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Sincerely,

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Fourth Quarter 2009, Volume XXXIII, Issue 4

How will baby boomer retirements affect teacher labor markets?

Daniel Aaronson and Katherine Meckel

Introduction and summary

Teachers play a vital role in their students' educational performance. In addition, there is a correlation between a teacher's experience and her effectiveness in the classroom—at least in the first few years of her career. These intuitive outcomes are supported by a large body of research literature.[1] With this in mind, it is reasonable to view rising rates of teacher turnover (since the early 1990s) as a cause for concern. Further, we expect that retirements, which have driven some of this increase, will accelerate to record levels in the coming decade as growing numbers of baby boomers reach retirement age.[2] This pattern will inevitably necessitate a significant increase in the demand for new teachers. Some communities—for example, poor urban districts, which tend to have especially high teacher turnover rates and severe recruitment problems[3]—might be particularly susceptible to declining teacher quality as a result of increased retirements.

In this article, we use a simple model of teacher demand and supply in order to gauge the implications of baby boomer retirements on the projected demand for new teachers. Our forecast links estimates of demand for all teachers with the expected supply of returning teachers through 2020 (that is, the 2020–21 school year). We assume any shortfall would have to be addressed by hiring additional teachers. We discuss how projected demand for new teachers compares with the past half century and what types of schools are likely to have to augment their teacher hiring over the coming decade. We also calculate how much teacher salaries would have to increase in order to fill the gap between teacher supply and demand. To compute the supply and demand of the teacher market, we use a variety of data sets and sources—for example, the U.S. Census Bureau's *Decennial Census* and *Current Population Survey* (CPS) and various publications of the U.S. Department of Education's National Center for Education Statistics (NCES), including its 2003–04 *Schools and Staffing Survey* (SASS) and the accompanying 2004–05 *Teacher Follow-up Survey* (TFS).

We estimate the number of new full-time public school teachers[4] needed from 2009 through 2020 will be between 2.3 million and 4.5 million, with the range encompassing reasonable assumptions about fertility rates, student–teacher ratios, and turnover propensity. Our preferred calculations—based partly on the latest teacher data available from the 2003–04 school year (and therefore not accounting for the economic downturn that began in late 2007)—predict roughly 277,000 new full-time public school teachers needed in 2009–10, rising to 303,000 new teachers by 2020–21, or 3.5 million for all school years between 2009–10 and 2020–21. Retirements account for about one-third of the teachers who leave the teaching work force over this period. Adding the private school sector to these calculations raises the number of new teachers needed by about 20 percent, to 4.2 million, but lowers the fraction due to retirements by roughly 3 percentage points.

These numbers, in isolation, are difficult to assess without some historical context. Therefore, we provide rough estimates of projected demand for new teachers over the past six decades using U.S. *Decennial Censuses*, combined with analogous hiring projections for the years 2010 and 2020. We find that more teachers will retire between 2010 and 2020 than in any other decade since the end of World War II. But because of relatively slower projected growth in the school-age population, the total number of new teachers needed for all reasons (including retirements) is within historical norms. Indeed, normalized by the size of the aggregate labor

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force (one rough measure of the potential teacher work force), demand for new teachers will be similar in magnitude in the coming decade to that in past decades. Therefore, we would not expect the increase in forthcoming retirements, *in the aggregate*, to have a significant impact on national levels of teacher hiring much beyond the variation in teacher hiring needed in the past.

However, it is still possible that certain areas will be especially hard hit by teacher retirements. Therefore, we explore how demand for new teachers is likely to vary based on a school's regional location, urban or rural status, share of free or reduced price school lunch recipients, and racial composition. We find that the need for new teachers is likely to be notably elevated in schools with a high fraction of minority or low-income students. However, this is not driven by an abnormal number of upcoming retirements in these schools, but rather by a combination of elevated teacher turnover rates and expected student population growth. For example, schools with minority representation in the top quartile of the distribution are expected to require 65 percent more new teachers than schools with average minority representation, but only 4 percent of this *difference* is due to retirements.

It is important to emphasize that our estimates are based on a mechanical model of teacher labor markets that assumes some key factors related to the propensity to enter and exit the teaching profession—such as compensation, pension packages, certification requirements, and tenure decisions—will look like they have in the recent past.[5] Difficulty in hiring or retaining teachers could lead local communities to change policies in a way that influences the supply of available teachers. However, many communities, especially those that face the most significant change in hiring over the coming decade, could find making the necessary policy changes challenging. To quantify this difficulty, we provide a very simple calculation of how compensation would have to change in order to offset elevated hiring requirements and keep teacher quality relatively stable; this exercise assumes that salary adjustment is the sole tool schools use to satisfy their growing demand for teachers. We find that real salaries would have to rise by an additional 10 percent beyond historical averages between 2009 and 2020. Pay would have to be particularly bolstered in heavily poor and minority schools in order to offset their expected demand for new teachers over the coming decade.

This article is organized as follows. In the next section, we explain our algorithm for projecting the demand for new teachers in the coming decade. Then we describe the results and provide some historical context by comparing our projections with similar estimates from the past half century. We also explore how our estimates differ by various school characteristics. Next, we ask how compensation policy might have to be adjusted, given estimates of the labor supply elasticity of teachers, to account for any additional hiring requirements in the future. We acknowledge that our data do not cover the period since the current economic downturn began; therefore, we briefly explore some channels in which the current recession might affect short- and long-run demand for new teachers.

A mechanical model of demand for new teachers

In this section, we describe the algorithm we use to forecast demand for new teachers.[6] To provide further intuition for our methodology, we also present a very simple numerical example in the accompanying box that, for expository reasons, strips the model to its bare minimum. Similar models are presented in Hussar (1999).

Demand for teachers

We estimate future demand for teachers by projecting student enrollment through 2020 (that is, the 2020–21 school year). Student enrollment is forecasted based on projections of the five-year-old population, estimates of the propensity to attend public school kindergarten, and estimates of grade progression rates. We then apply a student–teacher ratio to get the total number of teachers needed to fill classrooms to accommodate these students.

We begin with a baseline of the most recent count of students, broken down by grade, compiled by the National Center for Education Statistics for the 2003–04 school year. Each of these students is assumed to advance through the public school system based on estimated grade-specific progression rates calculated by the NCES for the 1999–2000 school year through the 2002–03 school year and displayed in table 1.[7]

New cohorts are added each year to kindergarten based on U.S. Census projections of five year olds corresponding to that school year[8] and the average fraction of five year olds that have attended public school kindergarten in the recent past. Since the mid-1980s, the share of five year olds attending public school kindergarten has varied a bit over time, but not in a way that suggests a trend.[9] Therefore, we project forward using 0.878, which is the average share of five year olds who attend public school between 1999 and 2003, the last four years for which data are available.

To get a final count of classrooms, we apply student–teacher ratios to our student totals based partly on forecasts from table 33 of Hussar and Bailey (2007), which include high, middle, and low scenarios. In the

BOX 1

A very simple numerical illustration

Assume that there are five types of teachers in time t (that is, $T_{1t}, \ldots, T_{5t}$) distinguished solely by their age. Age determines turnover propensity. For example, the youngest group, T_{1t}, exits with probability $p_{1.} = 0.2$ in every year; groups T_{2t}, T_{3t}, and T_{4t} exit with probability $p_{2.} = p_{3.} = p_{4.} = 0.1$; and the oldest group T_{5t} leaves with probability $p_5 = 0.5$. We simplify the example by assuming that all teachers work full time and are of the same gender and that experience, tenure, and age are perfectly collinear; but in the article, probabilities for various transitions, including exits, are estimated by age, experience, tenure, and gender (see table 2, p. 7). Further, assume there are 1,000 teachers in the initial year ($t = 1$), split evenly across the age groups. At the end of that initial year, total turnover $Exit_{.1} = \sum_{i=1}^{5} T_{i1} \times p_{i.} = 200 \times 0.2 + 200 \times 3 \times 0.1 + 200 \times 0.5 = 200$, where $i = 1$ to 5 is the age group. The overall turnover rate is 200/1000 = 0.2.

Further, assume that the demand for teachers grows by 1 percent per year because of changes in student–teacher ratios and growth in the new kindergarten cohorts. We abstract from issues related to grade progression rates for simplicity. Prior to the second year, the teacher work force must therefore expand by $Expand_{.1} = \Sigma T_{i1} \times 0.01 = 1000 \times 0.01 = 10$ *teachers*.

Together, this implies that $Demand\ for\ New\ Teachers_{.1} = Exit_{.1} + Expand_{.1} = 210$, the number of teachers that must be hired prior to the second year.

These 210 new teachers are assumed to have certain characteristics estimated from earlier cohorts of new teachers. To keep the example simple, assume that half the new teachers fall in age group 1 ($z_{1.} = 0.5$) and the other half in age group 2 ($z_{2.} = 0.5$). Therefore, we add 105 teachers to T_{12} and T_{22}. Note, that in our simulations, all returning teachers are made older by a year between the first and second years. We abstract from that in this example, but it would imply that teachers would move between age types each year.

