$

and that $|\partial(w_j^{nc} - p)/\partial\delta_j| > |\partial(w_j^{c} - p)/\partial\delta_j| = 0$, where these additional comparisons are unambiguous irrespective of the magnitude of Ω. Consequently, the real wage of a worker in a contracting sector is also more volatile in the face of an aggregate productivity shock. By definition, however, the real wage in a sector with nominal wage contracts is unaffected by sector-specific disturbances.

These comparisons imply that, from the perspective of a risk-averse worker whose utility declines with greater real wage volatility, the choice between nominal wage contracts and Walrasian wage setting hinges on the relative importance of aggregate versus sector-specific shocks. As the variances of aggregate disturbances increases relative to the variance of sectoral shocks, workers in more sectors will tend to prefer nominal wage contracts, *ceteris paribus*.

III. Determining the Equilibrium Contracting Share

According to the theory, what portion of sectors in the economy will rationally adopt nominal wage contracts? To address this question, we begin by making three final assumptions. First, monetary policy is determined exogenously, with the normalization $m = 0$. Second, firms are risk-neutral.[4] This leaves the decision about the type of wage contract to the worker, given that both contracts yield the same mean real wage, employment, output, and profit.

Finally, we assume that the variance of sector-specific shocks is equal to $(1-j)j^{-1}\sigma_\delta^2$, with $\sigma_\delta^2 \geq 0$. Because sectors are distributed uniformly over the interval [0, 1], the variances of sector-specific disturbances decline from an infinite value for sector zero to a value of zero for sector 1. This implies that workers in sector zero always prefer to adopt nominal wage contracts for $\sigma_\delta^2 > 0$ and that workers in sector 1 never wish to contract. At a sector located halfway along the unit interval, the sectoral variance is exactly equal to σ_δ^2; whether or not workers in this sector or in sectors to the right or left of it will contract depends on relative real wage variances.

Suppose that sector j^* is the sector for which workers are indifferent between nominal wage contracts and Walrasian wages. For the next sector, therefore, it must be the case that real-wage volatility under a nominal wage contract would exceed real-wage volatility that would arise without such a contract. Firm j^*, therefore, is the "critical sector" along the unit interval for which a point of indifference is reached concerning adoption of a nominal wage contract.

We assume that workers seek to minimize a weighted average of expected deviations of the real wage and employment from their Walrasian, market-clearing values, or

$$(16)\quad L_j = E\,[(l_j - l_j^*)^2 + \Gamma(w_j - p)^2].$$

where l_j^* is the Walrasian, full-information level of employment (that is, the employment level that each firm would attain under full information if no firms contracted), given by

$$(17)\ l_j^c = \frac{-\lambda}{\lambda\,(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon} \times \left([\lambda\,(\alpha+\varepsilon-\alpha\varepsilon) + (\varepsilon-1)](m + v) + \frac{\lambda[(\varepsilon-1)[(1-\alpha)\lambda+1]+(1+\lambda)\lambda\,(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon]}{[\lambda\,(\alpha+\varepsilon-\alpha\varepsilon)+\varepsilon][(1-\alpha)\lambda+1]} \right) + \frac{\lambda\,\delta_j}{(\alpha + \varepsilon - \alpha\varepsilon) + \varepsilon},$$

It follows that $L_j^{nc} \geq L_j^{\,c}$ for workers at all sectors in the interval $[0,\ j^*]$. Hence, j^* defines the cutoff point beyond which workers at additional firms choose not to contract, with $j^* = \Omega$ in equilibrium.

After computing these expressions from (14) and (15) and rearranging, we find that the equilibrium portion of sectors that contract solves the following equation:

$$
\begin{aligned}
(1-\Omega)/\,\Omega = {} & \Big\{(A^2 - \Gamma\varepsilon^2)B^2C^2[(1-\Omega)AC+\Omega B]^2\sigma_\delta\Big\}^{-2} \\
& \times \Big\{A^2B^4C^2\Big[[(\alpha\Omega-AC)^2-\Omega^2\alpha^2]\sigma_v^2 + \{[(1+\lambda)A+\alpha\Omega(\varepsilon-1)]^2-[A+\alpha\Omega(\varepsilon-1)]^2\}\sigma_\theta^2\Big] \\
& + (1/\Gamma)\Big[A^2B^2C^2\Big(B^2\varepsilon\, C[BC^2\varepsilon+2\lambda\alpha\Omega] +[AC(1-\Omega)+B\Omega]\{\lambda\alpha\Omega[1-[\lambda A+(\varepsilon-1)]+\varepsilon C\} \\
& \qquad +\lambda^2[AC(1-\Omega)+B\Omega]^2\{[1-[\lambda A+(\varepsilon-1)]^2\}\Big)\sigma_v^2 \\
& \qquad + BC\Big(BC\lambda^2[\alpha^2(\varepsilon-1)^2(1-2\Omega)-A^2] \\
& \qquad +2\{\lambda^2\alpha(\varepsilon-1)(1-\Omega)[AC(1-\Omega)+B\Omega][AC(\varepsilon-1)+(1+\lambda)B][\alpha(\varepsilon-1)+A]-\lambda^2A\alpha\Omega(\varepsilon-1)\}\Big)\sigma_\theta^2\Big]\Big\},
\end{aligned} \tag{18}
$$

where $A \equiv \alpha+\varepsilon+\alpha\varepsilon$, $B \equiv \lambda(\alpha+\varepsilon+\alpha\varepsilon)+\varepsilon$, and $C \equiv \lambda(1-\alpha)+1$. If Γ is sufficiently large, so that workers place a sufficiently high weight on reducing real-wage volatility, equation (18) indicates that as $\sigma_v^2/\sigma_\delta^2 \to \infty$ or $\sigma_\theta^2/\sigma_\delta^2 \to \infty$ (that is, as the volatility of aggregate disturbances dominates the volatility of sectoral shocks), then $\Omega \to 0$, and the equilibrium share of sectors that contract approaches zero. In contrast, as $\sigma_v^2/\sigma_\delta^2 \to 0$ or as $\sigma_\theta^2/\sigma_\delta^2 \to 0$ (that is, as the volatility of sectoral disturbances overwhelms the volatility of aggregate shocks), then $\Omega \to 1$, and *all* workers contract. Therefore, as discussed above, the model predicts that a smaller portion of sectors will use wage contracts as the variability of aggregate disturbances increases. Greater sectoral volatility, however, leads to a larger share of contract sectors over a significant range of parameter values.

The direction of the effect of higher relative sectoral variability on the equilibrium extent of contracting is not monotonic, however. In the limiting case in which $\Gamma \to 0$, for instance, greater sectoral variability provides a disincentive for firms to contract, because higher aggregate *or* sectoral disturbances push equilibrium employment farther from its full-information, market-clearing level. Thus, there is a critical value of Γ (expressed in terms of other parameters of the model) below which greater sectoral volatility *reduces* the equilibrium portion of sectors with wage contracts.[5] On net, therefore, the model leaves open the net direction of the effect of a small change in sectoral volatility on the equilibrium contract share.

Equation (18) is a cubic polynomial in Ω. Nevertheless, we can rearrange (18) and implicitly differentiate to find that $\partial\Omega / \partial\varepsilon < 0$ for sufficiently small values of Ω, so that a relatively small portion of sectors contract. A further requirement for $\partial\Omega / \partial\varepsilon < 0$ is that Γ must be sufficiently large, so that workers place sufficiently large weight on the real wage component of their loss function. [6] Consequently, for a large range of parameter values the model indicates that the equilibrium share of sectors that use wage contracts will decline as the degree of goods-market competition increases.

The intuition behind this result as follows. A rise in the elasticity of demand caused by an increased degree of goods-market competition reduces the responsiveness of a firm's equilibrium price to a sectoral disturbance. This results in a decline in the sensitivity of the marginal revenue product to sectoral shocks. Therefore, the positions of firms' labor demand schedules and their employment choices respond less completely to sector-specific shifts in the output demand schedules that they face. The result is a smaller change in the nominal wage. Because sectoral shocks leave the aggregate price level unaffected, the real wage is unambiguously less volatile in the face of sectoral shocks when the demand elasticity increases.

