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Product Markets and Paychecks: Deregulation's Effect on the Compensation Structure in Banking

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Abstract

The author investigates how deregulation designed to promote competition in the commercial banking industry affected the compensation structure for banking employees. Using establishment-based data from the Employment Cost Index (ECI) Survey and, secondarily, Current Population Survey (CPS) data, she compares changes in the distribution of wages and benefits in the banking industry to changes in unaffected industries across states and over time. In the ECI, deregulation had no impact on overall compensation inequality in banking, but this concealed off-setting changes within the compensation structure. Manager wages fell while non-manager wages held steady, reducing between-occupation compensation inequality. In contrast, between-establishment inequality increased dramatically. Comparisons with CPS data reveal that including gender and educational controls modifies the declines in between-occupation inequality. Additionally, deregulation led to shifts in the types of non-wage benefits banking employees received, and increased between-establishment inequality is due to increased heterogeneity among small banking firms following deregulation.

KEYWORDS: pay effects of deregulation, banking

PRODUCT MARKETS AND PAYCHECKS: DEREGULATION'S EFFECT ON THE COMPENSATION STRUCTURE IN BANKING

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The author investigates how deregulation designed to promote competition in the commercial banking industry affected the compensation structure for banking employees. Using establishment-based data from the Employment Cost Index (ECI) Survey and, secondarily, Current Population Survey (CPS) data, she compares changes in the distribution of wages and benefits in the banking industry to changes in unaffected industries across states and over time. In the ECI, deregulation had no impact on overall compensation inequality in banking, but this concealed off-setting changes within the compensation structure. Manager wages fell while non-manager wages held steady, reducing between-occupation compensation inequality. In contrast, between-establishment inequality increased dramatically. Comparisons with CPS data reveal that including gender and educational controls modifies the declines in between-occupation inequality. Additionally, deregulation led to shifts in the types of non-wage benefits banking employees received, and increased between-establishment inequality is due to increased heterogeneity among small banking firms following deregulation.

Markets for a range of goods and services have become more competitive in recent decades as governments dismantle industry-specific regulations and as increasing numbers of firms compete to sell their goods internationally. These forces have had predictable effects on goods prices and availability, but their effects on labor market outcomes for workers who produce for these markets are less well known. This paper exam-

ines the effects of increased competitiveness in an industry's product market on the compensation structure for the affected industry's employees. To identify effects of increased competition with changes in an industry's compensation structure, I take advantage of state-level changes to laws regulating entry into the commercial banking industry. The repeal of these laws lowered barriers to entry and reduced restrictions on scale, leading to a more competitive environment for a state's banking firms.

In contrast to its effects on good markets, effects of product market liberalization on

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This research was conducted with restricted access to the Bureau of Labor Statistics (BLS) data on-site at BLS. The views expressed here are those of the author and do not necessarily reflect those of BLS.

The author will gladly share all programs and log files upon request. Due to the confidential nature of the Bureau of Labor Statistics data, interested researchers must apply to BLS for researcher access to the data.

labor market outcomes are difficult to predict. First, these effects will generally depend on conditions of wage setting under the regulatory régime, and therefore predictions do not necessarily generalize from one industry setting to another. Additionally, liberalization may operate on worker compensation through a number of direct and indirect channels, which in turn may also differ across industries. Despite the numerous changes in worker compensation that liberalization might effect, few studies have undertaken an exhaustive review of the compensation structure changes actually associated with industry deregulation. Among those that have, the focus has generally been on unionized, capital-intensive industries, rather than growing service-sector industries like commercial banking.

This paper uses a restricted data set to go beyond earlier studies of deregulation's effect on banking industry wages. The data come from a survey of establishments conducted by the U.S. Bureau of Labor Statistics to construct their quarterly Employment Cost Index (ECI) and are used by special permission. The ECI data have two distinct advantages. First, the data include establishment identifiers, which allow me to examine effects of deregulation on dimensions of the compensation structure that are not generally observable in the standard individual-level data sets. Second, I am able to use the ECI's information on benefits expenditures to construct measures of total compensation, which allow me to compare deregulation's effects on wages to its effects on compensation more generally. Since the ECI data contain annual observations on a representative set of establishments, I can use state-level changes in banking laws to identify deregulation's effects on the banking compensation structure at the state-year-industry level.

Despite its strengths, the ECI has a major shortcoming: it includes no background information on workers. I supplement findings from the ECI with analysis using the Current Population Survey (CPS). I am thus able to compare results obtained using an important restricted-access data set with those obtained from a detailed individual-level data set. The comparison reveals a number of conflicts

between the CPS and ECI. The ECI provides a unique set of facts about the various ways in which deregulation altered the compensation structure for banking employees, but the CPS shows that several of these results are confounded by omitted variable bias.

Commercial Banking under Regulation and Deregulation

In 1920, 98% of banking firms in the United States were single office banks, or unit banks, and states used a chartering process to determine which banks operated within their borders. The required charter typically allowed a bank to serve a limited geographic area from one office.¹ Federal legislation known as the McFadden-Pepper Act enshrined unit banking as the legal default in 1927 and banned both intrastate and interstate branching unless a state passed legislation expressly permitting these corporate structures. States retained their monopoly over bank charters within their borders.²

The unit-bank or near unit-bank (extremely limited branching) market structure remained the norm in commercial banking until the late 1960s, when states began passing laws permitting various forms of bank branching. The legislative battles leading to state-level deregulation during this period rehashed the fight that led to passage of McFadden-Pepper, as more successful banks continued to press for the right to expand into other markets.³ In the intervening decades, events had also occurred to change the calculus in the bank branching debate. The savings and loan failures of the 1980s and deep recessions in some states led many

¹For a longer history of commercial banking in the United States, see White (1983) and P. Rose (1987). For details on the origins of the state charter system of banking regulation, see Sylla et al. (1987).

²The facts in this paragraph were drawn largely from Chapter 3 of White (1983) and Chapter 7 of P. Rose (1987). For details of the McFadden-Pepper legislation, see the White chapter.

³See Kroszner and Strahan (1999) and Kane (1996) for empirical studies of the political economy of banking deregulation. Economides et al. (1995) developed and tested a political economy theory of commercial banking deregulation.

to question the wisdom of a strictly atomized system of credit (Kroszner and Strahan 1999). The advent of ATMs also helped to spur a new round of debate as regulators argued over whether a freestanding ATM constituted a branch of its parent bank.

In most states, the watershed event in commercial banking deregulation was legalization of branching via mergers and acquisitions.⁴ M&A branching allowed an existing bank to create a branch by merging with or acquiring another bank and enabled banks to consolidate operations across different physical locations for the first time. Moreover, successful banking firms could finally compete directly in geographic markets previously controlled exclusively by other banking firms.

Changes to the product market environment in commercial banking following deregulation potentially affected the industry compensation structure through two broad channels. First, existing banks may have changed their behavior with the onset of deregulation. For example, if non-competitive rents disappeared following deregulation, banks may have lowered wages for employees who had previously been successful at capturing part of these rents. Alternatively, existing banks may have decided to hire a different skill mix of workers to meet the pressures of a more competitive environment.⁵

The second channel is less direct. Deregulation enabled substantial changes in the composition of the industry, which may have led independently to changes in the banking compensation structure. In particular, deregulation did away with many of the barriers to entry in the banking industry, allowing a large number of new and perhaps more specialized firms to enter. Deregulation also

allowed banks to achieve a much larger scale of operations. This led to a shift in the distribution of firm sizes throughout the industry, which was achieved at least in part through growth of existing firms. Changing either dimension of industry composition—the firm size distribution or firm heterogeneity—could lead to significant changes in industry compensation structures through well-known relationships between compensation levels and firm size and specialty.⁶

Previous authors who have explored the relationship of product market competition to wages have typically been forced by data considerations to study its effect on industry-wide wage levels.⁷ In a relevant study, Black and Strahan (2001) showed that male banking employees were adept at rent-taking under the regulatory regime, and they attributed a large share of the narrowing gender wage differential in banking to deregulation.⁸ In two of the few firm-level studies of competition's effects on banks, Jayaratne and Strahan (1998) and Flannery (1984) found evidence that more regulated banks earn excessive profits.⁹ Deregulation may also lead firms to change compensation for reasons other than rent erosion. Guadalupe (2003) examined a range of industries and found that increasing within-occupation inequality accompanies deregulation. She argued that the pressures of increased product market competition raise the cost of worker mistakes, leading firms to reward

⁴Laws permitting multibank holding companies were the first deregulatory legislation to be passed. This first wave of laws allowed holding companies to control ownership stakes in multiple banks, but the consolidation of operations across these banks was severely curtailed (Watt 1994). See Berger, Kashyap, and Scalise (1995) for a thorough overview of regulatory changes affecting the banking industry after 1979.