In the second year, the new distribution of teachers for group i is: $T_{i2} = T_{i1} \times (1 - p_i) + Exit_{.1} \times z_{i.} + Expand_{.1} \times 0.01 \times z_{i.}$. The first term is the number of returning teachers of type i, the second term is the number of new teachers of type i replacing any who exit, and the third term is the number of new teachers of type i hired because of expansions in the teacher work force. So, for example, the number of young teachers in the second year is $T_{12} = 200 \times (1 - 0.2) + 200 \times 0.5 + 1000 \times 0.01 \times 0.5 = 265$. Once we have a new distribution of teacher types, we again apply turnover propensities, add in growth to demand, and then compute the number of new teachers needed prior to the third year. This algorithm continues through the forecast horizon.

middle scenario, student–teacher ratios in public schools decline by roughly 1.2 students per teacher between 2004 and 2016. However, because of concern that these trend projections overweight the substantial decline in student–teacher ratios seen in the 1990s, our baseline assumption uses the average of the high scenario and a flat student–teacher ratio. Empirically, this is roughly equivalent to using the average decline between the school years 1999–2000 and 2003–04, the most recent years of data available.[10]

Figure 1 summarizes the projected demand for public school teachers. The solid line provides our best guess estimate, with the shaded range allowing for plausible deviations for the student–teacher ratio and the population growth rate of five year olds, as explained previously.

Supply of teachers

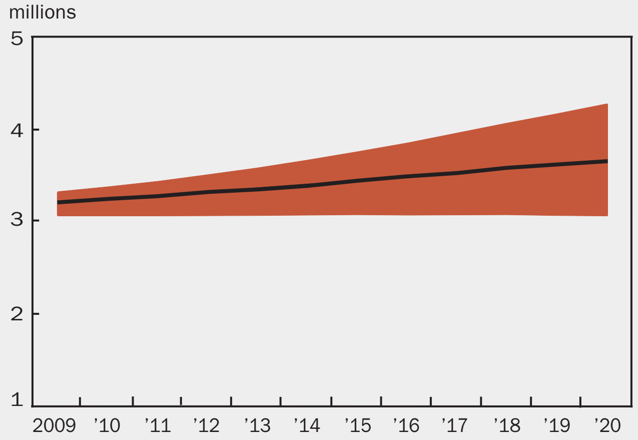
To project the supply of teachers, we begin again with the latest detailed accounting of market size—this time taken from the 2003–04 SASS. That survey tells us there were just over 3 million full-time public school teachers.[11] From the 2004–05 TFS, we can compute that 7.8 percent left public school teaching the following year. An additional 89.6 percent continued as full-time teachers and 2.6 percent as part-time teachers. Note that the latter two rates encompass teachers who switch their work time commitment (full-time versus part-time positions), as well as those who keep the same work time commitment as the previous year.

But all of these rates differ substantially by age, gender, experience, and tenure in the current job.[12] Examples of this heterogeneity are displayed in figures 2's panels A and B, which show how rates of exiting the teaching profession and staying as full-time or part-time teachers differ by age and gender. For example, not surprisingly, exit rates rise monotonically after age 50, and part-time status appears to be higher among women throughout most of the age distribution.

To compute future changes in teacher hours, we simulate the fraction of hours that are likely to return each year by using a simple ordered probit model that allows for three possible transitions—exit, full-time to part-time, and part-time to full-time—and accounts for differences by age, gender, experience, and tenure.

FIGURE 1

Projected demand for public school teachers, 2009–20



Notes: The solid line is our preferred forecast. The shaded region encompasses the plausible range of estimates. Demand for public school teachers was estimated for 2004–08 (not shown), since actual hiring data for those years are not available, and then carried through 2020. See the text for details.
Sources: Authors' calculations based on data from the U.S. Census Bureau, population estimates and projections; U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey* and compilation of yearly national student counts by grade and type of school, 1999–2003; and Hussar and Bailey (2007), table 33.

Table 2 reports the results of our baseline regression. We use the coefficient estimates from this regression to assign an end-of-school-year outcome for all individuals in the 2003–04 cohort based on their personal characteristics. This computation provides us with a forecast of the number of returning teacher hours in 2004–05. We then add a year to each returning teacher's age, experience, and tenure. We continue to project this cohort through the forecast horizon (2020–21), using the same procedures (appropriately adding a year to each returning teacher's age, experience, and tenure).

Three points about the simulations thus far are worth noting. First, it is well known that exits are especially high in the first few years of teaching. That pattern is clearly evident in figure 2, panel A, which displays the hump in exits for women in their late twenties and early thirties. A similar pattern exists when exits are plotted against tenure or experience. We include a dummy for the first five years of experience to account for this nonlinearity.

Second, because of data limitations, tenure is measured as *consecutive* years as a public school teacher, whereas experience is measured as the total number of years as a public school teacher. Unsurprisingly, our results are nearly identical if we exclude the tenure measure.

Third, we do not focus solely on exits because transitions between part-time and full-time teaching positions clearly affect changes in the total number of teacher hours. Specifically, figure 3 (p. 8) shows that part of the growth in teacher hours between 2003–04 and 2004–05 arises from a larger fraction of teachers switching from part-time to full-time positions than vice versa. The exit rate alone is, on average, 7.8 percent; however, accounting for changes in hours from switches between part-time and full-time positions effectively lowers the overall hours turnover rate to 6.5 percent.[13]

To this point, we have described how we project staffing levels due to the work choices of the existing cohort of 2003–04 teachers. But, each year, demand for classrooms exceeds the number of returning teachers; therefore, new instructors must be added to account for those hours. In the simulation, this deficit is filled by adding the appropriate number of "missing" hours to the model each year, while assigning them the age, gender, experience, tenure, and part-time/full-time status that replicates the distribution of characteristics of new teachers in the most

TABLE 1

Average public school grade progression rates

	Progression rate
Kindergarten to 1st grade	1.064
1st to 2nd grade	0.987
2nd to 3rd grade	1.008
3rd to 4th grade	1.003
4th to 5th grade	1.004
5th to 6th grade	1.016
6th to 7th grade	1.015
7th to 8th grade	0.997
8th to 9th grade	1.133
9th to 10th grade	0.892
10th to 11th grade	0.910
11th to 12th grade	0.934

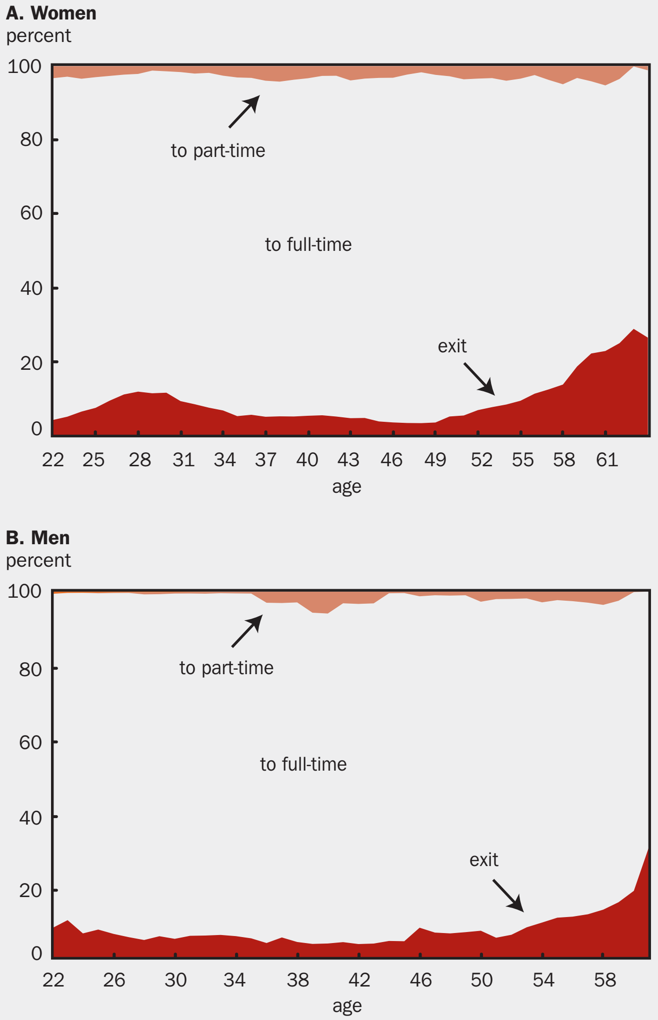
Note: A number above 1 implies a net influx of students coming into public schools in that particular grade from either private schools or schools without grade levels or home schooling; or it implies an influx of children entering the U.S. school system for the first time (recent immigrants).
Source: Authors' calculations based on data from the U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, compilation of yearly national student counts by grade and type of school, 1999–2003.