Aggregate demand shocks also induce less employment volatility for both contracting and noncontracting firms when goods markets are more competitive. For most ranges of parameter values, however, the contribution of common demand disturbances to aggregate price volatility rises, thereby adding to variability of the real wage at a contracting firm. As long as real wage variance is a significant factor in workers' loss valuations, therefore, greater aggregate demand variability in the presence of increased goods-market competition tends to reduce the extent of nominal wage contracting.

Aggregate productivity shocks contribute directly to labor demand variability, which causes greater employment variability at both type of firms. The real wage effects of aggregate supply disturbances differ, however. At noncontracting firms, a negative productivity shock reduces the equilibrium nominal wage while raising the aggregate price level; at contracting firms such shocks affect the real wage only through the price-level response. Under increased goods-market competition, an aggregate supply shock induces a smaller price response at firms, as discussed above, and so the sensitivity to of the value of labor's marginal product to such a shock declines. Consequently, the effect that a productivity disturbance has on labor demand is reduced with a higher elasticity of demand, making the nominal wage at noncontracting firms less responsive to aggregate supply disturbances. The result is that greater goods-market competition reduces the real wage volatility that is induced by supply shocks at noncontracting firms relative to contracting firms. *Ceteris paribus*, this leads to a reduction in the equilibrium share of firms that use wage contracts.

IV. Empirical Analysis

Using two-step cointegration methods, we test our model's implications that contract use declines if the degrees of goods-market competition, aggregate demand shock variance, or aggregate supply variance rise, or, for low contracting economies like the United States, if the variance of sector-specific shocks falls. We first test for cointegration among contract share, inflation risk, the degree of goods-market competition, aggregate supply variance, and the variance sectoral nominal output growth. Then we perform second-stage regressions of the change in the log of contract share.

A. Data and Variables

The variables used fall into five categories: contract share, aggregate demand variance, supply shock variance, goods-market competition, and sectoral variance.

Contract Share Given the absence of comprehensive measures of the share of U.S. workers under contracts, we use data on unionization rates in the private sector. Although many workers have annual pay adjustments, we are more interested in contracts longer than a year, because such contracts expose agents to a greater risk of expectational errors, which play a key role in determining contract share in our model. Union contracts generally exceed a year in length, and for this reason unionization rates appear to be a good proxy for contract share (see Figure 1).

Figure 1 Goes Here

Contract share is measured by splicing annual data from the Bureau of Labor Statistics (BLS) and Troy and Sheflin (1985) on private sector unionization rates, including employee association members (see Appendix A). Our theory, of course, relates more directly to nominal wage contracting than to unionization. There is, of course, a broad literature that assesses determinants of labor membership, including rent-sharing and other factors that can contribute to a potential linkage between the extent of unionization and the degree of goods-market competition [see, for instance, Ashenfelter and Pencavel (1969), Oswald (1982), Rau (1985), Neumann and Rissman (1984), and Mason and Bain (1993)]. Consequently, our theoretical model's macroeconomic-based prediction complements existing microeconomic rationales for a channel from goods-market competition to unions and longer-term contracting. One difference is that our multisector macroeconomic model predicts that the variances of sectoral and aggregate disturbances are additional key determinants of the extent of nominal wage contracting via a mechanism such as unionization.

Aggregate Demand Variance In our model, aggregate demand variance reduces the incentive to contract mainly by introducing inflation risk, which we proxy using annual CPI inflation (INF).[7] Holland (1995) uses the level of inflation to measure inflation risk on grounds that prior studies [e.g., Evans (1991)] have found that inflation uncertainty is increasing in the level of inflation.

Aggregate Supply Shock Variance Aggregate supply shock variance (OILVAR) is proxied by the 3-year moving average of the squared percent change in real consumer energy prices [using (CPI energy/CPI) in levels back to 1957 and (PCE gas and oil/PCE fixed weight deflator) in levels prior, with a break adjustment (see Appendix B).]

Goods-Market Competition The degree of goods-market competition (ε) is measured by the inverse of a measure of the aggregate markup. The raw markup [(price divided by marginal cost) minus 1] is measured by the ratio of after-tax profits to GDP for nonfinancial corporate businesses. This markup is adjusted for cyclical and short-run effects by subtracting off estimated impacts of the unemployment rate, oil prices, and the real exchange rate. The long-run, cyclically adjusted degree of competition equals the inverse of this adjusted markup (see Appendix B). Figure 2 shows that ε was low over the 1960s and 1970s but moved into a higher range in the 1980s and 1990s, coinciding with waves of deregulation in the transportation, energy, and communications industries and with stiffer foreign competition for manufacturers.

Figure 2 Goes Here

Sectoral Variance Shifts in relative sectoral output (SSHIFT) are proxied by shifts in the share of nominal output produced by different U.S. industries, defined as the sum of squared percentage-point changes in industry output shares according to 2-digit SIC data. (Using shares implicitly controls for relative sizes of particular industries, given that data for real industry output levels are unavailable for most of the 1950s — see Appendix C for details.) Data were carefully adjusted for shifts in category definitions and other sample breaks. To capture more medium-run swings in the importance of relative shocks, we use the three-year moving average of SSHIFT (SSHIFT3), which is plotted in Figure 3. Compared with using Lillien (1982)-type dispersion measures of employment across industries, our variable more closely corresponds to our theoretical model and avoids the labor-hoarding-type effects stemming from delays by firms in adjusting employment. (The inclusion of the aggregate supply and demand variance terms in the log cointegrating vectors implicitly controls for oil and cyclical influences while allowing us to use a log specification.) In other first- and second-stage runs, we obtained qualitatively similar results when, following the example of Otoo (1994), we first adjusted SSHIFT for the estimated impact of real oil price variance and deviations of output from trend (we used the gap between actual and DRI's estimate of equilibrium unemployment). To conserve space, we do not report these results in the tables.

Figure 3 Goes Here

Unfortunately, our empirical measure of sector-specific variance reflects both sector-specific demand and supply shocks. The theoretical model implies that workers who place relatively high weight on reducing real-wage volatility are more likely to prefer to contract in the presence of significant sector-specific *demand* variance. An expanded version of the model that includes sectoral *supply* variances indicates that an increase in the variance of sector-specific supply shocks has an ambiguous effect, which is consistent with the nonmonotonic relationship implied by our theoretical framework. Hence, the model has ambiguous predictions for the sign on our empirical measure of sectoral variance.

B. First-Stage Cointegration Results

The theoretical model implies that the use of contracts is related to economic variables in a nonlinear manner. Consequently, we use log specifications to test our hypotheses. Because the log level of union share is I(2) and other long-run variables are I(1), the first-stage cointegration tests use the first difference of the log of union share with logs of other long-run determinants. For this reason, the second-stage error-correction models use ΔΔlog(union share) as the dependent variable. Table 2 reports results for the most significant cointegrating vector for selected combinations of long-run variables. The Johansen-Juselius (1990) procedure is used, partly because annual data precluded using the Stock and Watson DOLS approach, which consumes many degrees of freedom. Results for three combinations of long-run variables that each include Δlog(union share) are presented: (1) INF and OILVAR (the "conventional model"); (2) INF, OILVAR, and SSHIFT3 (the "noncompetition model"); and (3) INF, OILVAR, SSHIFT3, and ε (the "competition model"). In each case, there was only one vector that was significant at the 5 percent level according to rank significance. We find in each instance that higher sectoral variance tends to reduce the extent of contracting via unionization, suggesting that our sectoral variable reflects sector-specific supply shock variance (or other factors) more than sector-specific demand shock variance.[8] In addition, as noted earlier, even though the existence of sectoral demand shocks can create an incentive to use nominal wage contracts, a marginal change in SSHIFT3 on net has a theoretically ambiguous effect on the extent of contracting — especially for countries, like the U.S., that have low shares of contracting/unionization. For this reason, it is plausible to observe a negative effect of SSHIFT3 on U.S. unionization and a higher degree of unionization in sectors that face greater sectoral demand volatility, such as manufacturing, energy (mining), *construction, and airlines* (transportation) — see Table 1.