⁵For a theory of skill segregation across firms, see Kremer and Maskin (1996).

⁶See Brown and Medoff (1989) for a discussion of the robustness of firm size wage premiums. Berger et al. (2005) found that lending practices differ across small and large banks, suggesting product specialization across firm sizes.

⁷Examples include Card (1996), N. Rose (1987), and Black and Strahan (2001), who also studied wage changes in banking following deregulation. See Winston (1993) or Peoples (1998) for an overview of these and other studies. See OECD (2002) for a cross-country study of the relationship between product market regulations and labor market outcomes in European countries.

⁸Black and Brainerd (2004) also found evidence of gender-biased rent allocation across a number of regulated industries.

⁹In keeping with the theory of inefficiency in regulated industries, Jayaratne and Strahan (1996) also found that deregulation improved efficiency in the banking industry.

competence and performance more highly and increasing inequality among workers in a given occupation. Hubbard and Palia (1995) found evidence that this phenomenon occurs in banking. They reported that the pay of CEOs in more competitive banking markets is, on average, both higher and more sensitive to performance than that of their counterparts working in less competitive markets.

While little is known about ensuing labor market effects, the shifts in industry composition along size and product diversity dimensions have been explored by finance economists. Expansion of efficient firms following deregulation led to a well-known increase in measures of industry concentration at the national level. At the local level, however, the industry's tendency toward increased concentration was far from uniform. Bergstresser (2001) showed that half of the MSAs affected by branching reform experienced declining levels of concentration among their commercial banking firms.¹⁰ Rhoades (2000) noted that the number of banking firms in the United States fell by approximately 40% between 1984 and 1998, but the number of bank establishments (offices) continued to grow at a stable rate over the same period. Indeed, the growth in the number of bank branches appears to have prevented any real decline in employment for bank tellers, for whom ATMs became a substitute during this period. The efficiency effects of this increase in scale among a subset of firms have received comparatively little study. One reason may be that scale effects in commercial banking are not large, as some authors have suggested (Nakamura 1993), and thus repealing restrictions on scale had little effect on the industry.

Empirical Strategy and the Employment Cost Index Data

The empirical approach in this paper is straightforward. I use the timing of state-level deregulation legislation to identify exogenous changes to the product market environment for a state's commercial banks. Following other authors, I use the year that a state passed legislation permitting branching via M&A to mark the beginning of the deregulated period.¹¹ Table 1 lists the 50 states in the order in which they adopted key legislation deregulating their commercial banking markets. Due to concerns about the ECI's representativeness prior to 1987 (discussed in the Data Appendix), the treatment group consists only of states that deregulated after 1987.

In the specifications that follow, I regress measures of compensation and the compensation structure on an indicator for a state banking industry's deregulation status to arrive at reduced-form estimates of deregulation's effect on the compensation measures. Because the relevant legislative changes targeted a single industry, this identification strategy produces triple difference estimates of deregulation's effect on the banking compensation structure. Measures of the banking compensation structure are compared with those in other industries in the same state and year. The differences are then adjusted for state, year, and industry effects.¹² The innovation in this paper lies not in its empirical methodology but in the setting in which these methods are applied. By using a data set in which it is possible to match wage and benefits information across employment establishments, I am able to construct a number of compensation and compensation

¹⁰Conflicting trends in industry concentration at the local level were the result of a more competitive market for banking services. Efficient firms expanded relative to less efficient firms, increasing concentration in their local markets. At the same time, new firms entered to compete against incumbents. In areas where incumbent banking firms were particularly inefficient, the market was divided among newcomers and a stagnating incumbent, leading to decreased concentration.

¹¹Once a state passed laws allowing M&A branching, laws permitting branching via other channels generally followed soon after. However, the short intervals between M&A branching legislation and subsequent deregulatory legislation make it difficult to identify separate effects for each level of deregulation.

¹²Despite some correlation of banking deregulation with state macroeconomic trends, I maintain, as have other authors, that the exact timing of the legislation was uncertain and that no evidence exists to suggest that workers anticipated the legislation's effects (Black and Strahan 2001; Jayaratne and Strahan 1998). Moreover,

Table 1. Years of Deregulation: States and the Years they Permitted M&A Branching.

| Early Deregulators: M&A Branching Permitted Prior to 1970 (12 states) | |
|---|-----------|
| Alaska | |
| Arizona | |
| California | |
| D.C. | |
| Delaware | no years |
| Idaho | available |
| Maryland | for this |
| Nevada | group |
| North Carolina | |
| Rhode Island | |
| South Carolina | |
| South Dakota | |
| M&A Branching Permitted Starting 1970–1979 (6 states) | |
| Vermont | 1970 |
| Maine | 1975 |
| New York | 1976 |
| New Jersey | 1977 |
| Virginia | 1978 |
| Ohio | 1979 |
| M&A Branching Permitted Starting 1980–1986 (12 states) | |
| Connecticut | 1980 |
| Alabama | 1981 |
| Utah | 1981 |
| Pennsylvania | 1982 |
| Georgia | 1983 |
| Massachusetts | 1984 |
| Nebraska | 1985 |
| Oregon | 1985 |
| Tennessee | 1985 |
| Washington | 1985 |
| Hawaii | 1986 |
| Mississippi | 1986 |
| M&A Branching Permitted Starting in 1987 (5 states) | |
| Kansas | 1987 |
| Michigan | 1987 |
| New Hampshire | 1987 |
| North Dakota | 1987 |
| West Virginia | 1987 |
| Treatment States: M&A Branching Permitted Starting 1988 or Later (16 states) | |
| Florida | 1988 |
| Illinois | 1988 |
| Louisiana | 1988 |
| Oklahoma | 1988 |
| Texas | 1988 |
| Wyoming | 1988 |
| Indiana | 1989 |
| Kentucky | 1990 |
| Missouri | 1990 |
| Montana | 1990 |
| Wisconsin | 1990 |
| Colorado | 1991 |
| New Mexico | 1991 |
| Minnesota | 1993 |
| Arkansas | 1994 |
| Iowa | [1994] |

Notes: Data are based on Black and Strahan (2001), Amel (1993), and Kroszner and Strahan (1999). Iowa had not yet deregulated in 1994, the year in which IBBEA passed allowing interstate branching and superseding state branching laws. For the analysis, Iowa is considered to have deregulated in 1994, but results are not sensitive to this.

Treatment states are those in which the deregulatory change is observed in the BLS data.

structure measures that cannot be observed with more traditional micro data.

The data are from the U.S. Bureau of Labor Statistics Employment Cost Index (ECI) micro data, collected by BLS to track wages, benefits coverage, and benefits costs.¹³ The ECI data are collected using a survey of establishments randomly drawn from within strata defined by establishment size, industry, and location. In keeping with the ECI's purpose as an index of employer costs, its data contain no individual worker-level or demographic information. Instead, the observational unit in the ECI is the "job." One should think of a job as an occupation–union-status–full-time status cell. The establishment provides data averaged over all employees in sampled jobs. Data are collected from each establishment for between two and eight jobs, depending on the establishment's size. Banking establishments typically provide data on four to six jobs.

The analysis focuses on changes in wages and benefits coverage among banking employees following deregulation. Due to the nature of the ECI data, these are observed only at the job level. Because wages are constructed at the job level and equal a job's average wage within the establishment, they likely understate within-job wage variance. Job-level benefits data reflect per employee expenditures by the firm.¹⁴ Following Pierce (2001), I construct a total hourly compensation figure at the job level by summing the wage and hourly benefits expenditures reported for each job.

Means for several ECI variables appear in Table 2. Throughout the paper, data are weighted using BLS-constructed weights to

the inclusion of several levels of industry, state, and time controls should alleviate concerns that my results stem from slow-moving trends in the wage structure for affected employees.

¹³For detailed discussions of the ECI data, see the Data Appendix to this paper, Pierce (2001), U.S. Department of Labor (1997), or the most current version of the National Compensation Survey Documentation (NCS) at <http://www.bls.gov/ncs/ect/home.htm>. The NCS absorbed the ECI.

¹⁴See Pierce (2001) for a discussion of how employer benefits costs relate to employee evaluation of benefits.