recent SASS.[14] We then update all of the returning teacher characteristics by making them older an additional year (and giving them an additional year of total work experience and tenure) and rerun the simulations to the following year, using the transition probabilities inferred from table 2. We continue this algorithm through the 2020–21 school year, continuously replacing missing teacher hours with representative entrants and updating the tenure, total experience, and part-time/full-time status of those remaining. This methodology assumes that schools will continue to hire teachers from the same demographic (that is, gender, age, experience, and tenure) background as they have in the recent past and that the part-time and full-time fractions by age and gender stay constant.[15]

We can ascertain the importance of teacher retirements in two ways. First, the follow-up survey asks the reason why teachers exit the profession. Retirement is listed as the reason for roughly 32 percent of all exits at the end of the 2003–04 school year, including 70 percent or higher among exits of teachers aged 55 and older. However, our prediction methods cannot distinguish between reasons for exiting the teaching profession. Therefore, we compute the probability a future exit is due to retirement based on the actual gender and age exit rates displayed in panels A and B of figure 2 and their correlation with the reason for exit in the 2003–04 SASS. Exits rise by 1.2 percentage points, on average, per year for ages 50–60, with overall turnover rates hitting close to 30 percent shortly after age 60. Note that turnover is also high among young and inexperienced teachers, especially women, who represent the bulk of the new teachers. Again, because we tend to exchange exiting teachers with these high-turnover replacements, retirements further amplify demand for new teachers by temporarily introducing high-turnover employees into the system.

FIGURE 2

Teacher transition rates, by age



Notes: The transition rates are computed as five-year weighted moving averages. To full-time and to part-time rates encompass teachers who switch their work time commitment as well as those who keep the same work time commitment as the previous year.

Sources: Authors' calculations based on data from the U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey* and 2004–05 *Teacher Follow-up Survey*.

Demand for new teachers

Lastly, for each year, we compare returning teacher supply (that is, how many teachers are left from the 2003–04 cohort and each subsequent cohort of new teachers) with demand. The additional teachers needed to fill the gap between supply and demand are what we call the demand for new teachers.

Basic estimates of demand for new teachers

Figure 4 provides several estimates of demand for new teachers through 2020. First, concentrate on the solid line, which is our preferred estimate of future demand for new teachers. In this scenario, just under 280,000 teachers are added in the 2009–10 school year, or about 9 percent of the projected 3.2 million teacher work force. Over the coming decade, the total number of new teachers needed to fill growth in demand,

TABLE 2

Impact of age, gender, experience, and tenure on teacher labor market transitions

	Coefficient	Standard error		Coefficient	Standard error
Age 22, Male	8.093646	29105.29	Age 22	0.584441	0.000148
Age 23, Male	0.024761	0.000281	Age 23	0.50437	0.00011
Age 24, Male	–0.10479	0.000203	Age 24	0.561063	0.000106
Age 25, Male	–0.15153	0.000187	Age 25	0.34095	0.000103
Age 26, Male	1.350913	0.000296	Age 26	0.133373	0.000095
Age 27, Male	0.206976	0.000176	Age 27	0.11059	9.81E-05
Age 28, Male	0.930092	0.000201	Age 28	–0.19276	8.95E-05
Age 29, Male	0.804145	0.000193	Age 29	–0.10115	8.75E-05
Age 30, Male	0.392089	0.000191	Age 30	–0.02861	8.89E-05
Age 31, Male	0.176934	0.000179	Age 31	0.266742	0.000106
Age 32, Male	0.563326	0.000184	Age 32	–0.10671	0.000104
Age 33, Male	–0.11639	0.000198	Age 33	0.37932	0.000102
Age 34, Male	0.486308	0.00019	Age 34	0.04453	8.92E-05
Age 35, Male	0.060937	0.000193	Age 35	0.220451	0.000109
Age 36, Male	0.456112	0.000205	Age 36	0.318921	0.000109
Age 37, Male	0.591015	0.000249	Age 37	0.394561	0.000106
Age 38, Male	–0.27538	0.000178	Age 38	0.198718	0.000121
Age 39, Male	–0.05095	0.000223	Age 39	0.027583	0.000117
Age 40, Male	1.000113	0.000365	Age 40	0.10183	0.000102
Age 41, Male	-0.64211	0.000175	Age 41	0.614629	0.000113
Age 42, Male	0.492821	0.000233	Age 42	0.279795	0.000106
Age 43, Male	0.462676	0.000204	Age 43	0.161943	9.89E-05
Age 44, Male	0.017997	0.000201	Age 44	0.039813	9.88E-05
Age 45, Male	0.784613	0.000246	Age 45	0.10397	9.84E-05
Age 46, Male	–0.00701	0.000216	Age 46	0.528427	0.000114
Age 47, Male	0.305273	0.000236	Age 47	0.645033	0.000105
Age 48, Male	–0.79063	–0.00018	Age 48	0.37092	0.000106
Age 49, Male	0.209362	0.000194	Age 49	0.390664	0.000109
Age 50, Male	0.480825	0.000239	Age 50	0.257623	0.000102
Age 51, Male	0.014518	0.000198	Age 51	0.247397	9.66E-05
Age 52, Male	0.126609	0.000162	Age 52	0.003844	8.52E-05
Age 53, Male	0.047295	0.000181	Age 53	0.049014	9.13E-05
Age 54, Male	0.361217	0.000156	Age 54	–0.19105	8.79E-05
Age 56, Male	0.007383	0.000149	Age 55	–0.10263	0.000082
Age 57, Male	0.214801	0.000159	Age 57	–0.24693	9.67E-05
Age 58, Male	0.510085	0.000205	Age 58	–0.4417	0.000124
Age 59, Male	0.77383	0.000189	Age 59	–0.58112	9.96E-05
Age 60, Male	–0.23118	0.000199	Age 60	–0.33222	0.000104
Age 61, Male	0.055861	0.000245	Age 61	–0.84614	0.000143
Age 62, Male	0.262981	0.00029	Age 62	–1.05536	0.000138
Age 63, Male	–9.97802	19421.77	Age 63	–0.05081	0.000185
Age 64, Male	–1.72697	0.000889	Age 64	–1.13825	0.000121
Age 65, Male	–1.89667	0.000572	Age 65	0.190768	0.000311
Age 66, Male	–9.83992	54831.4	Age 66	–0.24195	0.000197
Age 67, Male	–1.72607	0.000544	Age 67	–0.83847	0.000181
Male				–0.11158	0.000108
≤5 years of total experience				–0.17221	3.26E-05
≤3 years of current experience				–0.15405	3.08E-05
4–32 years of current experience				0.057591	3.17E-05
≥33 years of current experience				–0.86998	8.49E-05
Full-time in 2003				0.969324	3.59E-05

Notes: There are three types of transitions from year to year: exiting out of the teacher work force; becoming or remaining a part-time teacher; and becoming or remaining a full-time teacher. The model is estimated using an ordered probit with the data sources.
Sources: Authors' calculations based on data from the U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey* and 2004–05 *Teacher Follow-up Survey*.

FIGURE 3

Transition rates to full-time and part-time teaching positions, by age

Note: These transitions rates are computed as five-year weighted moving averages.
Sources: Authors' calculations based on data from the U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey* and 2004–05 *Teacher Follow-up Survey*.

as well as replace exiting teachers, grows by just over 2,000 per year, hitting 303,000 by 2020. From 2009 through 2020, roughly 3.5 million net teachers need to be added.

The shaded region provides alternative estimates, with the outer ranges suggesting plausible upper and lower bounds of hiring required when we adjust three key factors: the U.S. Census's assumed fertility rate, the estimated teacher turnover rate, and the estimated student–teacher ratio. The fertility rate is allowed to vary by plus or minus 1 percentage point from our baseline, with this range determined by U.S. Census's high and low population projections.[16] The teacher exit rate is also allowed to vary by plus or minus 1 percentage point from our estimated average 7.8 percent baseline rate.[17] This encompasses several alternative estimates, including the 2004–05 TFS turnover rate that weights full-time and part-time teachers equally (8.4 percent) and turnover rates from the *Current Population Survey*'s 2003–04 outgoing rotation file for full-time teachers (8.4 percent) and full-time college-educated teachers with annual incomes between $10,000 and $150,000 (6.9 percent). Finally, the bounds on the student–teacher ratio are allowed to range between the NCES's high assumption projection and a constant ratio based on values from the 2003–04 SASS. The edges of the shaded region use all three assumptions that result in the highest or lowest projection of the demand for new teachers. Taken together, these adjustments broaden the range of plausible new teacher demand to between 2.3 million and 4.5 million from 2009 through 2020. Approximately 42 percent of this range is due to changes in the assumed birth rate, 33 percent to changes in the assumed turnover rate, and 25 percent to changes in the assumed student–teacher ratio.