Table 2 Goes Here

However, only in the fully specified model that includes the sectoral variance and competition variables are the signs of the estimated cointegration consistent with contract theory. In particular, in the other models the cointegrating vectors imply a positive relationship between inflation and unionization. One plausible explanation for this stems from the fact that inflation and the measure of competition tend to be negatively related over the long run. Hence, omitting ε inadvertently mixes the positive effect of low competition on unionization with the negative impact of high inflation on the incentive to contract.

C. Second-Stage Error-Correction Results

As noted earlier, because the log level of union share is I(2), the second-stage error-correction models use ΔΔlog(union share) as the dependent variable with error-correction terms based on cointegrating relationships from Table 2. Table 3 presents six models, the first three of which have no short-run terms but differ according to which cointegrating vector from Table 2 was used to define the error-correction term. The other three models correspond to models 1, 2, and 3 in terms of which error-correction term is used but include two short-run variables — namely, the t lag of the change in aggregate supply shock variance (ΔOILVAR) and a dummy for recessions (RECESS) which equals 1 in years containing any NBER-recession quarters. Of the first differences of the various long-run variables, only the contemporaneous first difference of OILVAR proved to be a statistically significant short-run variable. The recession dummy was tested because during recessions, Δlog(union share) tends to be higher, plausibly reflecting that union share tends to be bolstered in recessions, during which larger and more unionized establishments are more likely to survive and during which larger firms tend to cut back on subcontracting work to smaller, less unionized firms. (More established firms tend to be more unionized, given that it takes time to organize a new firm.)

Table 3 Goes Here

Four patterns arise across the models. First, the competition models substantially outperform the corresponding conventional and noncompetition models, as evidenced by significant improvement in R^2. Second, the corresponding noncompetition models outperform the conventional models, indicating that accounting for sectoral shocks through the error-correction terms assists in explaining the changing pace of deunionization. Third, the error-correction terms are more significant and have larger magnitudes in the competition models, implying that shifts in competition have had large effects on the speed of U.S. deunionization. Fourth, recessions have significantly, albeit only temporarily, slowed the pace of deunionization.

Figures 4 and 5 shed further light on these results. Figure 4 illustrates the importance of controlling for the long-run impact of competition and sectoral shocks (through the error-correction terms) by plotting the fitted values from models 4 and 6 with the actual values of $\Delta\Delta$log(union share). Clearly, the competition model (model 6) more closely tracks movements in the changing pace of deunionization. To illustrate both the impact of recessions and the more accurately measured long-run relationships in the competition model, Figure 5 plots the pace of deunionization (Δlog(union share)) with the equilibrium levels implied by the error-correction terms from the competition and conventional models, with shaded areas indicating recessions. In both cases, Δlog(union share) tends to be notably higher than the equilibrium levels during recessions, while in nonrecessionary periods the equilibrium values from the competition model more closely track Δlog(union share) than do the equilibrium values from the conventional model.

Figures 4 and 5 Go Here

V. Conclusion

Conventional models of wage contracting typically focus on the role of aggregate demand and supply shocks, largely because they assume that the economy is homogeneous and that goods markets are characterized by perfect competition or a constant degree of competition. By examining endogenous wage contracting in a multisector economy, we conclude that sectoral output variance and changes in the degree of goods-market competition theoretically should affect the aggregate extent of wage contracting. Our empirical results indicate that greater sectoral output variance and greater competition both reduce the use of wage contracts.

Although direct comprehensive time-series measures of wage contracting in the United States are not available, we argue that the degree of private sector unionization is a reasonable proxy. Consistent with our theoretical model, we find that unionization is decreasing in the aggregate degree of goods-market competition. Consistent with our time-series findings, Figure 6 illustrates that the most pronounced declines in unionization since the early 1980s have occurred, on average, in sectors that faced stiffer foreign competition (e.g., manufacturing) or were deregulated in the late 1970s or early 1980s (communications, transportation, finance/insurance/real estate, and mining). While differences in factors can affect the cross-industry pattern of unionization at a given point in time, deregulation and foreign competition appear to affect changes in these relative cross-industry patterns across time.

Figure 6 Goes Here

Furthermore, unless one includes measures of both competition and sectoral output variance, misspecification may lead one to conclude empirically that inflation risk boosts unionization, contrary to contract theory in general. Our modified empirical model outperforms a conventional model, particularly in the low- to moderate-inflation period since the early 1980s and especially in the low-inflation environment of the early 1990s. Our theoretical and empirical results also confirm the conventional wisdom among economists that unionization largely depends on barriers that impede goods-market competition, whether that protection is from international or intranational competition.

Appendix A: Measuring Contract Share

Lacking comprehensive measures of the share of U.S. workers under contracts, we used data on unionization rates. Although many workers have their pay adjusted annually, we are more interested in contracts over a year in length, because such contracts expose agents to a greater risk of expectational errors resulting from underlying aggregate and sectoral shocks, which determine contract share in our theoretical model. Union contracts generally exceed a year in length, and for this reason unionization rates appear to be a reasonably good proxy for contract share.

For 1983-1996, we use the Bureau of Labor Statistics (BLS) annual series on private-sector unionization rates, including employee association members. Prior to 1983, we use Troy and Sheflin's (1985) annual series for 1956-1982, which also includes employee association members. The two series have one overlapping data point in 1983. For this year Troy and Sheflin report a unionization rate of 17.8 percent, whereas the BLS reports a rate of 18.8 percent. To adjust for this small difference, we added one percentage point to the Troy and Sheflin 1956-1982 data to create a spliced series covering the entire 1956-1996 period.

Our sample begins in 1956 for five reasons. First, by this year labor markets had adjusted to the pro-union policies of the Roosevelt and early Truman administrations, which had fostered a sizable run-up in union membership in the 1930s and 1940s. Second, the series begins after labor markets adjusted to the Taft-Hartley Act of 1949, which worked to slow the growth of unions. Third, the sample begins after unionization rates adjusted to the end of the Korean War (union share peaked in 1953 and leveled off in 1956-57). Fourth, this sample allows us to use post-Korean War profit data to construct lags of first differences of the degree of goods-market competition, which we use in the second-stage error-correction models of the change in unionization share. Fifth, our aggregate demand shock proxy, CPI inflation, was negative in levels in 1955, precluding the use of a log (interactive) specification on samples starting before 1956.

Appendix B: Measuring the Degree of Goods-Market Competition

Following Duca and VanHoose (1997 and 1998), we measure the aggregate price elasticity of demand by deriving a cyclically adjusted time series of the average markup of price over marginal cost. Our focus is in long-run movements in the markup, in contrast to studies of whether markups are cyclical [e.g., Basu and Fernald (1994), Ramey (1991), Rotemberg and Woodford (1991)]. We assume constant returns to scale, consistent with Basu and Fernald, and derive the markup using profits data.

We estimate the following specification of profits:

(B1) π = constant - β_1(FC) + β_2X + trend,

where $\pi \equiv$ profits/sales, FC $\equiv$ fixed costs/sales, and X is a vector of variables controlling for cyclical and other short-run factors. From (B1), profit share is adjusted for short-run factors ($\pi^a \equiv \pi - \beta_2 X$), which implies a cyclically adjusted markup as $\pi^a = 1 - 1/\mu^a$, where $\mu^a \equiv$ price/marginal cost. The standard model of imperfect competition implies $\mu^a - 1 = 1/(\varepsilon - 1)$ and $\varepsilon = 1/\pi^a$, where $\varepsilon \equiv |\text{price elasticity of demand}|$.