Table 2. Variable Means in the ECI Sample.

| | Banking | | | Overall ECI Sample | | |
|-------------------------------|----------------|----------------|----------------|--------------------|----------------|----------------|
| | 1987 | 2001 | Period Avg. | 1987 | 2001 | Period Avg. |
| Hourly Wage | 8.75 (6.00) | 8.49 (5.07) | 8.33 (5.30) | 8.45 (5.06) | 8.48 (5.56) | 8.23 (5.33) |
| % Non-Managerial | 0.72 (0.45) | 0.72 (0.45) | 0.74 (0.44) | 0.78 (0.41) | 0.78 (0.41) | 0.79 (0.41) |
| % Union | 0.01 (0.08) | 0.00 (0.00) | 0.01 (0.11) | 0.17 (0.38) | 0.10 (0.30) | 0.14 (0.34) |
| % Full-Time | 0.91 (0.28) | 0.83 (0.37) | 0.88 (0.33) | 0.84 (0.37) | 0.78 (0.42) | 0.79 (0.41) |
| % Receiving End-of-Year Bonus | 0.39 (0.49) | 0.56 (0.50) | 0.43 (0.50) | 0.33 (0.47) | 0.44 (0.50) | 0.37 (0.48) |
| % Eligible for Severance | 0.19 (0.39) | 0.29 (0.45) | 0.23 (0.42) | 0.10 (0.30) | 0.12 (0.32) | 0.10 (0.30) |
| Establishments/Year | 99 | 248 | 140 | 3,679 | 6,940 | 5,038 |
| Median Establishment Size | 468 | 179 | 341 | 198 | 198 | 219 |
| Observations | 692 | 1,121 | 10,378 | 18,822 | 29,147 | 332,989 |

Notes: Data are from the U.S. Bureau of Labor Statistics Employment Cost Index Survey. Standard errors are in parentheses. Data are weighted using adjusted BLS weights as described in the Data Appendix. Hourly wages are real; deflator is 1982–1984 average CPI.

ensure representativeness of the original stratified sample.¹⁵ The final set of columns gives means for the ECI as a whole, and the preceding columns give means for the commercial banking industry alone.¹⁶ Several comparisons within Table 2 are relevant to questions about the generalizability of the results in this study. Wage levels in commercial banking are close to the economy average over the period. The entire series of mean wages is shown in Figure 1. Wages in banking declined slightly over the period while wages in the economy as a whole were flat. In results not shown here, wage variance in banking was lower than in the economy as a

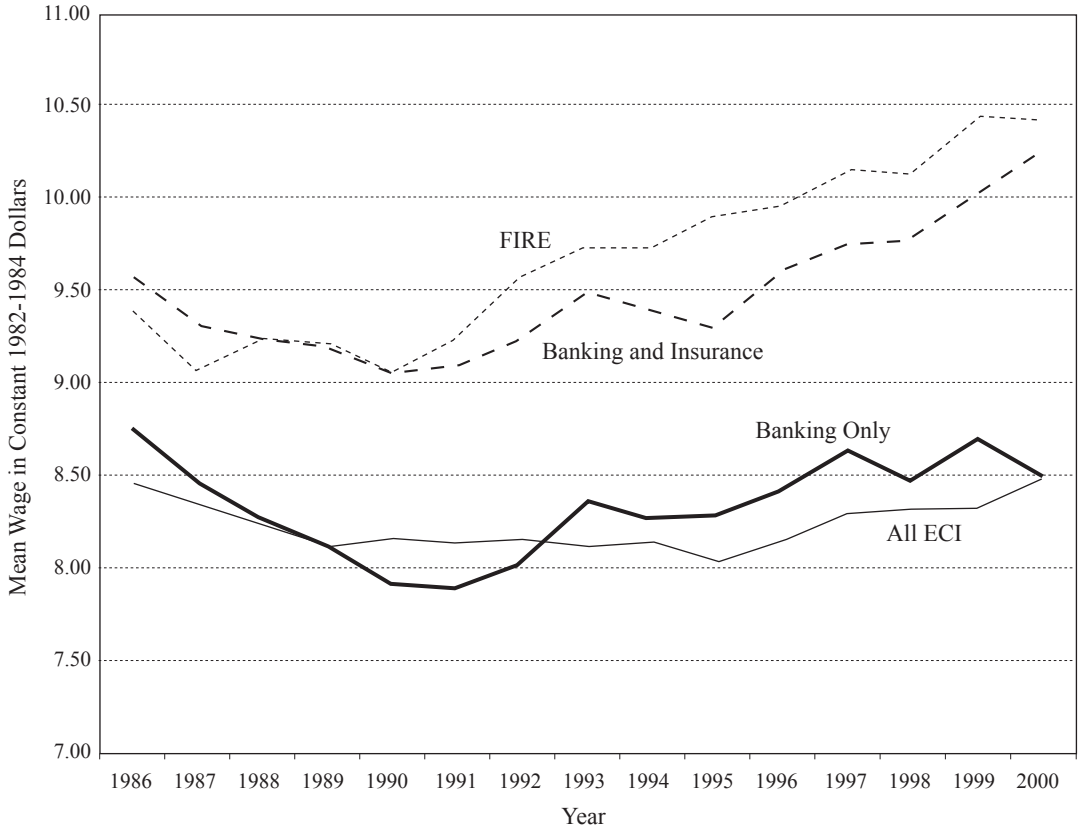
whole during the period covered by the data. In terms of the share of non-managerial jobs and the share of union jobs, banking differs from the ECI average, but it is similar to the service jobs in FIRE (finance, insurance, and real estate) on these measures. Banking is not unionized and seems to have a slightly higher fraction of managerial jobs than the economy as a whole. The fraction of full-time jobs appears high relative to similar figures from other sources, such as the CPS, but this is consistent with other estimates from the ECI. Banking employees also tend to receive two key forms of non-wage compensation, severance and bonus packages, more often than the average ECI employee.

A final trend to notice in Table 2 is the sharp decline in the size of the median banking establishment. Because the sampling procedure for the ECI is biased toward larger establishments, the levels may appear high relative to other sources, and some of the decline in median size is probably due to an increase in the sample size in 2000, which would favor smaller establishments. Nevertheless, the size of the median banking establishment fell by almost 300 employees, from 468 in 1987

¹⁵Weights are adjusted to correct for attrition and make the rotating panel representative of the national economy each year. For details on this procedure, see the Data Appendix. Note that the number of state-year-industry cells is less than what might be expected given 50 states, 14 years of data, and two-digit industries. The discrepancy stems in part from the fact that jobs in government and agriculture are excluded. Moreover, the ECI's procedure does not necessarily result in sampling from all industries in all states. It is intended to be representative of national employment in an industry, not state-level employment.

¹⁶Throughout this paper, banking industry observations include only commercial banks, savings institutions, and credit unions.

Figure 1. Mean Wages in the ECI.



to 179 in 2001. This is likely due in part to the effects of deregulation over the period, which allowed banks to create smaller partial-service branches.¹⁷

The major shortcoming of the ECI data is their lack of demographic information on the workers included in the sampled jobs. I compensate somewhat for this shortcoming by using the March extracts from the Current Population Survey to replicate and extend the ECI analysis in instances where a comparison with estimates from a more rep-

resentative sample with known background characteristics would be useful. The CPS sample spans 1977 to 1999 and includes all workers who were employed during the year prior to the survey.

Results: Deregulation's Effects on the Banking Compensation Structure

I first examine the effects of banking deregulation on broad measures of compensation inequality, since this is an area in which the ECI data have much to offer. I then turn to the question of how the changes in inequality that I identify came about. It is in this context that I examine deregulation's effect on wage levels in banking. The final results in this section exploit the ECI's second strength and consider deregulation's effects

¹⁷ Establishment size numbers are unweighted. The number of unique banking establishments increases over the life of the sample, from 112 in 1987 to 265 in 2001. Due to the ECI's sampling procedure, these contain many of the largest banks in the United States, which accounts for the high average establishment size.

on the types of non-wage benefits banking employees receive.

Broad Changes in Compensation Inequality

I decompose variance in the banking compensation structure into between-establishment, between-occupation, and residual components. Following Groshen (1991), I use two variants of dummy variable regressions to decompose compensation variance into contributions from establishment and occupation.¹⁸ The variants both include main effects for establishment and occupation, but only one includes an interaction between establishment and occupation. Both methods produce establishment and occupation variance measures for state-year-industry cells as well as measures of residual variance. However, measures from the decomposition that includes an establishment-occupation interaction (hereafter the interaction decomposition) include separate estimates of the variance explained by the interaction. Measures from the main-effects-only decomposition (hereafter the main effects decomposition) assign this variance, due to covariance between establishment and occupation, to its respective measures of establishment and occupation variance, since these are estimated from separate decompositions.