The dashed line shows the number of new teachers arising from retirements. We find that roughly 30 percent to 35 percent of demand for new teachers between 2009 and 2020 is due to openings created by retirements. Retirements rise from about 82,000 in 2003–04 to just under 96,000 in 2009–10 and average around 96,000 per year over the next decade.

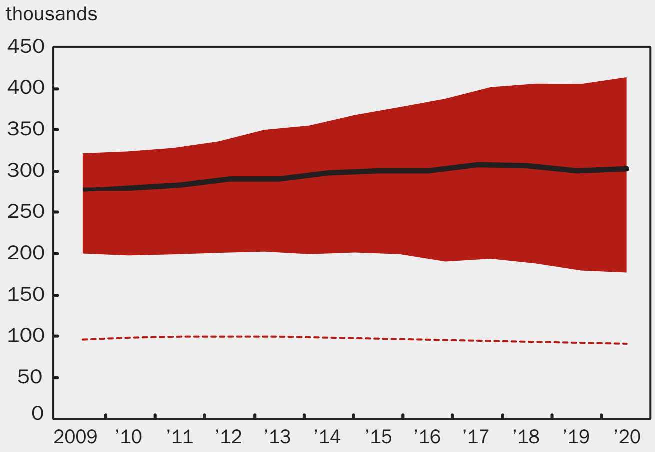
Including private schools

Thus far, we have only included public school teachers. We can compute simple projections for private school new teachers by applying the overall turnover rate of 10.7 percent among private school teachers in the SASS to current staffing levels and NCES projections of the demand for private school classrooms through 2015.[18] In our baseline scenario, private school demand for new teachers rises from roughly 55,000 in 2009 to almost 62,000 in 2015. Projecting this trend forward to 2020 would imply about 725,000 new private school teachers between 2009 and 2020—about a fifth of the public school net demand for new teachers over the same time period.

The ratio of private school students to public school students is about 13 percent, significantly less than the ratio of projected private school to public school new teacher demand. That is mostly explained by a higher overall teacher turnover rate in the private sector (roughly 3 percentage points higher). One consequence of these sector-specific dynamics is that under 10 percent of net private school hiring through 2020 is driven by retirements, suggesting that the retirements of baby boomers will have significantly less impact in private schools over this period. If we aggregate the public school and private school sectors, 29 percent of net teacher hiring is due to retirements (in the SASS)—which is less than the 32 percent of net teacher hiring due to retirements among public schools alone.

FIGURE 4

Projected demand for new public school teachers, 2009–20



Notes: Demand for new teachers was estimated for 2004–08 (not shown), since actual hiring data for those years are not available, and then carried through 2020. The solid line is our preferred forecast. The shaded region encompasses the plausible range of estimates. The dashed line shows openings created by retirements. See the text for further details.

Sources: Authors' calculations based on data from U.S. Census Bureau, population estimates and projections; U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey*, 2004–05 *Teacher Follow-up Survey*, and compilation of yearly national student counts by grade and type of school, 1999–2003; and Hussar and Bailey (2007), table 33.

Are these projections historically high?

Of course, there are always retirements. The key question is how unusual hiring might be given the baby boomer retirements. We provide some historical context by comparing U.S. Census-based estimates of future changes in new full-time public school teacher demand with past changes.[19]

Because of data limitations, we provide very rough approximations of changes in demand for new full-time public school teachers during a decade by adding growth in the full-time teacher labor force to the number of teachers who are of retirement age. The idea behind this calculation, which clearly understates year-to-year hiring, is that it consistently measures all well-observed new full-time public school teachers that 1) fill newly created positions and 2) replace retirees. The number of retirees is conservatively estimated as those who are at least age 55 at the beginning of the decade (and thus retire by age 65). The demand for new full-time public school teachers is plotted in figure 5. We make comparable projections for 2010 and 2020, which vary from our more detailed projections reported previously but are consistent with the historical data.[20]

The red line again shows the rise in new full-time public school teachers needed in the coming decade. But it also shows that the 1970s was a time when hiring was brisk. The reasons, of course, differ. In the 1970s, 72 percent of our new teacher demand measure was necessitated by growing populations of school-age children. By contrast, between 2010 and 2020**,** we expect that only around 31 percent of this measure of the demand for new teachers will be due to student population growth. The remainder will be due to teacher retirements.

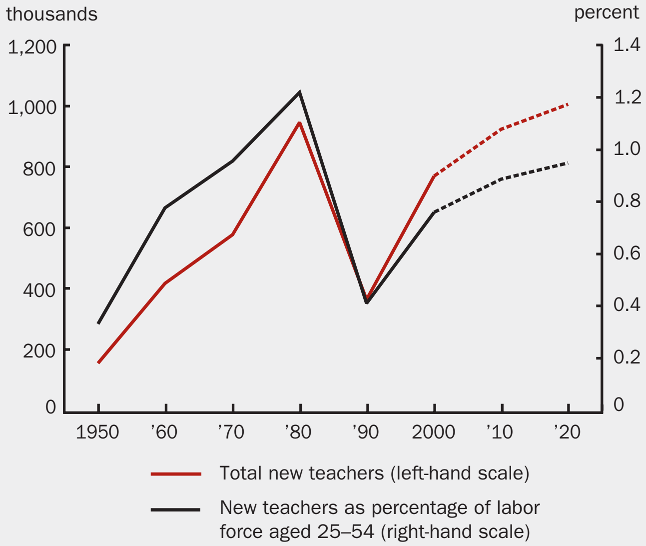
We recognize that comparing absolute numbers is misleading because the size of the aggregate population, and consequently the potential and actual teacher work pool, has grown over time. Therefore, the black line normalizes our new teacher numbers by the population aged 25–54. Here, we find that this ratio is not unusually high right now, nor do we expect it to become unusually high in the near term. Demand for new teachers as a percent of the labor force aged 25–54 is expected to average 0.91 percent between 2010 and 2020—just above the 0.83 percent average between 1960 and 2000 (and below the 0.96 and 1.20 percent levels reached in 1970 and 1980). This suggests modest concerns about filling teacher vacancies in the aggregate.

Demand for new teachers by school demographics

Obviously, not all schools face the same future hiring requirements; the effect of the baby boomer retirements could put particular strain on some more than others. We explore this issue by looking at the key parameters in our forecasts when schools are stratified by region,[21] urban or rural status, share of students receiving free or reduced price lunch, and share of students who are minorities. For each of these categories, table 3 reports the share of teachers over age 50 and 55 (first and second columns of data), the hours turnover rate for new and experienced teachers (third, fourth, and fifth columns), the student–teacher ratio (sixth column), and the growth rate of the student population (seventh column).[22] The eighth column provides results from simulations of new teacher demand, using the same methodology as in figure 4, but assuming that the entire teacher labor market takes on parameters of a subpopulation (as described in the leftmost column). Those numbers are reported relative to the baseline forecasts of the nationally representative population.

FIGURE 5

Demand for new full-time public school teachers



Notes: All part-time teachers are dropped from our calculations. Dashed lines indicate forecasts.

Source: Authors' calculations based on data from U.S. Census Bureau, population estimates and projections.

On average, schools that should expect to see unusually high demand for new teachers are in the West and South; they are located in large cities and small towns; and they educate high shares of minority and low-income students. The particular explanations vary somewhat by school characteristic. However, retirements do not seem to be driving any of the results in an economically significant way. For example, we stratified schools into quartiles based on the fraction of minority students (the bottom two quartiles are aggregated for simplicity).[23] While the top quartile has a higher fraction of teachers aged over 55, there is no statistical difference in the share of teachers aged over 50 across the racial minority representation quartiles. If we leave all parameters at the top quartile's level but switch the age distribution of the teachers so that it matches the schools at the bottom half of the minority representation distribution, overall new teacher demand increases by only 4 percent. Consequently, there is little evidence that baby boomer retirements will affect schools with a high proportion of minority students any *more* than other schools.

Additional hiring demands in schools that have student populations with high minority representation or those with many low-income members (who receive free or reduced price lunches) are driven almost entirely by higher turnover propensity and expected student population growth. Putting all these pieces together, we would predict that if all schools had the characteristics of schools with a high proportion of minority students, the demand for new teachers would be 65 percent higher than the baseline forecasts over the forecast horizon. Over 60 percent of this gap is explained by differences in turnover rates across the age/experience distribution, and just under half by differences in expected student population growth rates. Similar issues arise for schools with high fractions of free or reduced price lunch program participants or for those in urban areas, many of which are also schools with a high fraction of minority students.[24]

What can policy do? The case of teacher compensation

Finally, we ask how policy can respond if community demand for new teachers increases beyond historical norms. Obviously, there are many factors that affect teacher labor supply—a short list of which would include salaries, pension systems, classroom and school conditions, and certification requirements and other barriers to entry. We concentrate on teacher financial compensation because of its relevance to policy discussions and because of the attention that has been paid to its estimation in the literature.