To implement (B1), we include nominal nonfinancial corporate GDP to proxy for sales, two measures of fixed costs, and several terms to control for short-run movements. In principle, fixed costs can be measured by the ratios of consumption of fixed capital and net interest to nonfinancial corporate GDP (DEP and NET, respectively).[11] However, because of shifts in the use of debt and equity and the difficulty of disentangling equity investment from stock price shifts, NET reflects swings in leverage, as well as in interest rates, spreads between corporate and Treasury rates, and in inventory financing. Because NET thereby will give a distorted picture of fixed costs, we also subtract off its estimated impact in measuring the long-run markup.[9]

To cyclically adjust profits (after IVA and CCAdj adjustments), we include the *t* through *t*-3 lags of real GDP growth (Δy), year-over-year GDP growth lagged four quarters (YOYΔy), and the t and t-1 lags of the civilian unemployment rate (U), adjusted for the 1994 survey change.[10] To control for temporary profit swings stemming from real-exchange-rate swings, the t-1 through t-4 lags of the real dollar exchange rate (RER) are included.[11] RER tracks the mid-1980s decline in profits stemming from the dollar's huge appreciation and recovery owing to the subsequent depreciation. We use levels of RER since the RER series is dominated by a mid-1980s hump that depressed profits even as the dollar fell off its 1985 highs. Of the oil variables tested, the most significant were changes (t to t-3 lags of ΔOIL and the t-4 lag of the year-over-year change in real oil prices, YΔOIL) in the real retail price of energy (CPI energy/CPI).[12] In the oil-induced recessions of the 1970s, profits fell less than in other recessions because oil profits jumped. To control for such differences, the lags and construction of oil terms parallel those of the GDP terms. In addition to cyclical variables, we adjust for profit margin restrictions during the Nixon price controls by including dummy variables for each quarter when controls were in effect.[13]

Using a 1953:Q1-96:Q4 sample, the model used to construct ε $(-1/\pi^a)$ is:

(B2) $\pi_t = 0.1643^{**} - 0.3827DEP_t^{**} - 0.7293INT_t^{**} - 0.0026U_t^{*} + 0.0021U_{t-1}^{*}$
(9.28) (-2.92) (-4.00) (-2.57) (2.24)

$- 0.0151D534TAX_t^{**} + 0.1730\Delta y_t^{**} + 0.1454\Delta y_{t-1}^{**} + 0.0914\Delta y_{t-2}^{*}$
(-7.09) (5.45) (4.23) (2.53)

$+ 0.0990\Delta y_{t-3}^{**} + 0.0504YOY\Delta y_{t-4}^{*} + 0.0252\Delta OIL_t^{*} + 0.0189\Delta OIL_{t-1}$
(2.97) (2.31) (2.32) (1.57)

$+ 0.0516\Delta OIL_{t-2}^{**} + 0.0338\Delta OIL_{t-3}^{*} + 0.0236\Delta YOYOIL_{t-4}^{*} - 0.00009RER_{t-1}$
(3.76) (2.53) (2.36) (-1.08)

$- 0.00009RER_{t-2} + 0.00003RER_{t-3}^{**} - 0.0002RER_{t-4}^{*} - 0.0018D714$
(-1.08) (0.24) (-2.66) (-0.61)

$- 0.0055D721 - 0.0104D722^{*} - 0.0106D723^{*} - 0.0097D724^{*}$
(-1.46) (-2.34) (-2.27) (-2.01)

$- 0.0065D731 - 0.0052D732 - 0.0064D733 - 0.0064D734^{*}$
(-1.36) (-1.14) (-1.59) (-2.11)

$\bar{R}^2 = 0.979$, D.W. = 2.03, Q(24) = 16.63, and $\rho(1) = 0.95$.
** (*) denotes significance at the 99% (95%) level.
t statistics are in parentheses.

(Note that in other runs, linear, quadratic, and cubic time trend terms were individually and jointly insignificant and did not affect the qualitative results.)

The model results are not surprising. The positive GDP coefficients reflect the procyclicality of profits. The negative sign on U_t and the positive sign on U_{t-1} reflect that profits are reduced by the level and change in unemployment.[14] The positive $\Delta CPIOIL_{t-1}$ coefficients mainly reflect that profits did not fall as much in oil-price-hike-induced recessions as in other recessions, while the negative effects of real exchange rates track the negative impact of the dollar's appreciation on profits in the mid-1980s.

The estimated effects of all the variables except DEP and the time trends are subtracted from π_t to yield π^a_t. The averages of quarterly data on π^a_t are used to construct annual measures of ε.

Appendix C: Measuring Sectoral Employment Variance

The variance of sectoral shocks (SSHIFT) is proxied by shifts in the share of nominal output produced by different U.S. industries, defined as the sum of squared percentage point changes in industry output shares:

$$\text{(C1)} \quad \text{SSHIFT} = \sum_{j=1}^{55}[100 \times (\text{SHARE}_{j,t} - \text{SHARE}_{j,t-1})]^2,$$

where subscript j denotes industry j, subscript t denotes the year, and $\text{SHARE}_j \equiv [(P_jY_j)/PY]$, with $P_j \equiv$ industry price and $Y_j \equiv$ industry real output, and $PY \equiv$ aggregate nominal output. Based on the Bureau of Economic Analysis's estimates of gross product originating (GPO) by 2-digit SIC-code industries, we use 55 different categories of industries. SSHIFT corresponds to calculating a weighted average of the variance of sectoral growth rates across industries. The degree of shifts in nominal output share roughly proxy the extent of shifts in expenditure share, which empirically corresponds to the relative demand shock term (δ_j) in the theoretical model. However, the measure likely reflects sectoral supply shocks as well. Industry output shares are used instead of expenditure share data, because the latter are very inconsistently aggregated across time, do not easily permit combining consumption and investment purchases by product type, and do not readily allow for precise allocations of domestic and net export spending categories to particular industries.

The gross domestic product industry data were adjusted for several breaks in data categories stemming from changes in industry classification in 1959 and 1987. First, the broader pre-1959 category of "real estate" was used. Detailed data on the division between "nonfarm housing services" and "other real estate" were not available for the pre-1959 period.

Second, separate data for "social services" and "membership organizations" were not available prior to 1959. Hence, we treated these post-1958 categories as a single category.

Third, the pre-1987 category of "miscellaneous professional services" was replaced by an "other services" category, and some "business services" output was reclassified into this new "other services category." To minimize data distortions, the pre-1987 categories of "business services" and "miscellaneous professional services" were combined to create a "business and other services" category that is defined using post-1987 data as the sum of "business services" and "other services."

Fourth, the pre-1987 categories of "banking" and "credit agencies" were replaced by two different post-1987 categories of "depository institutions" and "nondepository institutions." Because the sums of output in the two respective groups of industries were equal in 1987 while the detailed breakdowns differed greatly, these four categories were replaced by the created category of "depository and nondepository institutions," which we defined as the sum of the pre-1987 categories of "banking" and "credit agencies" before 1988 and the sum of the post-1987 categories of "depository institutions" and "nondepository institutions" after 1987.

Fifth, there were several shifts in classification that involved the renaming of two categories in 1987 and the redefinition of a category that retained its name. To prevent these classification shifts from creating a spike in SSHIFT, we combined the involved categories into a "general machinery" category, defined as the sum of nominal output in the "machinery except electrical," "electric and electronic equipment," and "instruments and related products" categories before 1987 and afterward as the sum of nominal output in the "industrial machinery and equipment," "electronic and other electrical equipment," and "instrument and related products."

Sixth, we excluded the category "holding and other investment offices." Nominal output estimates for this very small sector were negative in recent years, making these estimates highly dubious.

Finally, overlapping data for 1987 using the pre-1987 and post-1987 classification schemes revealed some slight differences in the categories of "furniture and fixtures," "stone, clay, and glass products," "paper and allied products," "rubber and miscellaneous plastics products," "telephone and telegraph," "radio and television," "legal services," and "amusement and recreation services." To prevent the relatively small differences in output levels arising from more minor reclassification shifts from creating an artificial spike in SSHIFT, the value of SSHIFT for 1987 was based on comparing 1986 output shares with 1987 output shares using the pre-1988 data basis, whereas the 1988 value of SSHIFT was based on the post-1987 classification scheme.