Table 3 provides summary statistics for measures from the main effects and interaction decompositions of wage and total compensation variance. The relative contributions of establishment and occupation differ little across the two types of compensation. Also, the contributions of establishment and occupation are quite similar across the two decomposition methods. This is due to the relatively small contribution of the establishment-occupation interaction to overall variance in both wages and total compensation. In all four decompositions, a large share of the variance remains unexplained.

The results of interest are from secondary regressions like the following:

$$(1) \quad y_{kst} = \beta_0 + \beta_1 \text{BANKING}_k^* \text{DEREG}_{st} + \beta_2 \text{DEREG}_{st} + \theta_k + \theta_s + \theta_t + \varepsilon_{kst}$$

¹⁸For more details on this procedure, see Groshen (1991).

Table 3. Means of Constructed Variance Components.

| Measure | Decomposition of Wages | Decomposition of Total Compensation |
|-----------------------------------|------------------------|-------------------------------------|
| Interaction Decomposition | | |
| Total Variance | 0.155 (0.002) | 0.145 (0.002) |
| Between-Occupation | 0.044 (0.001) | 0.037 (0.001) |
| Between-Establishment | 0.035 (0.001) | 0.039 (0.001) |
| Between-Estab. & Occ. Cell | -0.017 (0.001) | -0.014 (0.001) |
| Residual | 0.079 (0.001) | 0.076 (0.001) |
| Main Effects Decomposition | | |
| Total Variance | 0.155 (0.002) | 0.145 (0.002) |
| Between-Occupation | 0.059 (0.001) | 0.051 (0.001) |
| Residual, Occupation Reg. | 0.097 (0.001) | 0.093 (0.001) |
| Between-Establishment | 0.049 (0.001) | 0.052 (0.001) |
| Residual, Establishment Reg. | 0.106 (0.001) | 0.101 (0.001) |

Notes: Data are from the U.S. Bureau of Labor Statistics Employment Cost Index Survey, 1987-2001. The table shows means and standard errors (in parentheses) of variance components constructed as described in the text. The wage decomposition and total decomposition samples each have approximately 19,000 observations. Details of the decomposition are given in Groshen (1991).

where y is one of the measures of compensation variance, k indexes two-digit industries, s indexes state, and t indexes year. θ_k , θ_s , and θ_t are dummy variables for industry, state, and year. DEREK is equal to one in years following a state's deregulation of its banking industry.

The banking-deregulation coefficient from the second stage regressions should be interpreted as a triple difference. It is deregulation's effect on the banking industry wage structure relative to non-affected industries after controlling for state and year effects.¹⁹ As mentioned earlier,

¹⁹Allowing the effect of deregulation to vary according to time since deregulation produced results that were too imprecise to be useful.

Table 4. Deregulation's Effect on Banking Compensation Structure (Interaction Decomposition).

| | <i>Dependent Variable</i> | | | | |
|---|---------------------------------------|--|---|--------------------------------------|------------------------------|
| | <i>Total Variance of Log Wage</i> | <i>Between Occupation Var.</i> | <i>Between Establishment Var.</i> | <i>Between Job-Cell Var.</i> | <i>Residual Variance</i> |
| Wage Decomposition Estimates | | | | | |
| Banking*Dereg. Dummy | -0.0231 [0.0175] | -0.0374 [0.0106]** | 0.0158 [0.0079]** | 0.0034 [0.0100] | -0.0127 [0.0134] |
| Deregulation Dummy | -0.0115 [0.0039]** | 0.0029 [0.0024] | -0.0077 [0.0018]** | -0.0047 [0.0023]** | 0.0017 [0.0030] |
| Total Compensation Decomposition Estimates | | | | | |
| Banking*Dereg. Dummy | -0.0274 [0.0162]* | -0.0332 [0.0093]** | 0.0134 [0.0082] | 0.0100 [0.0088] | -0.0232 [0.0121]* |
| Deregulation Dummy | -0.0162 [0.0037]** | -0.0019 [0.0021] | -0.0078 [0.0018]** | -0.0010 [0.0020] | -0.0031 [0.0027] |

Notes: Data are from the U.S. Bureau of Labor Statistics Employment Cost Index Survey, 1987–2001. Results shown are from the second regression in a two-part estimation; each column is a separate OLS regression. The first stage regresses individual wages on the establishment-occupation interaction and on the main effects of establishment and occupation separately and together. The second stage regresses R-squared differences from the first stage regression on state, year, and two-digit industry dummies. Details of decomposition are given in Groshen (1991). Wage decomposition and total compensation samples each have approximately 19,000 observations.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

a state's decision to deregulate its banking industry may depend on prevailing economic conditions in the state. To account for this, I verified that all results are robust with respect to the inclusion of state-year fixed effects, which control for any non-linear, state-level compensation structure trends over the period; only results from the specification in (1) are reported.

The tables show results using all ECI industries as the comparison group of non-affected industries. Unreported results using alternative comparison industries, such as FIRE, were not substantively different. Also note that the decompositions only included two occupation groups, managers and non-managers, so between-occupation results should be interpreted as reflecting changes²⁰ between these two groups.

²⁰I considered only two occupation groups because the ECI only categorizes occupations at the one-digit level. At this level of classification, one managerial category and one non-managerial category comprised virtually all commercial banking jobs, so I grouped the few remaining jobs into either the manager or non-manager categories.

Results from estimation of equation (1) are summarized in Tables 4 and 5, which present estimates using measures from the interaction and main effects decompositions, respectively. Each column in the tables represents a separate regression of the dependent variable in the column heading on dummy variables for deregulation, state, year, and industry, as well as the interaction of banking and deregulation. The first column in the two tables is the same: it estimates the effect of deregulation on total variance in banking industry compensation. In addition, each table shows results using measures from decomposition of total compensation (wages plus benefits) and from decomposition of wages alone.

The first result to notice in Tables 4 and 5 supports a null hypothesis: deregulation had no effect on wage inequality among banking industry employees and little effect on total compensation inequality in the ECI data. However, this conceals two pronounced but conflicting trends within the compensation structure. First, deregulation led to a significant reduction in between-occupation inequality. The estimated effects of deregula-

Table 5. Deregulation’s Effect on Banking Compensation Structure (Main Effects Decomposition).

| | <i>Dependent Variable</i> | | | | |
|---|-----------------------------------|--------------------------------|---------------------------------------|-----------------------------------|--|
| | <i>Total Variance of Log Wage</i> | <i>Between Occupation Var.</i> | <i>Residual Var., Occupation Reg.</i> | <i>Between Establishment Var.</i> | <i>Residual Var., Establishment Reg.</i> |
| Wage Decomposition Estimates | | | | | |
| Banking*Dereg. Dummy | -0.0231 [0.0175] | -0.0296 [0.0127]** | 0.0065 [0.0122] | 0.0236 [0.0107]** | -0.0467 [0.0141]** |
| Deregulation Dummy | -0.0115 [0.0039]** | -0.0009 [0.0029] | -0.0107 [0.0027]** | -0.0115 [0.0024]** | -0.0001 [0.0032] |
| Total Compensation Decomposition Estimates | | | | | |
| Banking*Dereg. Dummy | -0.0274 [0.0162]* | -0.0291 [0.0114]** | 0.0040 [0.0118] | 0.0175 [0.0107] | -0.0474 [0.0128]** |
| Deregulation Dummy | -0.0162 [0.0037]** | -0.0053 [0.0026] | -0.0112 [0.0027]** | -0.0112 [0.0024]** | -0.0060 [0.0029]** |

Notes: Data are from the U.S. Bureau of Labor Statistics Employment Cost Index Survey, 1987–2001. Results shown are from the second regression in a two-part estimation; each column is a separate OLS regression. The first stage regresses individual wages on the main effects of establishment and occupation separately and together. The second stage regresses R-squared differences from the first-stage regression on state, year, and two-digit industry dummies. Details of the decomposition are given in Groshen (1991). Wage decomposition and total compensation samples each approximately 19,000 observations.

*Statistically significant at the .10 level; *at the .05 level; **at the .01 level.

tion on this measure in the second columns of the two tables are large and statistically significant. The coefficient from the decomposition of wages in Table 4 (using the interaction decomposition) represents a decline of 25% from the mean of pre-deregulation between-occupation variance.²¹ The Table 5 estimate (using the main effects decomposition) is somewhat smaller but still implies a reduction of between-occupation variance of 19% from the mean. Estimates using the decomposition of total compensation are very similar.