That attention in the literature certainly does not imply a consensus. A number of recent papers have established a link between teacher salary, outside work alternatives, and turnover (for example, Dolton and van der Klaauw 1995, 1999; Murnane and Olsen 1989, 1990; Stinebrickner, 1998; and Harris and Adams, 2007). But others (for example, Scafidi, Sjoquist, and Stinebrickner, 2006; Hanushek, Kain, and Rivkin, 2004; Clotfelter et al., 2008; and Ondrich, Pas, and Yinger, 2008) cast doubt on these findings. We concentrate on the larger estimates of the impact of salary on turnover in the literature and therefore consider our results to be a lower bound estimate of the effect of raising teacher salaries on future demand for new teachers.

We mechanically introduce the impact of an across-the-board salary adjustment to our transition probabilities by adjusting potential exit rates using salary–turnover elasticities from various studies.[25] For our original cohort of 2003 public school teachers, we use the "overall" (that is, representative of the entire public school teacher labor force) elasticity estimate from Harris and Adams (2007). For the

TABLE 3

Key parameters and projected demand for new teachers, by school characteristics

	Age distribution of teachers, 2003–04		Teacher hours turnover rate, 2003–04 to 2004–05			Student–teacher ratio, 2003–04	Percent change of student population growth rate, 2008–20	Percent change in new teachers relative to baseline
	Age >50	Age >55	New	Age 25 to 35	Age >50			
Northeast	0.317***	0.135***	0.063	0.073	0.129	12.8***	–1.9	4.4
Midwest	0.313***	0.121*	0.000*	0.071	0.108	14.2***	0.6	–10.3
South	0.273***	0.118***	0.093	0.055	0.104	14.2***	25.4	27.4
West	0.304	0.140***	0.054	0.107*	0.102	17.9***	24.8	6.2
Percentile of free/reduced price lunch students								
>75 percentile	0.279***	0.120**	0.136**	0.124	0.123	14.4***	—	53.7
50–75 percentile	0.297	0.124	0.029	0.062	0.125	14.3***	—	13.8
<50 percentile	0.305***	0.130**	0.018	0.045	0.098	15.2***	—	–20.9
Large or mid-sized central city	0.301	0.131**	0.137**	0.103*	0.130*	15.4***	—	23.7
Urban fringe	0.296	0.128	0.000**	0.058*	0.096	15.3***	—	–10.4
Small town/rural	0.293	0.114***	0.104	0.063	0.124**	13.3***	—	26.5
Percentile of minority students								
>75 percentile	0.296	0.141**	0.203***	0.114*	0.122	15.5***	21.6	65.4
50–75 percentile	0.291	0.118***	0.034	0.069	0.106	15.1***	10.0	23.0
<50 percentile	0.300	0.122***	0.000**	0.045*	0.106	14.3***	–0.8	–26.2
Observations	42,310		675	1,629	1,673			

*Statistically significantly different from the remaining population at the 10 percent level.
**Statistically significantly different from the remaining population at the 5 percent level.
***Statistically significantly different from the remaining population at the 1 percent level.
Notes: Student growth rates are from U.S. Census Bureau's population projections. All other parameters are computed from the 2003–04 *Schools and Staffing Survey* and 2004–05 *Teacher Follow-up Survey*. Hours turnover accounts for changes between part-time status and full-time status. If switches from part-time to full-time more than offset lost hours through exits and switches from full-time to part-time, we report the hours turnover rate as 0. The final column reports the percent change in net demand for new teachers relative to the baseline if the full population had the subpopulation characteristics.
Sources: Authors' calculations based on data from U.S. Census Bureau, population estimates and projections; U.S. Department of Education, Institute of Education Sciences, National Center for Education Statistics, 2003–04 *Schools and Staffing Survey*, 2004–05 *Teacher Follow-up Survey*, and compilation of yearly national student counts by grade and type of school, 1999–2003; and Hussar and Bailey (2007), table 33.

cohorts of new teachers introduced into our model, we use the "new teacher" salary–turnover elasticities (calculated for teachers during their first five years of teaching) reported in Dolton and van der Klaauw (1995), Stinebrickner (1998), and Harris and Adams (2007).[26] For those new teachers that survive past their fifth year, we switch them to the "overall" exit elasticity once they complete that fifth year. We continue to assume that the fraction transitioning to part-time or full-time teaching positions remains the same; that is, these transitions are unaffected by new salary levels. We also assume that these salary effects do not differ across school types, as described in Hanushek, Kain, and Rivkin (2004).

In the aggregate, we calculate that annual wage growth about 0.8 percentage points beyond average pay growth would offset much of the additional net new demand for teachers over the coming decade, relative to the early 1990s. Specifically, the ratio of teacher hiring to the size of the general labor force between 1988 and 1995 was about 0.00144. We project that this ratio will fluctuate between 0.00178 and 0.00184 during the 2010s. To get the ratio back to 0.00144 by 2020 would require roughly an additional cumulative wage growth of 10 percent between 2009 and 2020. By comparison, cumulative real weekly wage growth of teachers in the *Current Population Survey* was 9 percent between 1989 and 2004, suggesting that the pay hike needed to reach this fairly ambitious target is relatively large.

Teacher pay would have to be especially bolstered in schools with high proportions of poor and minority students in order to offset their expected teaching needs over the coming decade. For example, if a goal was to reduce the demand for new teachers in schools with a high fraction of minority students from 65 percent to 30 percent above baseline national needs by 2020 (thereby offsetting the turnover and retirement rate differences in these schools), average real pay for teachers would have to rise by well over 25 percent.

How does the recent recession impact our estimates?

The data underlying our projections are not available for the current downturn; this is unfortunate, since our predictions could be affected by a significant decline in economic activity. Economic theory predicts at least three ways in which demand for new teachers might be altered by a recession: through changes to the fertility rate, immigration, and teacher attrition (including retirements).[27]

Both fertility rates and net migration flows are commonly observed to fall during recessions, and early indications are that both measures have fallen during the current downturn.[28] Lower migration will reduce demand for teachers now, and lower fertility will reduce demand five years hence. Moreover, children of immigrant parents tend to be disproportionately from low-income families and clustered in a few large urban areas.[29] Therefore, it is possible that lower net migration will help to relieve some constraints in schools where the demand for new teachers is projected to be relatively high.

Lower teacher attrition, and consequently lower replacement hiring, is also possible as household wealth declines and alternative labor market opportunities evaporate. Because of these factors, we would speculate that some additional weight should be placed on our lower bound projections in figure 4 (p. 9) for the next couple of years. Beyond that, we think that consensus economic forecasts[30] imply that these cyclical effects will fade away.

Conclusion

In this article, we provide a simple model of teacher demand and supply in order to gauge the implications of baby boomer retirements on demand for new teachers over the coming decade. We find that the demand for new teachers will rise over the coming decade—and a good portion of this will be due to retirements. That said, we do not expect that this increase in teacher demand will be significantly different from that of past decades, especially relative to the size of the aggregate labor force. However, the added hiring requirements are likely to play out longer than they have in the past, and they will not be equally dispersed across the nation. Moreover, simply raising pay, unless substantially unanchored from past trends, is unlikely to keep teacher quality constant, especially at schools that have traditionally had the most difficulty recruiting and retaining teachers.

NOTES

[1]See, for example, Murnane (1975); Rockoff (2004); Rivkin, Hanushek, and Kain (2005); and Aaronson, Barrow, and Sander (2007). Unsurprisingly, experience is correlated with productivity across a variety of professions. Recent estimates (for example, Aaronson and Sullivan, 2001) suggest that the loss of human capital from the baby boom generation, primarily through lost experience, will be enough to lower the potential rate that the economy can grow during the 2000s by one-tenth to two-tenths of a percentage point per year from its 1990s levels.

[2]Over the past two decades, the share of teachers that have left the profession has been increasing, from roughly 5 percent of teachers in the early 1990s to over 8 percent a decade later. For an example of popular press concern about teacher turnover, see Dillon (2007). See Gordon, Kane, and Staiger (2006) for a discussion of the impact of baby boomer retirements on the teaching profession.

[3]Hanushek, Kain, and Rivkin (2004); and Jacob (2007).

[4]We define a full-time public school teacher as one who works 35 hours per week. Teachers who work fewer hours are counted as fractions of full-time teachers. Unless otherwise indicated, this is how we count teacher demand and supply in our calculations.

[5]Our model also assumes that variability in the business cycle looks similar to that of the past. Given the current deep recession, there may be some concern that this assumption is inaccurate. Unfortunately, key data from recent years are unavailable. Later in the article, we briefly discuss how our estimates might change as a result of the current recession.

[6]Detailed calculations and data are available from the authors upon request.