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FOOTNOTES

[1] Duca's (1987) framework in turn derives from Blinder and Mankiw (1984). Waller (1992) and Walsh (1995) have applied this multisector approach to examinations of political tensions that can arise in the choice of a central banker and of a central banking contract. Ramagopal (1990, 1994) has extended the Duca-VanHoose (1991) framework to explore tax-policy issues, and Ghosal and Loungani (1996) find evidence supporting its essential implications.

[2] The labor supply function is consistent with a utility function that is additively separable in consumption and leisure. While an alternative approach, such as a shopping-time framework, would yield more a more general framework, our model is includes aggregate, sectoral, supply, and demand shocks but is sufficiently tractable to yield interpretable comparative static results. In addition, by extending the frameworks of seminal papers by Gray, Ball, Woglom, and others, we are better able use our macroeconomic approach to analyze the determinants of contract usage and to develop better-performing empirical models of aggregate wage contracting.

[3] In a less stylized framework, of course, mean real wages could differ at contracting and noncontracting firms.

[4] This assumption contrasts with Wogom (1990), in which risk sharing between workers and firms is a central feature.

[5] If sectoral productivity shocks are incorporated into the model, which we have not exposited because of the exceeding complicated expressions that result, the critical value of Γ changes, but this qualitative conclusion is unaffected.

[6] The critical value of Γ relevant for this comparison need not be as large as the critical value determining the effect of greater sectoral volatility on equilibrium contract share, however.

[7] We use inflation rather than velocity shocks to measure aggregate demand variance mainly because long-run instability in money velocity and redefinitions of monetary aggregates (both M1 and M2) make it difficult to construct real-time measures that were widely observed by workers and firms.

[8] For example, if increased variances in sectoral output shares are correlated with surges in the pace of business formation, unionization rates could fall because it takes time to organize the greater number of workers employed at new firms into unions.

[9] For example, the profit-output ratio would fall if firms levered up and then would rebound if firms delever, because firms would make greater payments to debt capital holders than to equity capital holders. Not adjusting for this swing, the cyclically

adjusted profit-output ratio would be U-shaped, as it was during the late-1980s rise in leverage and the early-1990s fall in leverage.

[10] The dy_{t-i} control for short-run growth effects, and the four-quarter lag of YOYΔy controls for slightly longer lagged effects. To control for the level of resource utlization, we include the unemployment rate [downwardly adjusted by 0.2 points after 1993 since the 1994 change in the household survey added 0.2 points to the unemployment rate (BLS estimates)]. Other lags of U were insignificant.

[11] Contemporaneous RER was very insignificant. The Federal Reserve Board's real-exchange-rate measure (covering 10 nations) is used because it covers the early 1970s, in contrast to other measures. Because the series begins in 1967:Q1 and because exchange rates were fixed earlier, before 1967:Q1 we set RERE at its 1967:Q1 level. This assumption likely has little effect on the result that the real-exchange-rate time series is dominated by the mid-1980s hump.

[12] Because CPI energy data begin in 1957, prior real energy prices equal the ratio of the personal consumption expenditures (PCE) index for gasoline and oil (1987 weights) prices to the overall PCE prices. The PCE rations were break-adusted by the 1957:1 ratio of real CPI energy prices to real PCE energy prices. Since real oil price movements are dominated by changes in 1973-74, 1979-80, and 1986, this reasonable assumption has little effect on coefficient estimates, while permitting us to extend the samples by two more annual observations.

[13] The Nixon controls capped profit margins at low recessionary levels and delayed a cyclical recovery in profits (*Economic Report of the President*, 1974, p. 91 and 1973, p. 65). Separate quarterly estimates are used since a single dummy for the period 1971:4-73:4 would not reflect how different phases of the controls and their bindingness changed during this economic recovery.

[14] Denoting the coefficients on U_t and U_{t-1} as β_1 and β_2, the unemployment effects can be expressed as $(\beta_1+\beta_2)U_t - (\beta_2 U_t - \beta_2 U_{t-1}) = (\beta_1+\beta_2)U_t - \beta_2 \Delta U_t$.

Table 1: Changes in the Cross-Sector Pattern of Unionization

1983 and 1996 Unionization Shares, Inclusive of Employee-Association Members

	1983 Unionization Share	1996 Unionization Share	Percent Decline
All NonAg. Private	18.8	11.2	-40.4
Transportation & Public Utilities	46.2	28.3	-38.7
Construction	27.5	19.2	-30.2
Manufacturing	30.5	18.3	-40.0
Mining	20.7	15.0	-27.5
Wholesale & Retail Trade	8.7	6.2	-28.7
Services	7.7	6.9	-10.4
FIRE	2.9	3.1	+6.9

Decomposing the 1983-96 Decline in Unionization

	Absolute Change
Overall Unionization Share	-7.6
Due to Employment Shifts	-0.25[1]
Due to Within Sector Declines in Unionization	-7.35

1. Decline due to employment shifts equals employment-share adjusted 1983 unionization rate (11.45) minus the 1996 unionization rate (11.2). The employment-share adjusted 1983 unionization rate equals the sum of 1983 employment shares by industry multiplied by their corresponding 1996 sectoral unionization shares.

Table 2: Cointegrating Vectors For Δlog(Union Share) (without trend, 1956-96)[a]

Model:	*Conventional*	*NonCompetition*	*Competition*
log(INF) (aggregate demand σ^2)	+0.010	+0.015	-0.002
log(OILVAR) (aggregate supply σ^2)	-0.017	-0.014	-0.014
log(SSHIFT3) (sectoral output σ^2)		-0.016	-0.007
log(ϵ) (degree of competition)			-0.031
constant	-0.018	-0.108	-0.035
Max Eigen	36.044	34.788	31.726
Trace	52.525	66.643	79.477
# sign. vectors	1	1	1
Rank Sign.	5%	5%	5%

<u>Variable Definitions:</u>

INF = Annual CPI inflation. Proxy for inflation risk.
OILVAR = 3-year variance of real consumer energy prices. Proxy for aggregate supply shocks.
SSHIFT3 = 3-year variance of industry nominal output. Proxy for sectoral output shocks.
ϵ = Index of the degree of goods market competition.

a. Coefficients reported so that a positive sign on a variable X implies that indexation is increasing in X.

Table 3: Error-Correction Models of $\Delta\Delta$log(Union Share)
(1957-96, Annual Data, EC terms based on table 2 results)

	Conventional Model 1	NonCompetition Model 2	Competition Model 3	Conventional Model 4	NonCompetition Model 5	Competition Model 6
constant	-.0004 (-0.13)	0.0002 (0.06)	0.0007 (0.23)	-.0058 (-1.59)	-.0051 (-1.45)	-.0057* (-2.05)
EC_{t-1}	-.2240* (-2.14)	-.2412* (-2.25)	-.5269** (-3.86)	-.4419** (-3.20)	-.5151** (-3.63)	-.8943** (-6.42)
$\Delta OILVAR_t$				-.0034^{+} (-1.86)	-.0040* (-2.22)	-.0039** (-3.14)
$RECESS_t$				0.0175* (2.54)	0.0194** (2.86)	0.0237** (4.35)
$\bar{R}^2$	.0842	.0948	.2629	.2078	.2544	.5253
L.M.	3.4555^{+}	3.2896^{+}	2.862^{+}	0.2313	0.0774	0.2891
Q(10)	6.77	6.55	7.09	2.87	2.99	8.83
S.S.E.	0.01554	0.01536	0.01251	0.01274	0.01199	0.00763

*(**,$^{+}$) denotes significant at the 95% (99%, 90%) level, t-statistics in parentheses.

Variable Definitions:

EC = error-correction terms based on long-run determinants of Δlog(union share).
RECESS = recession year dummy.
ΔOILVAR = first difference of long-run oil variance measure.