A second, opposing effect of deregulation on the compensation structure is apparent: deregulation increased between-establishment compensation variance. This, combined with the changes in between-occupation variance, resulted in stable total variance following deregulation. The effect of deregulation on between-establishment wage variance in banking is 0.0158 using

the interaction decomposition and 0.0236 using the main effects decomposition; both are statistically significant. The estimates are similar in the decompositions of total compensation but not quite statistically significant, with p-values of 0.102 and 0.101 using the interaction and main effects decompositions, respectively. These results imply large increases of 30–50% from the mean in the level of between-establishment inequality in the banking industry.²²

Between-Occupation Changes

The broad compensation structure changes described in the previous section could

²¹Author’s calculations. Separate pre- and post-deregulation variance decompositions are not reported.

²²Note that in this setting, the dependent variables are measures of dispersion in predicted, rather than actual, values. In cases where the sample sizes used to form predicted values are very small, as is the case with the establishment dummies, variance in the predicted values may be large due to sampling error. This problem is not likely to be more severe for banks than for other industries, so the triple difference will largely difference out this noise. (I am grateful to a referee for pointing this out.)

Table 6. Effect of Banking Deregulation on Compensation by Occupation and Industry.

| Independent Variable | Log Wages | | | Log Total Compensation | | |
|---|--------------------|---------------------|---------------------|------------------------|---------------------|---------------------|
| | [1] | [2] | [3] | [4] | [5] | [6] |
| Banking*Dereg. Interaction | 0.015 [0.039] | -0.167 [0.064]** | -0.182 [0.064]** | -0.007 [0.047] | -0.203 [0.073]** | -0.222 [0.073]** |
| Banking*Dereg*Non-Manager | | 0.246 [0.061]** | 0.261 [0.061]** | | 0.261 [0.069]** | 0.281 [0.068]** |
| Banking*Non-Manager | | 0.108 [0.188] | 0.123 [0.189] | | 0.154 [0.207] | 0.165 [0.209] |
| Dereg*Non-Manager | | -0.068 [0.029]** | -0.079 [0.026]** | | -0.058 [0.030]* | -0.071 [0.026]** |
| Non-Manager Dummy | -0.683 [0.20]** | -0.849 [0.034]** | -0.607 [0.023]** | -0.067 [0.021]** | -0.865 [0.034]** | -0.604 [0.022]** |
| Banking Industry Dummy | 0.139 [0.091] | 0.574 [0.083]** | 0.548 [0.084]** | 0.148 [0.095] | 0.649 [0.091]** | 0.623 [0.093]** |
| Deregulation Dummy | N/A | N/A | N/A | N/A | N/A | N/A |
| Union Dummy | 0.281 [0.027]** | 0.002 [0.032] | 0.281 [0.027]** | 0.358 [0.033]** | 0.018 [0.029] | 0.358 [0.033]** |
| Full-Time Dummy | 0.418 [0.011]** | 0.199 [0.028]** | 0.418 [0.011]** | 0.539 [0.012]** | 0.303 [0.027]** | 0.539 [0.012]** |
| Non-Managerial*Union | | 0.296 [0.036]** | — | — | 0.361 [0.043]** | — |
| Non-Managerial*Full-Time | | 0.255 [0.029]** | — | — | 0.274 [0.030]** | — |
| HO: Banking*Dereg. + Banking*Dereg*Non-Manager = 0 | | | | | | |
| F statistic | | 3.91 | 4.08 | | 1.45 | 1.60 |
| p-value | | 0.053 | 0.049 | | 0.233 | 0.212 |

Notes: Data are from the U.S. Bureau of Labor Statistics Employer Cost Index surveys, 1987–2001. In addition to the covariates shown, regressions included a full set of state-year fixed effects, banking-specific year effects, banking-specific state effects, and interactions of banking with occupation, full-time, and employer size. Robust standard errors, clustered at the state level, are in brackets. N = 331,818.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

have come about in any of several ways. To understand more about how these changes occurred, I use more detailed wage equations to examine falling between-occupation inequality following deregulation. I estimate the following version of a wage equation for the ECI sample with unit of observation job j in industry k of state s at time t :

$$(2) \quad \ln y_{jkst} = \beta_0 + \beta_1 \text{BANKING}_k + \beta_2 \text{DEREG}_{st} + \beta_3 \text{MGR}_j + \beta_4 \text{BANKING}_k * \text{DEREG}_{st} + \beta_5 \text{BANKING}_k * \text{MGR}_j + \beta_6 \text{DEREG}_{st} * \text{MGR}_j + \beta_7 \text{BANKING}_k * \text{DEREG}_{st} * \text{MGR}_j + \beta_8 \text{UNION}_j + \beta_9 \text{FULLTIME}_j + \beta_{10} \text{NONMGR}_j * \text{UNION}_j + \beta_{11} \text{NONMGR}_j * \text{FULLTIME}_j + \beta_{12} \text{ADDTLCNTRLS} + \varepsilon_{jkst}$$

where y is either wages or total compensation. The additional controls consist of state-year fixed effects, banking-specific state effects, and banking-specific year effects. The results are presented in Table 6.

I first constrain the effect of deregulation in equation (2) to be the same for managers and non-managers. Results in columns (1) and (4) of Table 6 show that deregulation had no effect on the level of wages or compensation for the average banking employee in the ECI data.²³ I then allow the effect of deregulation on banking employee compensation to differ

²³In contrast, Black and Strahan (2001) found declining average wages in banking using CPS data.

for managers and non-managers. Results in columns (2)–(3) and (5)–(6) indicate that reduced between-occupation inequality following deregulation was accomplished through reductions in bank manager wages relative to non-manager wages. The level effect of deregulation on bank manager wages was large and negative, implying a decline of approximately 17% after deregulation and declines in total compensation of 20%. The triple interaction indicates that non-manager banking employees, however, were totally insulated from these effects. F-tests at the bottom of the table show that narrowing in wages led to significant net gains for non-managers, but the narrowing in total compensation only held non-manager compensation steady.²⁴ In general, estimates are similar across specifications using the two different dependent variables and including interactions of the non-manager indicator with the union and full-time dummies.

Without more detailed information about worker characteristics, it is difficult to know what is driving this relative decline.²⁵ In fact, a decline in manager wages relative to non-manager wages is precisely the opposite of the between-occupation changes documented by Black and Strahan (2001), who found that non-managers lost a larger share of their wages than did managers following banking deregulation. To assess the role that gender and other characteristics may have played in the ECI results, I turn to the CPS data. I first repeat the wage equation

estimates of Table 6 using CPS data. These results appear in the first column of Table 7. Using the specification I applied to BLS data, which excludes any controls for worker characteristics, I again find substantial between-occupation narrowing among banking industry employees following deregulation. Managers' wages declined by almost 10% following deregulation, while non-managers experienced no statistically significant wage changes.²⁶ Interestingly, these point estimates are much smaller using the CPS data; I provide an explanation for this below.

The results reported in Table 7 are from specifications that, reading from left to right, incorporate increasing numbers of controls for worker background characteristics. I find that the results in the first column are substantially robust with respect to including controls for age and race; controlling for education shows that the average banking employee did suffer wage losses after deregulation, although the losses for non-managers were much smaller than those for managers. Controls for gender, however, eliminate the between-occupation narrowing entirely. I conclude that between-occupation narrowing in the ECI data is not driven by trends in banking employee age or race. Rather, like Black and Strahan, I find that declining gender differentials in banking following deregulation are behind the observed decline in manager wages relative to non-manager wages, with a supporting role for changing educational attainment among non-managers.²⁷

Black and Strahan presented compelling evidence that wage declines among men in the banking industry were driven by erosion of rents that men had disproportionately captured in the form of higher wages. Rather than repeat their analysis, I present

²⁴Interestingly, Fortin and Lemieux (1997) found that deregulation explains little of the rising wage inequality between more and less educated workers in the 1980s. One reason may be that deregulation—especially in non-unionized industries—can reduce wage differentials between more- and less-educated workers, as was the case in the ECI data.

²⁵The establishment identifiers in my data should allow me to test the hypothesis that this between-occupation narrowing is an artifact of changing establishment composition in the industry, but unfortunately establishments remain in the ECI data for only five years. Thus the number of establishments identifying this estimate is extremely small. Unreported results suggested that the narrowing occurred within establishments, but, not surprisingly, the effect was not statistically significant.