[7]We thank William Hussar at the NCES for providing us with student counts, by type of school and grade, for the years 1970–2003. We used these data to calculate the progression rates. A number above 1 implies a net influx of students coming into public schools in that particular grade either from private schools or schools without grade levels or home schooling; or it implies an influx of children entering the U.S. school system for the first time (recent immigrants). This is particularly notice able in the transition between eighth and ninth grades and kindergarten and first grade. In order to put heavy weight on more recent history, we used the average from 1999–2003 only. Our results do not change appreciably if we take into account the longer time series.

[8]These are available at www.census.gov/ipc/www/usinterimproj/. These projections are an interim revision of more detailed forecasts released in 2000; the forecasts have been updated to take into account the 2000 U.S. Census. In the sensitivity analysis to follow, we use the high and low series from the original projections released in 2000. We have also tried using the five- and six-year-old population, but this resulted in kindergarten projections that were less accurate when compared with similar NCES student projections.

[9]We calculate the share of five year olds attending public school over time, using population estimates by age from the U.S. Census and the student counts by grade from the NCES. Some of the changes in the share of five year olds attending public school may line up with the business cycle. We found some very mild, but not particularly robust, evidence of procyclicality. But that appears to be driven primarily by a correlated drop in both gross domestic product (GDP) growth and public school attendance of five year olds in the early 1990s.

[10]The average annual drop in the public school student–teacher ratio for the school years 1991–92 through 1998–99 was –0.15, according to Hussar and Bailey (2007), whereas the average annual drop for the school years 1999–2000 through 2003–04 was –0.06. The average decline predicted by the NCES's high assumption for the school years 2004–05 through 2015–16 is –0.10. Mechanically, we compute the student–teacher ratio in our SASS data for 2003 and then apply the year-to-year differences in the average of the NCES's high assumption and a flat student–teacher ratio. Because the NCES ratio is projected out to 2016, we use the rate of growth since 2004 to project ratios further into the future.

[11]This count excludes pre-kindergarten teachers and short-term substitutes. We find that the count of full-time and part-time teachers in the 2003–04 SASS is consistent with similar counts in the U.S. Census Bureau's 2004 *American Community Survey* (ACS) and NCES *Common Core of Data* (CCD). To make the ACS sample comparable, it is necessary to exclude teachers categorized as "other," which includes short-term substitutes and instructors working outside of public elementary and secondary education, and to restrict the sample based on education and salary. Doing so provides a sample that is similar not only in count to the SASS but also in the distribution of education, age, and earnings.

[12]Obviously, there are other teacher characteristics that can affect turnover. To take one important example, Podgursky, Monroe, and Watson (2004) and Corcoran, Evans, and Schwab (2004) estimate the impact of teacher ability on teacher retention. Later, we discuss how the fraction of low-income students, racial composition, urban or rural status, and geographical location of a school affect teacher turnover propensities.

[13]In order to count the number of teacher hours lost in this transition, it is necessary to quantify the "amount of teacher" added or subtracted when a teacher switches from part-time to full-time status or vice versa. The average part-time teacher in the sample works roughly 60 percent of a full-time teacher's hours (which again is defined as 35 hours per week). Therefore, a teacher who switches from full-time to part-time status is counted as 0.6 of her original weight, and a teacher who switches from part-time to full-time status is counted as 1.67, or 1/0.6, of her original weight.

[14]In particular, we use the demographic distribution of teachers who were not teaching in public schools in the prior year. On average, these teachers are eight years younger (and eight years less experienced) than returning public school teachers. Gender composition is nearly identical between the two groups of teachers.

[15]Since the labor force—both teacher and overall labor force—has been growing older, this could reflect the distribution of new teachers as well. In order to examine how demographic change affects our forecasts, we used U.S. Census population projections by age and gender to project the age and gender distribution of new teachers from 2004 through 2020. We found that adjusting our estimates to take into account an increasing fraction of older teacher hires does not make a large difference in our projections. The baseline of new teachers increases by just under 2 percent and retirements increase by 4 percent by school year 2020–21.

[16]The U.S. Census does not report revised interim high and low projections. Instead, we applied the growth rates of the high and low projections from the detailed U.S. Census projections released in 2000 to recalibrate new high and low projections that are in line with the revised middle forecast (see note 8 for further details).

[17]A component of this assumption is the age of retirement over time. We currently assume that the age of retirement stays constant at the 2003–04 level. The median age of teacher retirements in the 2003–04 outgoing rotation files from the *Current Population Survey* is 60.0, quite close to the average median age of 60.3 between 1994 and 2005. We see little evidence of a trend in this series. However, if the retirement age declined by one year during the period 2004–20, this would be equivalent to an increase of

0.6 percentage points in the exit rate over this time period; this implies a 2.2 percent change in net new teachers between 2009 and 2020 if we assume uniform year-to-year increases in the exit rate.

[18]By using the NCES projections, we explicitly accept NCES's assumption that factors influencing private school and public school enrollments, such as transfers to and from private and public schools, migration, dropouts, deaths, and grade promotion will display future patterns consistent with past patterns. The NCES uses past grade progression rates to project enrollment.

[19]For these calculations, we restrict the U.S. Census sample to full-time teachers only. The U.S. *Decennial Censuses* include short-term substitutes and other instructors who do not teach in public schools. Moreover, education requirements and salary ranges change over time. Therefore, it is not clear how to restrict the historical Census samples to make them comparable to SASS counts.

[20]More specifically, these forecasts are computed by applying age-specific teacher retention rates using the U.S. Census Bureau's *American Community Survey* from 2001 through 2006 and the student–teacher ratios and U.S. Census projections of the school-age population as discussed previously. As with calculations using U.S. Census data for new teacher demand in previous decades, estimates for new teacher demand from 2000 through 2010 and from 2010 through 2020 are computed by adding the number of teachers over 65 to the growth in the demand for full-time public school teachers.

[21]Unfortunately, sample sizes preclude reliable results at smaller levels of geography.

[22]We take the student growth rate from U.S. Census population projections, which do not report results by urban or rural status or by free or reduced price student lunch status. Consequently, the rural versus urban status and lunch status simulations assume average student growth over the entire population for all school types. This mechanically mutes the comparison with minority or regional sub-populations. Note also that the hours turnover rate accounts for changes in part-time and full-time status, as well as exits. Because the transition from part-time to full-time is relatively common (see figure 3, p. 8), especially for new teachers, that transition can more than offset hours lost from exit and from full-time to part-time switches. In such cases, we report the hours turnover rate as 0.

[23]For racial composition, the top quartile comprises schools with at least 75 percent minority representation. The second quartile comprises schools with 29 percent to 75 percent minority representation. For free/reduced price school lunch composition, the top quartile comprises schools with at least 61 percent free or reduced price lunch students, and the second quartile comprises schools with 37 percent to 61 percent free or reduced price lunch students.

[24]Another example where inner-city schools might be disadvantaged is described in Boyd et al. (2005). That paper shows that teachers often work in areas near where they grew up. This can make it more difficult to hire teachers for districts with alumni who do not go into the teaching profession.

[25]Of course, an across-the-board salary increase affects the transition decisions of both high- and low-quality teachers. As much as it creates an incentive for the low-quality teachers to stay longer, it may not necessarily improve overall teacher productivity.

[26]It is important to note that in our model, some teachers may be returning to teaching after a break. In the literature, the elasticities tend to be computed for the first five years of teaching experience.

[27]Another channel is larger class sizes. However, we find little evidence of countercyclical movements in class size during the previous two business cycles. That said, pressures on state and local governments have been notably more severe during this recession.

[28]Analysis using the March *Current Population Survey* shows that the number of recent immigrants in 2008 (that is, foreign-born residents who moved to the United States last year) was down 7 percent from 2007 and 30 percent from 2006. The Centers for Disease Control and Prevention's National Center for Health Statistics reports a decline of 0.5 percentage points in fertility rates during early 2008 relative to the previous few years.

[29]Roughly one-fifth of school-age children are from immigrant families, up from 6 percent in 1970, and these children are concentrated in states with the largest urban school districts, for example, California, New York, Texas, Florida, Illinois, and New Jersey (Capps et al., 2005).

[30]See, for example, the Federal Reserve Bank of Philadelphia's *Survey of Professional Forecasters*, available at www.phil.frb.org/research-and-data/real-time-center/survey-of-professional-forecasters/.

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The recession of 1937—A cautionary tale

François R. Velde

Introduction and summary

The U.S. economy is beginning to emerge from a severe economic downturn precipitated by a financial crisis without parallel since the Great Depression. As thoughts turn to the appropriate path of future policy during the recovery, a number of economists have proffered the recession that began in 1937 as a cautionary tale. That sharp but short-lived recession took place while the U.S. economy was recovering from the Great Depression of 1929–33.[1]

According to one interpretation, the 1937 recession was caused by premature tightening of monetary policy and fiscal policy prompted by inflation concerns. The lesson to be drawn is that policymakers should err on the side of caution. An alternative explanation is that the recession was caused by increases in labor costs due to the industrial policies that formed part of the New Deal—the policies of social and economic reform introduced in the 1930s by President Franklin D. Roosevelt. If a policy lesson can be drawn from this, it might have more to do with the dangers of interfering with market mechanisms.