Figure 1: The Declining Share of Unionized Private Sector Workers in the U.S.

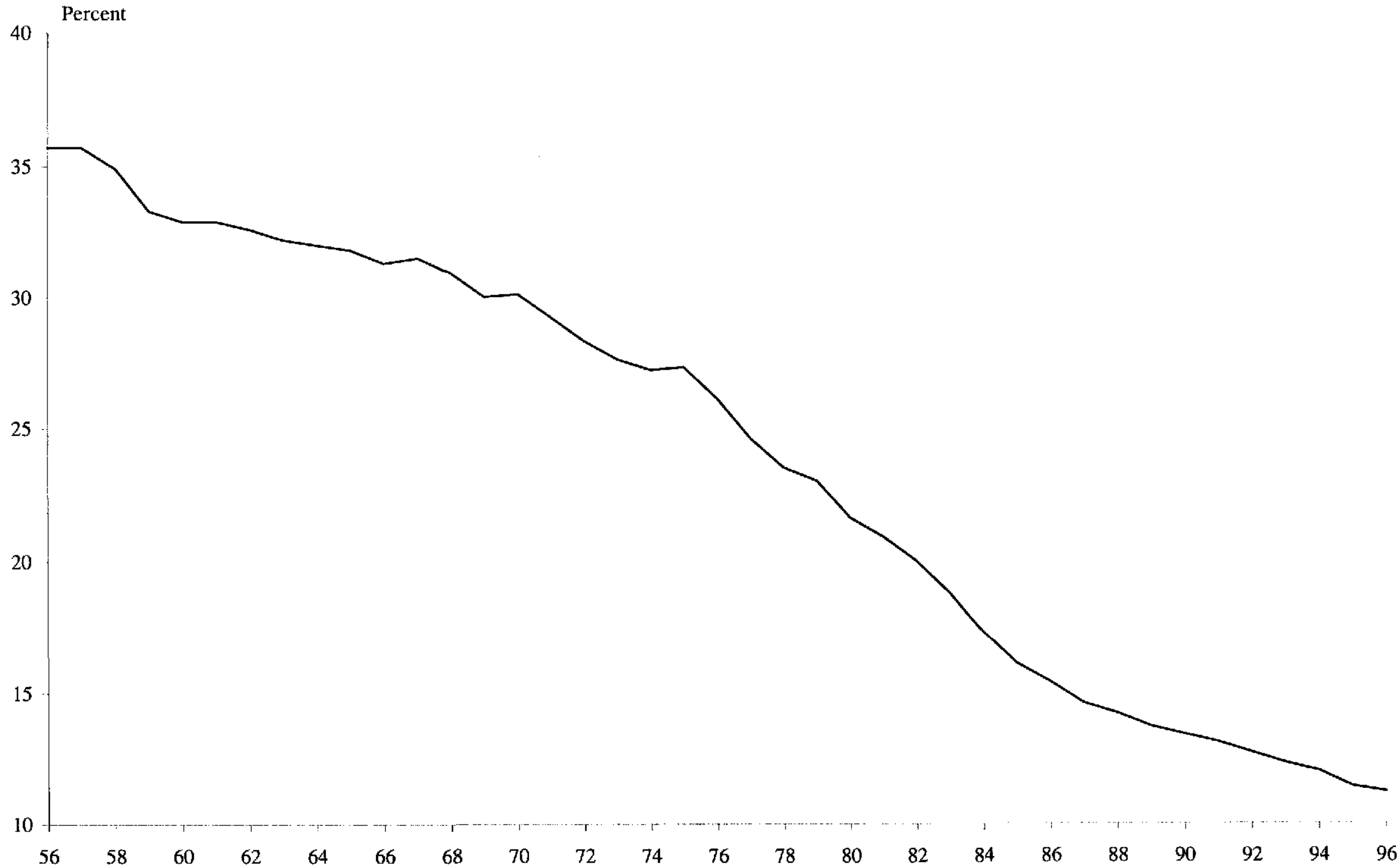


Figure 2: Movements in the Degree of Goods Market Competition (U.S. Non-Financial Corporate Sector)