²⁶Point estimates were similar but not statistically significant when I restricted the CPS data to the same period covered by the ECI data, 1987–2000.

²⁷For the most part, my estimates confirm results published in Black and Strahan (2001). An exception is my finding of no difference between the wage declines experienced by male managers and non-managers. This may be due to differences in coding occupations as managerial or non-managerial.

Table 7. Variations on the BLS Wage Equations Using CPS Data.

| Description | BLS | BLS + Education | BLS + Education, Age | BLS + Educ., Age, Race | BLS + All Background | BLS + All Background Women 1977-99 | BLS + All Back- ground Men 1977-99 |
|--|----------------------|----------------------|----------------------------|------------------------------|----------------------------|--|--|
| | 1977-99 | 1977-99 | 1977-99 | 1977-99 | 1977-99 | | |
| BLS Covariates | | | | | | | |
| Banking*Deregulation | -0.097 [0.023]*** | -0.082 [0.021]*** | -0.076 [0.021]*** | -0.078 [0.021]*** | -0.025 [0.017] | 0.02 [0.017] | -0.113 [0.024]*** |
| Banking*Deregulation* Non-Manager | 0.075 [0.024]*** | 0.061 [0.022]*** | 0.042 [0.021]* | 0.045 [0.022]** | -0.032 [0.016]* | -0.051 [0.015]*** | 0.024 [0.033] |
| Non-Manager | -0.365 [0.011]*** | -0.232 [0.008]*** | -0.187 [0.008]*** | -0.183 [0.008]*** | -0.195 [0.007]*** | -0.243 [0.005]*** | -0.153 [0.011]*** |
| Banking | -0.015 [0.025] | 0.033 [0.024] | 0.029 [0.025] | 0.025 [0.024] | 0.021 [0.022] | 0.015 [0.022] | 0.05 [0.042] |
| Deregulation*Non-Manager | -0.087 [0.014]*** | -0.045 [0.014]*** | -0.045 [0.013]*** | -0.044 [0.013]*** | -0.054 [0.010]*** | -0.036 [0.007]*** | -0.07 [0.015]*** |
| Banking*Non-Manager | -0.269 [0.020]*** | -0.314 [0.019]*** | -0.284 [0.019]*** | -0.283 [0.018]*** | -0.069 [0.014]*** | -0.039 [0.012]*** | -0.217 [0.030]*** |
| Full-Time | 0.458 [0.008]*** | 0.449 [0.008]*** | 0.37 [0.008]*** | 0.368 [0.008]*** | 0.302 [0.006]*** | 0.346 [0.008]*** | 0.236 [0.005]*** |
| Demographic Controls from CPS Data | | | | | | | |
| Age + Age* Year Dummies | — | — | Yes | Yes | Yes | Yes | Yes |
| Age-Sq + Age-Sq*Year Dummies | — | — | Yes | Yes | Yes | Yes | Yes |
| Educ. Dummies + Educ.*Year Dummies | — | Yes | Yes | Yes | Yes | Yes | Yes |
| Nonwhite Dummy + Nonwhite* Year Dummies | — | — | — | Yes | Yes | Yes | Yes |
| Female Dummy + Female*Year Dummies | — | — | — | — | Yes | N/A | N/A |
| Observations | 1,358,171 | 1,358,171 | 1,358,171 | 1,358,171 | 1,358,171 | 609,266 | 748,905 |
| R-Squared | 0.23 | 0.26 | 0.32 | 0.32 | 0.42 | 0.35 | 0.38 |

Notes: Data are from the March CPS, 1977-1999. All specifications also include state-year fixed effects, banking-specific state effects, and banking-specific year effects. Column [7] restricts the sample to women only, and Column [8] restricts the sample to men only. Standard errors are clustered on state. Robust standard errors, clustered on state, are in brackets.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

additional support for this claim in Table 8. Table 8 shows that banking employment overall, and for men in particular, increased following deregulation. Employment gains were greatest among managers, and growth in managerial employment occurred at roughly the same rates for men and women. I find no evidence that occupations in banking became “feminized”; indeed, I find that if anything, banking increased its share of male employees following deregulation. Strong absolute and relative employment growth among male banking employees, combined with evidence that changes in other characteristics of banking employees had little role in wage declines, confirms that wage declines among men in banking following deregulation were due to

the erosion of above-market wages under the regulatory régime.²⁸

In Table 9, I further reconcile the CPS and ECI results. Using the CPS data, I estimate the gender-specific wage equations of Table 7 separately for employees of large and small firms, where small is defined as having fewer than 25 employees.²⁹ Since large

²⁸Even with detailed information on individuals like that in the CPS, it is impossible to entirely rule out changing employee quality as an explanation for wage declines following banking deregulation. However, it seems unlikely that male and female employees would diverge in their levels of post-deregulation quality.

²⁹In the March CPS, firm size is reported in seven categories starting in 1987. Firm size and establishment

Table 8. Effect of Deregulation on Banking Employment from CPS Data.

| | <i>Dependent Variable is Log of:</i> | | | | | | | |
|----------------|--------------------------------------|----------------------|----------------------|------------------------|-----------------------------------|-------------------------------------|---------------------------------|---|
| | <i>Total Emp.</i> | <i>Emp. of Men</i> | <i>Emp. of Women</i> | <i>Managerial Emp.</i> | <i>Managerial Emp. (Men Only)</i> | <i>Managerial Emp. (Women Only)</i> | <i>Managerial Share of Emp.</i> | <i>Women's Share of Managerial Emp.</i> |
| Banking*Dereg. | 0.113 [0.023]*** | 0.199 [0.037]*** | 0.037 [0.025] | 0.187 [0.033]*** | 0.163 [0.044]*** | 0.222 [0.044]*** | 0.039 [0.007]*** | 0.005 [0.012] |
| Deregulation | -0.043 [0.016]*** | -0.063 [0.017]*** | -0.024 [0.016] | -0.038 [0.019]** | -0.037 [0.020]* | -0.047 [0.020]** | -0.004 [0.002]* | -0.001 [0.004] |
| Banking | -4.064 [0.018]*** | -4.853 [0.030]*** | -3.57 [0.020]*** | -3.724 [0.028]*** | -3.791 [0.037]*** | -3.778 [0.038]*** | 0.066 [0.006]*** | 0.012 [0.010] |
| Observations | 2,319 | 2,319 | 2,319 | 2,122 | 2,122 | 2,122 | 2,348 | 2,122 |
| R-Squared | 0.98 | 0.98 | 0.98 | 0.98 | 0.97 | 0.97 | 0.55 | 0.53 |

Data are from the March CPS, 1977–2000. All regressions include state and year fixed effects. Observations are weighted by the square root of observations in the state-year-industry cell.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

establishments are oversampled in the ECI, restricting the CPS sample to employees of larger firms more closely approximates the ECI data. I find large wage declines for male employees of large banking firms following deregulation of the industry. Male banking employees of large firms saw their wages fall by 17% on average following deregulation. This figure matches the declines estimated for managerial banking employees in the Table 6. Again, I find no evidence that wages for managers and non-managers converged following deregulation once gender-specific effects have been taken into account. This is because wages for women, who were over-represented in non-managerial positions, were unaffected by the regulatory change. Interestingly, I find statistically significant divergence between managerial and non-managerial wages for men employed in small banking firms. I return to this point in the following section.

Between-Establishment Changes

As was the case with narrowing between-occupation inequality, a number of channels

might explain the increases in between-establishment inequality that occurred in the banking industry following deregulation. One possibility is that firm size compensation premiums shifted following deregulation, which could have happened if small firms earned a disproportionate share of pre-deregulation rents or large firms switched to hiring more able people relative to small firms when faced with a more competitive product market. In unreported results, I tested for changes in the establishment size-compensation premium among banking employees using regressions of log compensation similar to those in Table 6 by adding interactions of establishment size with the BANKING and BANKING*DEREG indicators.³⁰ I then allowed the effect of establishment size to differ for banking industry employees before and after deregulation. I found no evidence that the establishment size premium for banking employees changed following deregulation. This casts doubt on any explanation for the increase in between-establishment variance

size are not generally the same, but they are likely to be most similar for larger establishments. Thus the comparison between ECI jobs and large firm employees in the CPS is probably one of like to like.

³⁰The results from regressions excluding the interactions showed a positive and statistically significant relationship between establishment size and compensation in the ECI data, just as in the firm size literature. The wage effect of establishment size did not differ between banking employees and employees in other industries.