The goal of this article is to present the relevant facts about the recession of 1937 and assess the competing explanations. Although overshadowed by its more dramatic predecessor, the recession of 1937 has received some attention before, in particular Roose (1954) and Friedman and Schwartz (1963). Then, as now, the competing explanations centered on fiscal policy, that is, the impact of taxation and government spending on the economy; monetary policy, or the management of currency and reserves; and labor relations policy, or more broadly government policy toward businesses.

The rest of this article is organized as follows. I first present the salient facts about the 1937 recession. I then review the competing explanations and finally provide a quantitative assessment of their likely contributions to the recession. I find that monetary policy and fiscal policy do not explain the timing of the downturn but do account well for its severity and most of the recovery. Wages explain little of the downturn and none of the recovery.

The recession

Before describing the salient features of the 1937 recession, I first take up the issue of its timing. The traditional National Bureau of Economic Research (NBER) business cycle dates put the peak of the recession in May 1937 and the trough in June 1938. Romer (1994) argues that there are inconsistencies in the way these dates were established over time, devises an algorithm that closely reproduces the dates of post-war business cycles, and applies it to the Miron and Romer (1990) industrial production series to produce new dates. In the case of the 1937 recession, Romer identifies August 1937 as the start of the recession. Cole and Ohanian (1999) implicitly use the same starting date when they state that industrial production peaked in that month. I will stick to the traditional date for several reasons. One is that Romer (1994) directs her argument mostly at cycles before 1927, when a shift in NBER methodology occurred. Another is that the NBER dating process considers a broader set of series than just industrial production. Roose (1954) lists the peaks of 40 monthly series and shows that 27 series peaked before August. Finally, industrial production as measured by the Board of Governors of the Federal Reserve System peaked in May 1937. There is no controversy over the end date of the recession, set by the NBER and Romer (1994) in June 1938.

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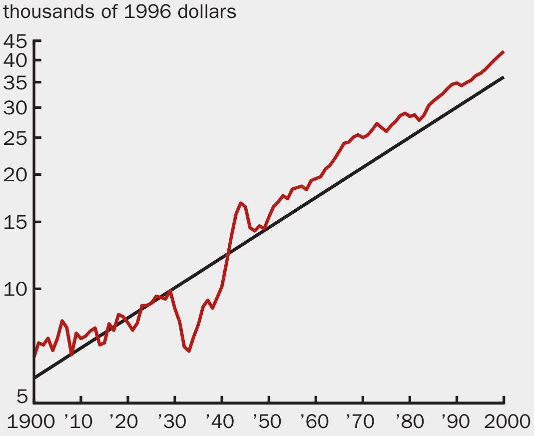
FIGURE 1

Gross domestic product per capita, 1900–2000

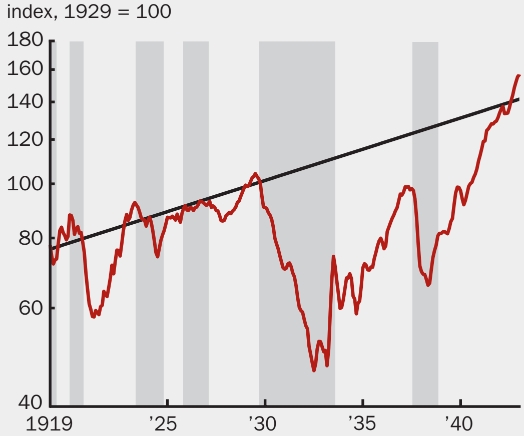


Notes: The population is age 16 and older; gross domestic product per capita is measured on an annual basis over the period 1900–2000. The trend line (black) grows at the average growth rate over the periods 1919–29 and 1947–97.

Source: Author's calculations based on data from Carter et al. (2006), tables Aa125–144 and Ca9–19.

FIGURE 2

Industrial production per capita, 1919–42



Notes: The population is age 16 and older; industrial production per capita is measured on a monthly basis over the period January 1919–December 1942. The trend line (black) is the average growth rate over the period January 1919–August 1929. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Source: Author's calculations based on data from the Board of Governors of the Federal Reserve System, G.17 statistical release, various issues.

Figure 1 plots real annual gross domestic product (GDP) per capita (population aged 16 years and older) over the twentieth century. The trend line follows that series' average growth rate over the periods 1919–29 and 1947–97, and is set to coincide with the series in 1929. This is the metric by which Cole and Ohanian (2004) show that the recovery after the Great Depression was weak, since the series does not return to trend until 1942. The exceptional nature of the Great Depression and the ensuing recovery is starkly evident, but the 1937 recession barely registers in the annual series. The reason is that the recession is so short, beginning in mid-1937 and ending in mid-1938.

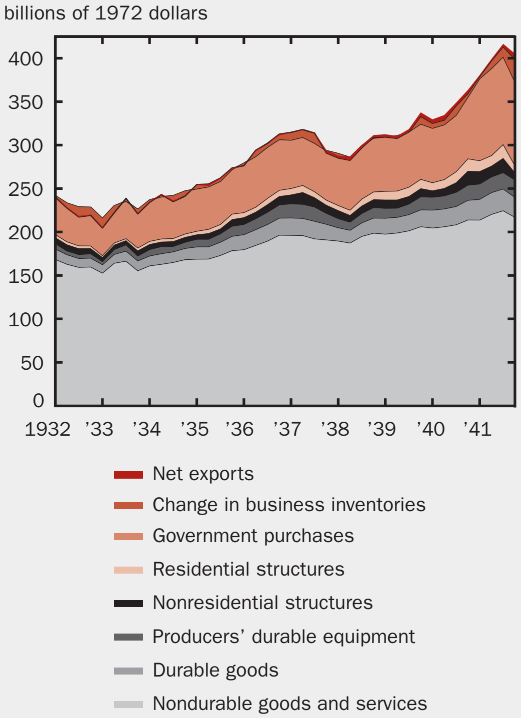
To get a better sense of the importance of this episode, we need to look at higher-frequency data. The national income and product accounts (NIPAs) are not available at the usual quarterly frequency before 1946, however, so we have to resort to other series. Figure 2 plots a monthly index of industrial production, which will be the main focus of my analysis in the final section. Again, a trend line has been added, growing at the average rate of growth for the period from January 1919 to August 1929. The severity of the 1937 recession is now apparent. In particular, it is striking to see that the speed at which industrial production contracted is greater than during the Great Depression. From its peak in July 1937 to its trough in May 1938, industrial production declined 32 percent. By comparison, it took two full years for industrial production to fall as much from its July 1929 peak. Other measures confirm the severity of the 1937 recession—for example, employment fell by 22 percent and stock prices declined by over 40 percent (Carter et al., tables Cb46 and Cb53).

Another striking aspect of the 1937 recession is the recovery that ensued. The rate of growth of industrial production was slightly higher than that which prevailed over the period 1933–37 (22 percent per year compared with 21 percent), and the recovery proceeded smoothly, without the pauses and reversals that marked 1934. Had it not been for the 1937 recession, industrial production would have returned to its trend three or four years earlier.

Although official NIPA data are not available on a quarterly basis during that period, Balke and Gordon (1986) have estimated the components of gross national product (GNP), using regression-based interpolation. Although these estimates should be taken with care, I show them in figure 3; I present the growth rates in table 1 for the period of interest, with the averages for the preceding expansion as the point of comparison. They display some interesting differences of timing with industrial production. Nondurables consumption growth, strong in the last three quarters of 1936, stalled in early 1937 and collapsed in the third quarter. The various components of investment do not show such a clear pattern until the fourth quarter of 1937, when all growth rates turn negative. In contrast, the recovery is firm across all sectors in the third quarter of 1938.

FIGURE 3

Components of gross national product, 1919–41



Note: Data are quarterly over the period 1919:Q1–1941:Q4.
Source: Balke and Gordon (1986).

Fiscal policy

In the 1930s, total government was still a relatively small but growing share of the economy: In 1929 total government consumption and investment represented 9 percent of GDP, and by 1939 it had reached 16 percent. During the same period, the federal government grew in importance relative to the states and local government: Federal spending grew from 1.6 percent to 6.4 percent of GDP.[2] However, figure 4 shows that the stance of fiscal policy at the state and local level did not change much during the period under consideration. I will therefore concentrate on federal finances.

Until the Great Depression, the traditional fiscal policy had been one of balanced budgets. During the early stages of the New Deal, the vast expansion of the federal government was financed through debt, but by the middle of the 1930s, concerns were growing over the size of the public debt, which had gone from 16 percent of GDP in 1929 to 40 percent in 1936.