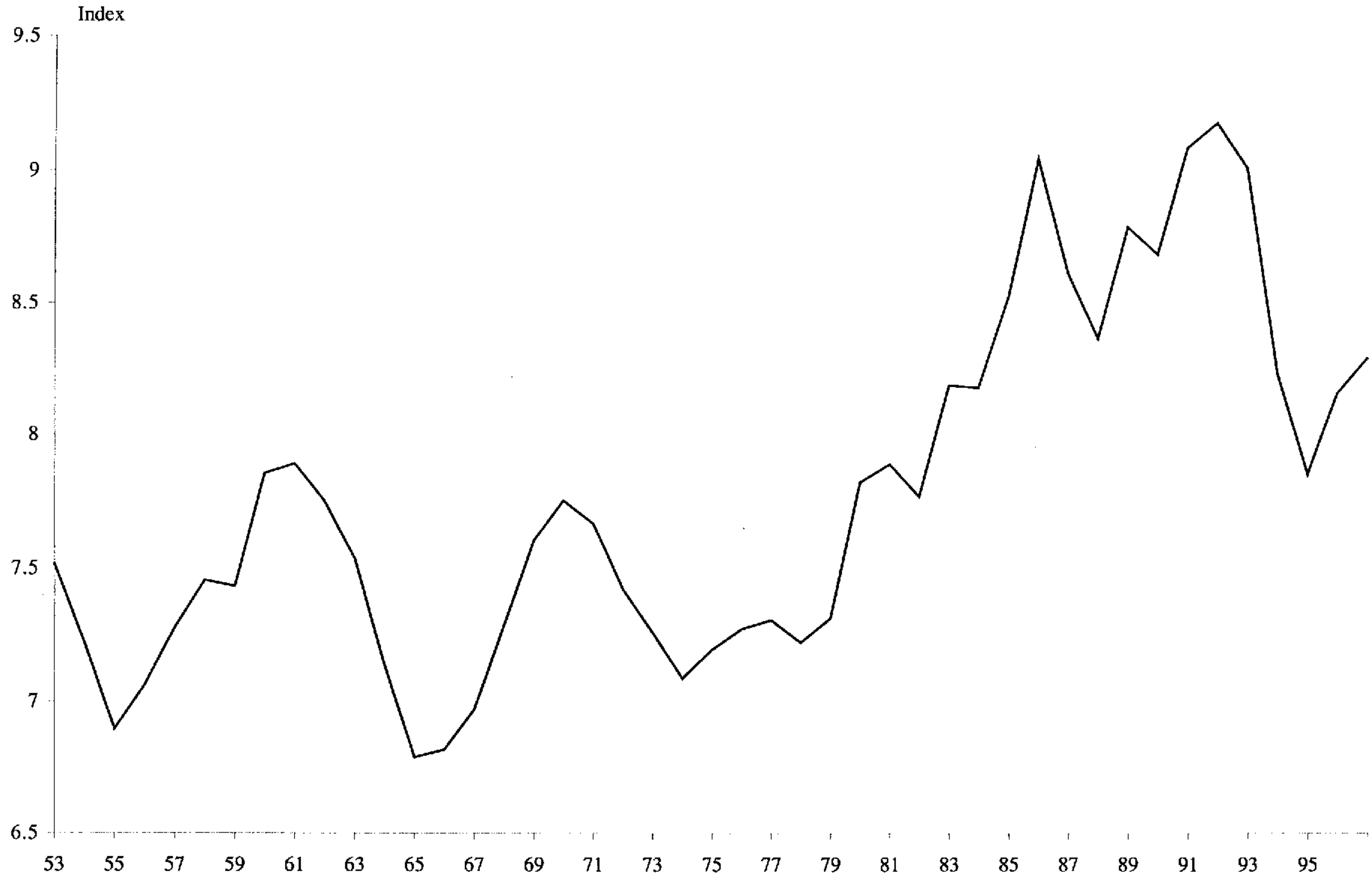


Figure 3: Sectoral Growth Variance

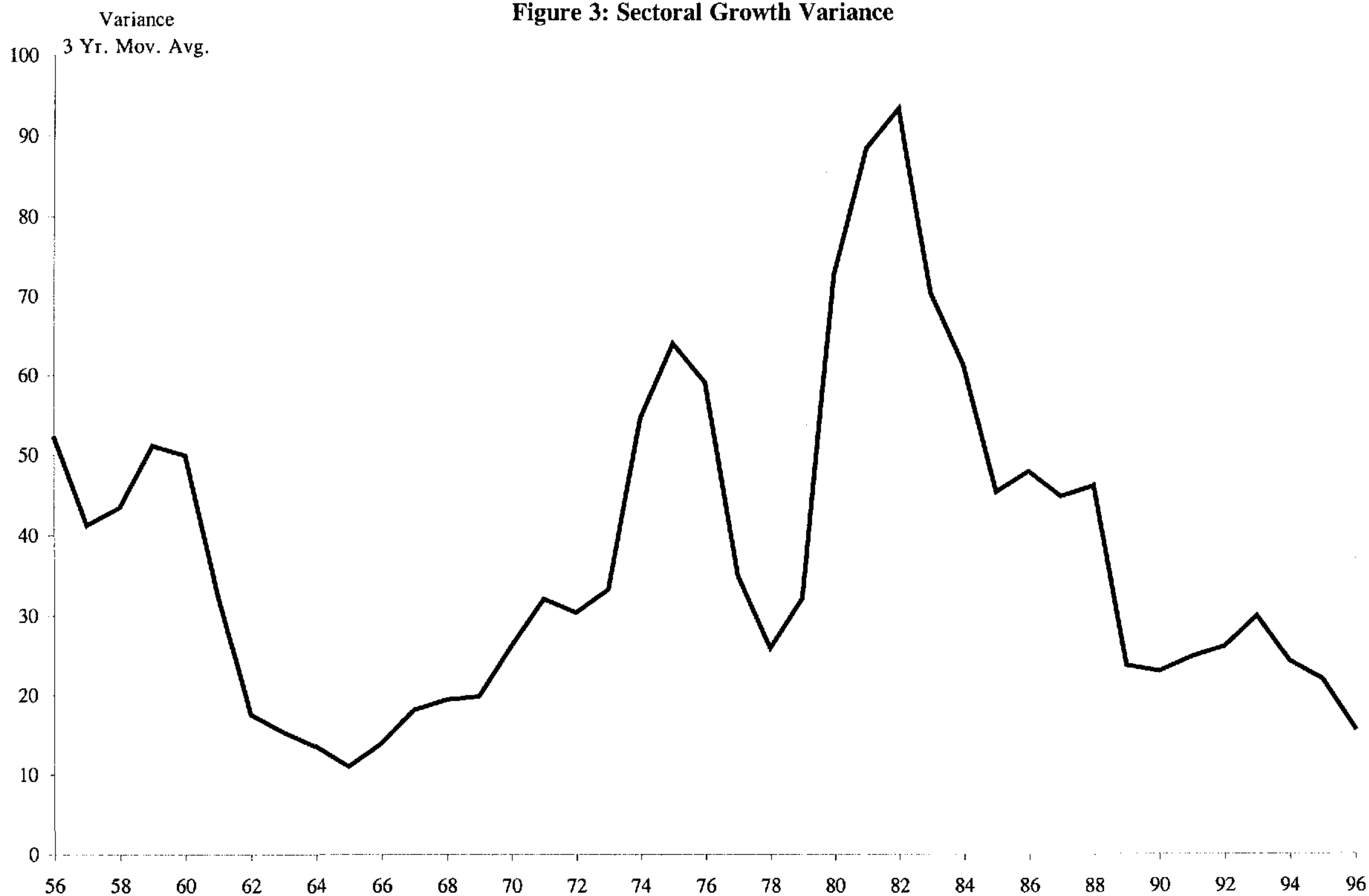


Figure 4: Actual and Fitted Change in the Growth of Union Share

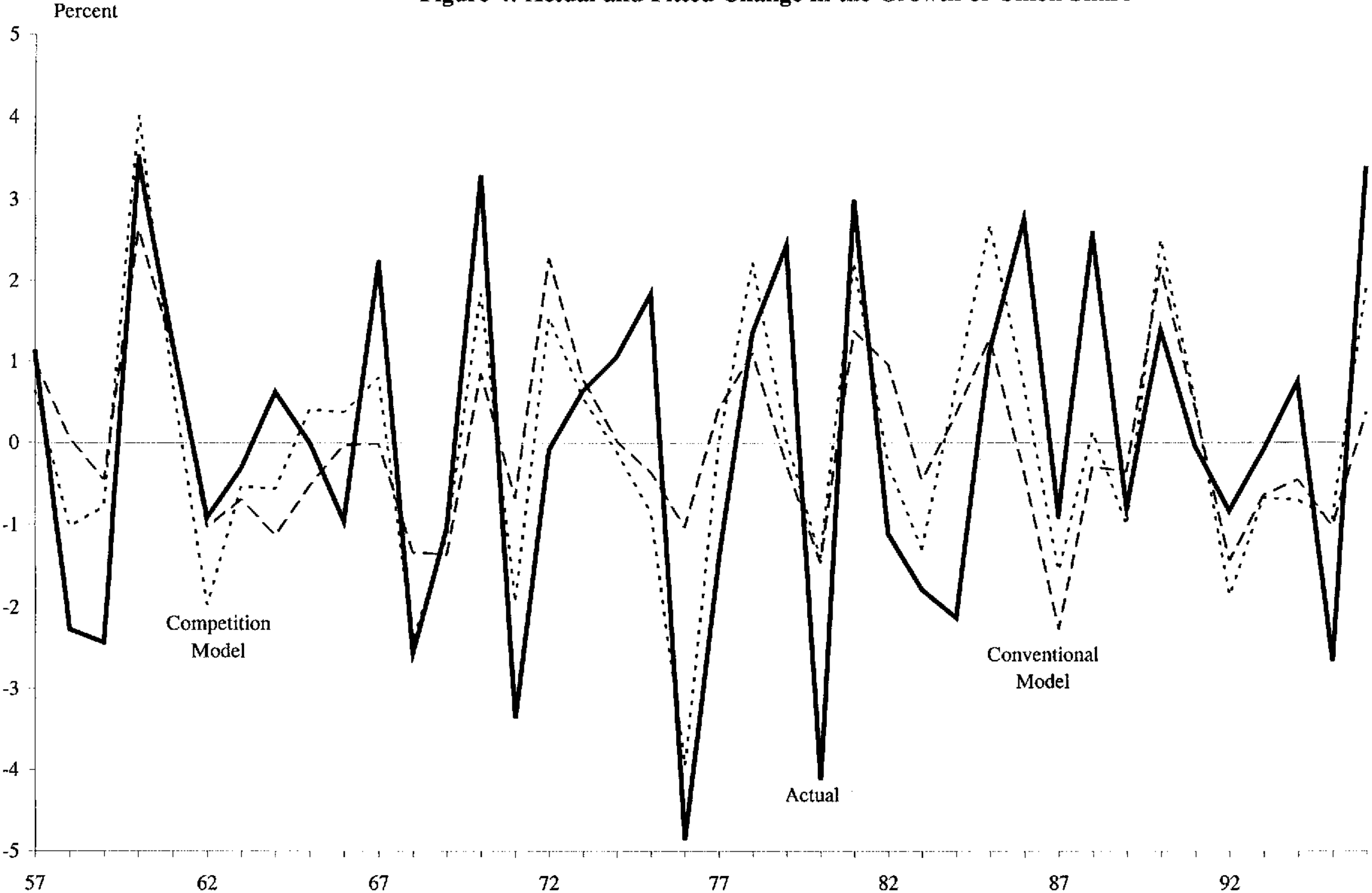


Figure 5:
Changing Competition Helps Explain Long-Run Movements in Union Share Especially After Abstracting From Recessions