Table 9. Firm-Size Specific Wage Equations Using CPS Data.

| | <i>Firm Size Less Than 25</i> | | <i>Firm Size More Than 25</i> | |
|---|-------------------------------|----------------------|-------------------------------|----------------------|
| | <i>Female</i> | <i>Male</i> | <i>Female</i> | <i>Male</i> |
| BLS Covariates | | | | |
| Banking*Deregulation | 0.089 [0.089] | 0.174 [0.097]* | 0.012 [0.057] | -0.172 [0.069]** |
| Banking*Deregulation*Non-Manager | 0.048 [0.123] | -0.468 [0.076]*** | -0.018 [0.058] | 0.085 [0.096] |
| Non-Manager | -0.21 [0.033]*** | -0.191 [0.036]*** | -0.287 [0.017]*** | -0.174 [0.018]*** |
| Banking | 0.028 [0.135] | 0.159 [0.224] | 0.036 [0.058] | 0.275 [0.080]*** |
| Deregulation*Non-Manager | -0.025 [0.033] | -0.079 [0.038]** | -0.006 [0.018] | -0.06 [0.021]*** |
| Banking*Non-Manager | -0.213 [0.114]* | 0.00 [0.000] | -0.079 [0.056] | -0.282 [0.093]*** |
| Full-Time | 0.338 [0.010]*** | 0.196 [0.006]*** | 0.339 [0009]*** | 0.249 [0.006]*** |
| Demographic Controls from CPS Data | | | | |
| Age + Age*Year dummies | Yes | Yes | Yes | Yes |
| Age-Sq + Age-Sq*Year Dummies | Yes | Yes | Yes | Yes |
| Educ. Dummies + Educ.*Year Dummies | Yes | Yes | Yes | Yes |
| Nonwhite Dummy + Nonwhite*Year Dummies | Yes | Yes | Yes | Yes |
| Observations | 63,175 | 77,129 | 250,483 | 277,916 |
| R-Squared | 0.27 | 0.28 | 0.37 | 0.39 |

Notes: Data are from the March CPS, 1988–2000. All specifications also include state-year fixed effects, banking-specific state effects, and banking-specific year effects. Robust standard errors, clustered on state, are in brackets.

*Statistically significant at the .10 level; **at the .05 level; ***at the .01 level.

that hinges on a change in the relationship between firm size and compensation in banking.

Alternatively, a change in the distribution of banking establishments over the range of establishment sizes following deregulation would have increased between-establishment inequality. Unfortunately, the ECI's sampling procedures make it an unreliable source of information on shifts in establishment size distributions. An alternative to examining the ECI's establishment size distribution directly is to modify the wage decompositions of Tables 4 and 5 by replacing establishment dummies with measures of establishment size. This procedure decomposes wages into components due to occupation dummies and establishment size using exactly the same method as the initial decomposition between

occupation and establishment dummies. If increases in wage variance attributable to establishment size are similar to those attributed to establishment itself in Tables 4 and 5, then changes in the establishment size distribution likely explain the increases in between-establishment inequality. Using this procedure, I obtain an estimate of the contribution of establishment size measures to post-deregulation wage variance in banking equal to 0.006. This effect is significant at the 10% level and implies that establishment size does explain some of the increase in between-establishment inequality. However, much remains unexplained. This estimate is less than half the increase in between-establishment variance in Table 4 and is only one-quarter the size of the estimate in Table 5.

The role of a changing firm or establish-

Table 10. Distribution of Banking Employment over Firm Size Categories in the CPS.

| | <i>Banking's Share of Employment in Firm Size Category</i> | | <i>Share of Banking Employment in Firm Size Category</i> | |
|----------------------|--|--------------------------|--|--------------------------|
| | <i>Pre-Deregulation</i> | <i>Post-Deregulation</i> | <i>Pre-Deregulation</i> | <i>Post-Deregulation</i> |
| Under 25 | 0.013 | 0.007 | 0.205 | 0.104 |
| 25 to 99 | 0.034 | 0.017 | 0.282 | 0.152 |
| 100 to 499 | 0.021 | 0.020 | 0.207 | 0.185 |
| 500 to 999 | 0.033 | 0.024 | 0.136 | 0.108 |
| 1000+ | 0.014 | 0.026 | 0.327 | 0.553 |
| P-value of Rank Test | 0.3500 | | 0.5000 | |

Data are from the March CPS, 1988–1999.

ment size distribution can also be assessed using CPS data. Here it is possible to look at the firm size distribution directly (albeit in a censored manner) to see if changes in the firm size distribution itself could account for changing between establishment inequality.³¹ If the firm size distribution became more bimodal following deregulation, then between-establishment wage variance in banking would increase as mass is shifted toward the extremes of the distribution, even if the firm size wage premium remained stable. In Table 10, I show the distribution of banking employees by firm size category in the CPS. Banking's share of employment (column 1) in the largest firm size category increased following deregulation, while its share in smaller firm size categories declined. In other words, banking employment at the largest firms grew faster following deregulation than employment at large firms generally, within a given state and year. These trends are consistent with the consolidation through purchases of existing banks that occurred following deregulation and with expansions in the service sector generally. Examination of the share of banking employment by firm size category (column 2) tells the same story,

although rank tests at the bottom of the table reject statistically significant distributional shifts across the two periods. Moreover, it does not appear that a proliferation of small firms drove this change, as both the growth rate and the share of banking employees in small firms fell after deregulation.

A final possible explanation for the increase in between-establishment compensation dispersion is that banking firms became more heterogeneous within establishment-size categories. The evidence on this possibility is multifaceted. First is the fact that establishment size measures only explain a small part of the increase in post-deregulation between establishment variance. Also, recall that the wage and compensation decompositions from the ECI revealed no change in between-job-cell inequality. Thus the increase in between-establishment inequality is not due to increasing inequality within some set of establishments. Finally, Table 9 shows that wages for male banking employees in smaller firms diverged markedly following deregulation.

Suppose we accept that the Table 9 result in the CPS data, which has no firm identifiers, is not driven by increasing within-establishment dispersion based on the ECI results for job-cell variance or to a changing firm size distribution based on establishment size's small contribution. Then we must conclude that increasing wage dispersion among male banking employees of small firms is due to increased heterogeneity in firm-specific pay. Moreover, it seems unlikely that the increase in employment at large banking firms drove

³¹It is possible that the relationship between firm size and establishment size (in banking) changed following deregulation as large firms acquired small establishments. Nevertheless, if larger firms pay higher wages across their establishments, changes in firm size may shed some light on changing between-establishment inequality.

the increase in between-establishment inequality when it is clear from Table 9 that the wages for the highest-paid employees at large banks were falling after deregulation. I conclude that increased heterogeneity among small banking firms following deregulation is most likely responsible for the increase in between-establishment inequality. It is noteworthy that the wages of female employees were insulated from these changes. This suggests a more complicated interaction of gender with increasing firm heterogeneity. Falling wages for non-manager men can be explained by eroding rents, but increasing wages for male managers may reflect a greater propensity for men to head banks, with wage rewards for competition swamping the rent erosion. See Hubbard and Palia (1995) on this effect. This is consistent with evidence from Rhoades (2000), who noted that the number of banking establishments continued to increase over this period despite high levels of consolidation and high turnover among small firms. In other words, following deregulation some employees of small banking firms worked for firms on the brink of collapse and exit while others worked for firms that would soon be merged with Citibank.³²

Effect of Product Market Competition on Non-Wage Compensation

One of the ECI's most interesting features is its detailed information on benefits costs. Although changes in relative levels of benefits do not drive the major changes in the post-deregulation banking compensation structure, changes may nevertheless occur within the benefit structure, particularly if providing incentives for performance becomes more important to firms following deregulation. In addition to hourly wages, the ECI collects information about the average costs of sixteen types of non-wage compensation.³³ I exam-

ined post-deregulation changes in coverage for seven benefits categories that are both large and voluntary: vacations and holidays, leave, non-production bonuses, severance packages, health insurance, retirement packages, and overtime pay.³⁴ The changing benefits choices of banks following deregulation also provide evidence about changing firm priorities following deregulation and insight into how overall compensation inequality followed the same patterns as wage inequality post-deregulation.

In unreported results, the differences in overall coverage rates between managers and non-managers in the ECI followed expected patterns. Managers were more likely than non-managers to be covered by health insurance, to have a retirement plan or a severance package, and to receive leave, but less likely to receive overtime pay. Vacations and holidays were offered to the vast majority of workers, although managers received this benefit at a slightly higher rate. Interestingly, the coverage rate of non-production bonuses, hereafter simply bonuses, was no different between managers and non-managers.