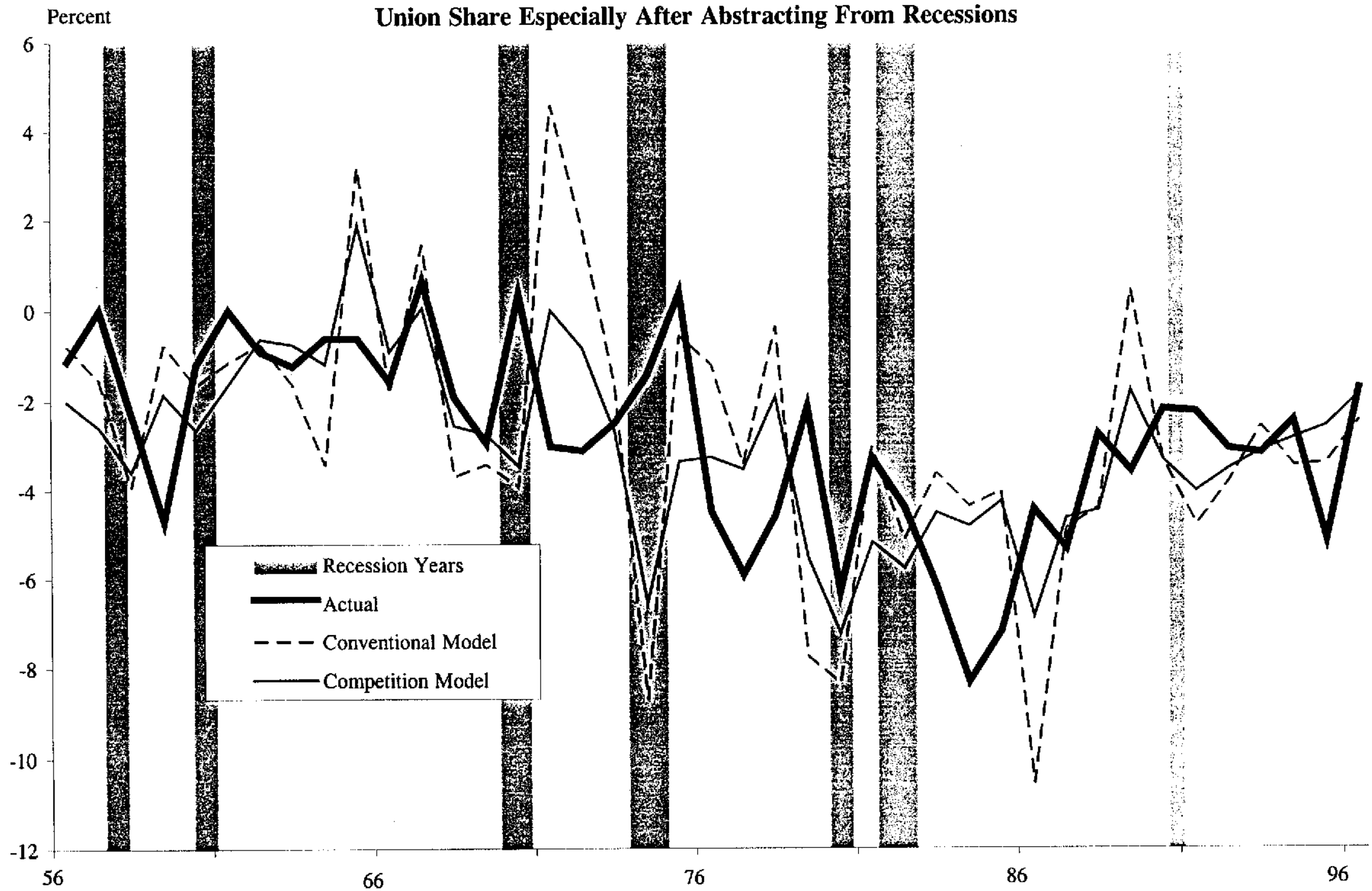
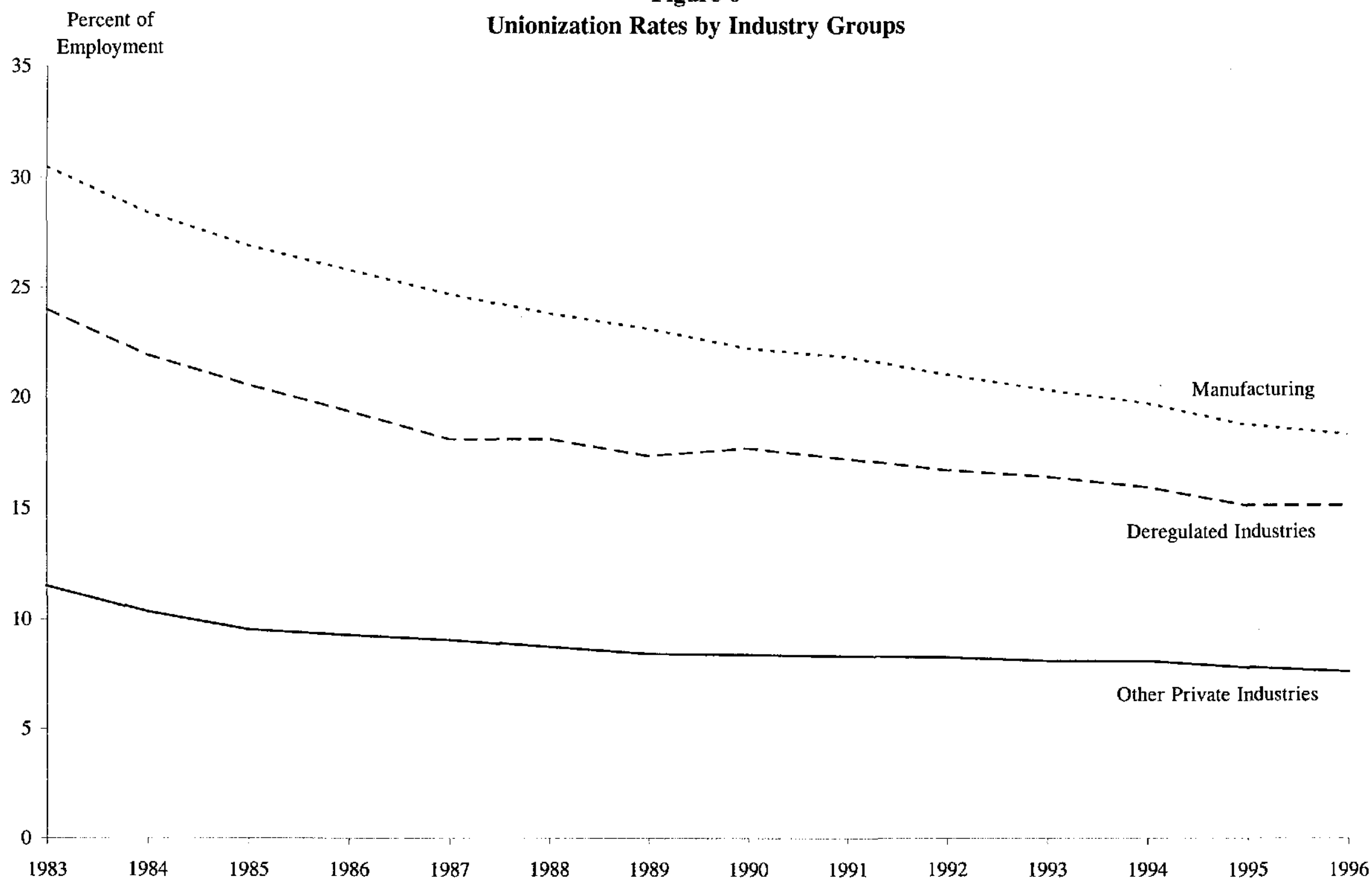


Figure 6
Unionization Rates by Industry Groups



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