To examine the effect of banking deregulation on benefits provision, I regressed coverage rates on a banking industry dummy, a deregulation dummy, their interaction, and state and year fixed effects. Benefits coverage rates were computed by state-year-industry-occupation cell, and separate regressions were run for each of the seven benefits categories. The rate of overtime pay and receipt of severance packages increased significantly for banking employees following deregulation.³⁵ The rate of bonus pay declined significantly, and the impact of deregulation on the coverage of other forms of compensation was not statistically significant. However, these changes generally fell on either managers or non-managers exclusively. The increase in the coverage of overtime pay was entirely accounted for by increasing overtime pay

³²Several authors have discussed implications of deregulation and increased product market competition for within-occupation inequality (Guadalupe 2003; Hubbard and Palia 1995). Unfortunately, the ECI data are not well suited to studying within-occupation inequality, and experimentation with analysis along these lines produced only inconclusive results.

³³See Pierce (2001) for data on the average employer's contribution by each ECI benefit category.

³⁴A job was considered covered by a benefit if an employer's expenditures on the benefit were non-zero for that job.

³⁵Mean coverage rates within cells were estimated using corrected BLS weights. Regressions were run without cell size weights, due to the difficulty of computing appropriate cell sizes in the size-stratified ECI data.

for non-managers. Conversely, the increase in severance packages and the decrease in bonus pay were both entirely accounted for by changes in the coverage rates for these benefits among managers.³⁶ Following deregulation, the likelihood of receiving leave increased for managers but decreased for non-managers. These off-setting changes appeared as zero net change in leave coverage for all banking employees.³⁷ Coverage rates were unchanged for the vacations and holidays, retirement plans, and health insurance benefits.

Conclusion

I have examined the effects that liberalization of the product market in one particular U.S. industry, commercial banking, had on various dimensions of the compensation structure for industry employees. The easing of banking industry regulations increased competition between banking firms by reducing barriers to entry and expansion. These changes had the potential to alter the banking compensation structure through a number of channels, including the erosion of any non-competitive rents paid out in compensation, a change in firm size distribution in the industry, increased firm heterogeneity, and shifting incentive problems.

I used data with establishment and occupation identifiers from the U.S. Bureau of Labor Statistics' Employment Compensation Index Survey (ECI) to construct measures of the distribution of compensation by state, year, and industry. Variation in the timing

of banking deregulation across U.S. states provided identification of deregulation's effects on these measures at the state-year-industry level. I supplemented this analysis with estimates using data from the March CPS, which has much more detailed data on worker characteristics than the ECI but little information about a worker's establishment of employment.

In the ECI data, deregulation had no impact on overall compensation inequality in the banking industry. However, this concealed a number of important changes within the banking compensation structure. Manager wages fell while non-manager wages held steady, leading to a large decline in between-occupation compensation inequality. In contrast, between-establishment inequality increased dramatically. These changes in inequality came about almost entirely through wage changes as opposed to compensation changes. Finally, deregulation led to shifts in the types of non-wage benefits banking employees received within the main occupation groups of managers and non-managers, although the level of non-wage compensation was unchanged.

Comparisons with CPS data reveal that the ECI does not tell the whole story. The absence of level effects of deregulation in the ECI data is due to the lack of educational controls in that data set. Once the analysis controls for changing educational attainment among banking employees, deregulation is found to have had a small negative effect on overall wages. The more dramatic reversal concerns the decline in between-occupation inequality. Using controls for gender available in the CPS, I found that the decline in between-occupation inequality is due to differential effects on men and women in the banking industry. I also used CPS data to complement results in the ECI regarding the increase in between-establishment inequality. Combining several pieces of evidence, I concluded that increased heterogeneity among small banking firms following deregulation is most likely responsible for the increase in between-establishment inequality.

In cases where my analysis coincides in focus with other banking industry compensation studies, its results generally support

³⁶This is interesting in light of a debate between Hart (1983) and Scharfstein (1988) concerning the effect of increased product market competition on managerial slack. Their opposing results hinged on different assumptions about managers' risk aversion, especially managers' concerns about "the worst state," in which they lost their jobs. Alternatively, this shift may indicate the advent of tough times for firms, with lower profits forcing them to cut bonuses and undertake layoffs, necessitating severance pay.

³⁷Gender trends may have played a role in these changes. Although effects of federally mandated maternity leave benefits are captured by year effects, banking may have operated on a different schedule with respect to these benefits.

theirs. The effects of deregulation on the banking compensation structure align well with the effects of liberalization that have been observed in other industries (Rose 1987; Card 1996). The banking example suggests that the compensation structure effects of product market liberalization depend primarily on wage setting institutions in the regulatory period rather than on the capital-labor ratio or other industry features, a point also made by Borjas and Ramey (1995).

For those concerned about the effects of deregulation on inequality among workers, the example of the banking industry provides mixed conclusions. The evidence contradicts popular assumptions that competition and liberalization inevitably increase wage inequality and harm low-skilled workers. For example, the cost-cutting pressures of increased product market competition likely eroded rents that were

the source of managerial wage premiums in the banking industry, accruing in particular to men. On the other hand, regulatory restrictions on entry and scale had created an industry in which banking firms were of similar size and quality and thus paid similar wages, leading to more homogeneity in compensation for banking employees in the regulatory era than would be the case following deregulation. Deregulation also led to changes in non-wage compensation that suggest a change in the work environment, with welfare consequences for workers that are difficult to quantify. As this example makes clear, no single measure of inequality will capture all relevant compensation structure changes that accompany liberalization. Taken together, these results underscore the importance of evaluating several dimensions of the compensation structure before drawing conclusions about the effects of a regulatory change on workers.

Data Appendix

A. ECI Sample Construction

The primary data source used in this paper is the U.S. Bureau of Labor Statistics (BLS) Employment Cost Index (ECI) micro data. ECI data are collected by BLS for the purpose of tracking wage and benefits levels, coverage, and costs (to employers) across industries and geographic areas. BLS gathers the data from surveyed establishments, where an establishment is defined as all employees of a firm at a given physical address. The establishment sample is drawn from within strata defined by establishment size, industry, and location, excluding establishments in government and agriculture. Establishments in larger size strata face higher probabilities of selection (U.S. Department of Labor 1997).

In keeping with their purpose as a measure of employer costs, the ECI data contain no individual worker-level or demographic information. Instead, the observational unit in the ECI is the "job." One should think of a job as an occupation–union status–full-time status cell.³⁸ Job-level data represent averages over all employees in sampled jobs. Data for between two and eight jobs are collected from each establishment, depending on the establishment's size. Given their size, banking establishments typically provide data on four to six jobs.

The resulting data set is a rotating panel. Establishments are surveyed annually but new samples are drawn from the universe only every five years, and jobs are not re-sampled over an establishment's time in the survey. Each job observation includes the job's average wage within the establishment, constructed from quarterly hours and earnings; the average expenditures on fifteen different benefit categories for employees in the job; a one-digit occupational code; a four-digit industry code; full-time or part-time status; union status; state; the establishment identifier; number of employees in the establishment; and, in some cases, a county FIPS code identifying the establishment's location.

B. Sample Refinements in This Paper

The ECI data begin in 1979, but the sample in this paper is restricted to the years 1987–2001. The earlier data are discarded because the ECI did not constitute a reliably representative sample of industries or jobs in those years. Rather, some emphasis was placed on collecting data on industries and jobs in which the government had a greater interest, and the ECI gave banking particularly short shrift in the 1979–1986 period. Job observations with zero hours recorded and with hourly wage rates in the top or bottom 0.5% of the distribution in each year were also discarded.³⁹ The remaining sample consists of approximately 330,000 job observations over a 15-year period. About 10,000 of these are in banking. The ECI constructs weights to produce representative estimates from its stratified random sample. Unfortunately, no effort is made to replace establishments that drop out of the sample prior to the 5-year replenishment, and the weights are only recalculated at the time of replenishment. To correct for the fact that recently sampled industries will constitute a disproportionate share of the ECI (relative to older industry samples), the ECI weights were rescaled to total to industry employment in each year as published in the County Business Patterns Data.⁴⁰

³⁸For example, two bank tellers in the same establishment would be considered to be in the same job if both were also full-time employees and not union members. The two would be considered to be in different jobs if one was employed part-time and the other full-time. In the first case, their information would be averaged together to form the job observation. In the second, their information would go into different job averages.

³⁹This follows established procedures for working with the ECI. See Cannon et al. (2000).

⁴⁰Thanks to Brooks Pierce for providing these data and coding for the adjustment.

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