In 1936, there was a deliberate attempt to return to a balanced budget. Figure 5 shows the components of federal revenues by source and also plots expenditures. On the expenditures side, there is little to note except a very large spike in the second quarter of 1936. This represents the payment of bonuses to World War I veterans, which Congress decided to accelerate that year before the November elections. This probably boosted demand in the last three quarters of 1936 well above its earlier levels (table 1), but it is hard to see how it could have precipitated a recession on its own.

TABLE 1

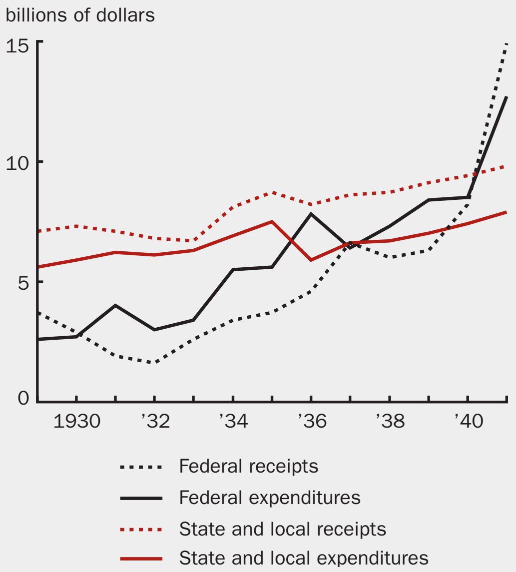
Growth rates of components of gross national product, annualized, 1933–38

	Nondurable goods and services	Durable goods	Producers' durable equipment	Nonresidential structures	Residential structures	Government purchases
	(- - - - - - - -	- - - - - - - -	- - - - - - - -	percent - - - -	- - - - - - - -	- - - - - - - -)
1933:Q1–1935:Q4	3.8	17.2	32.5	8.9	27.5	5.0
1936:Q1	2.5	18.4	4.5	7.1	–26.9	39.8
1936:Q2	10.9	18.1	29.6	–21.8	0.7	11.2
1936:Q3	10.5	21.5	44.6	59.4	93.0	3.8
1936:Q4	14.1	11.6	29.5	47.7	–15.7	–2.0
1937:Q1	–0.4	7.9	26.5	21.9	15.7	–19.9
1937:Q2	–0.3	–7.4	0.5	112.0	36.2	–2.4
1937:Q3	–7.8	10.8	3.9	–96.0	–63.5	2.8
1937:Q4	–2.5	–47.4	–106.0	–54.5	–63.6	7.1
1938:Q1	–2.7	–65.6	–73.0	–3.4	0.0	19.3
1938:Q2	–4.9	–17.4	–45.2	–71.3	13.6	6.9
1938:Q3	15.9	25.7	56.8	41.9	114.1	4.0
1938:Q4	7.7	44.0	43.5	21.2	45.4	3.8

Source: Author's calculations based on data from Balke and Gordon (1986).

FIGURE 4

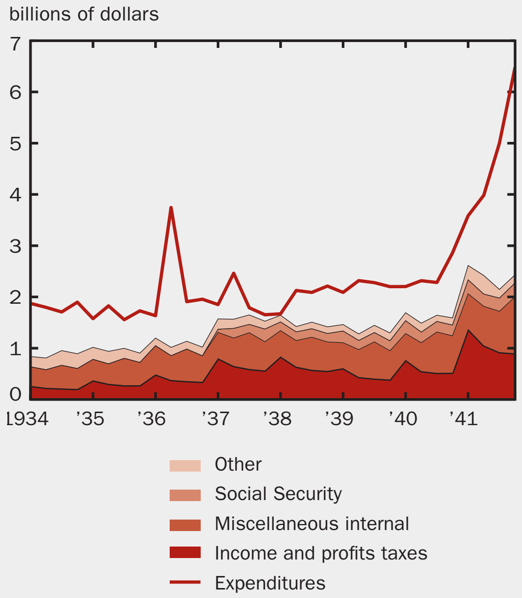
Federal and state and local receipts and expenditures, 1929–41



Note: Data are annual.
Source: U.S. Bureau of Economic Analysis, *National Income and Product Accounts of the United States*, Historical Tables, tables 3.2 and 3.3.

FIGURE 5

Federal government revenues, by source, and expenditures, 1934–41



Note: Data are quarterly over the period 1934:Q1–1941:Q4.
Source: Board of Governors of the Federal Reserve System (1943), table 150, pp. 513–515.

On the revenue side, it is apparent that revenues increased sharply in the first quarter of 1937. There are two main factors. The most important one is the increase in income tax revenue, which grew by 66 percent from 1936 to 1937. This was due to a significant increase in income tax rates in the Revenue Act passed in June 1936. The rates previously ranged from 4 percent (starting at $4,000) to 59 percent (above $1 million). They remained unchanged for income brackets below $50,000, but were increased above that threshold, to reach 75 percent on the top earners. As a result, the average marginal tax rate for incomes above $4,000 almost doubled, from 6.4 percent to 11.6 percent.[3]

The second factor, of lesser quantitative importance, is the beginning of Social Security taxation. The Social Security tax rate was 2 percent, with half paid by the employer, and the ceiling was $3,000. Collection began in January 1937, and represented 10.5 percent of total federal tax receipts for the year 1937.

The undistributed profits tax

One interesting component of fiscal policy in that period was the introduction of a tax on undistributed profits (Lent, 1948). The motivation for the tax was not so much to raise revenue as to encourage firms to pay out dividends. The government saw this as desirable for two reasons. First, the accumulation of earnings by corporations allowed some earnings to avoid income taxation. Second, it was thought that firms did not know the best uses of the capital they were retaining and could possibly spend it on wasteful projects. According to this view, it would be better to send the earnings to the shareholder and flowing back into general capital markets.

The tax was announced, without warning, by President Roosevelt in March 1936, and enacted in the summer as part of the Revenue Act of 1936. Earnings that were not distributed as dividends were subjected to an additional tax. Lent (1948) found that the tax generated little revenue because most corporations, especially the large ones, simply paid out larger dividends. Also, smaller corporations were able to use legal mechanisms to require their shareholders to reinvest the dividends into shares of the corporation. The firms that were the most affected (as shown by the increase in their tax liability) were the medium-sized corporations.

The tax, although it had little effect in terms of revenues, could have had two effects on the economy. First, to the extent that small and medium firms find it difficult to access credit and capital markets, they have to rely on internal sources of funds to finance investment. The tax would obviously increase the cost of investment for those firms. Second, the tax was reflective of a changed political climate and increasingly populist rhetoric coming from politicians and the Roosevelt administration. At the same time as the tax was announced, the Roosevelt administration was becoming increasingly vocal against "economic royalists," alleged monopolists, and business in general. Although the tax was widely considered a failure and was repealed in all but name after two years, it may have played a psychological role in increasing uncertainty about the profitability of investment. This assessment must be tempered by the fact that, as table 1 (p. 18) shows, there was a surge in investment in the second half of 1936, and all components of investment do not start falling uniformly until late in 1937.

By early 1938, the severity of the recession prompted a turnaround in fiscal policy. This was manifested in a dramatic announcement by President Roosevelt on April 14, 1938, of a new "spend–lend" program with a $2 billion increase in spending.

To sum up, fiscal policy became tighter in early 1937, with a brief return to a balanced budget due to tax increases. The stance was reversed in early 1938, shortly before the trough of the recession.

Monetary policy

Most of the recent discussions of the 1937 recession have centered on the monetary policy carried out by the Federal Reserve System. Because the 1930s were a period of great change for monetary policy, I will first provide some background on this change to show that the Fed abandoned its traditional instruments and adopted a passive attitude during the first half of the 1930s. When policy became active again in 1935, it was through the use of a new instrument, namely, changes in reserve requirements, coupled with actions by the U.S. Department of the Treasury. The stance of monetary policy, like that of fiscal policy, reversed as the 1937 recession took its toll. I will then examine in more detail the response of the banking system to monetary policy during the recession.

Background

The 1930s were a period of considerable change for U.S. monetary policy. The turning point was the Gold Reserve Act, passed on January 30, 1934. It nationalized all gold in the United States, including the gold reserve held by the Fed. It authorized the president to devalue the dollar, which he did immediately, changing the dollar price of an ounce of gold from $20.67 (its price since the 1830s) to $35. This implied that the Treasury made a capital gain of 60 percent, or about $2 billion, on its newly acquired gold holdings. The proceeds were used to create an Exchange Stabilization Fund under the sole discretionary control of the Treasury. The existence of the fund gave the Treasury a strong hand in its dealings with the Fed, and for the next 17 years the Treasury dominated monetary policy.

From its foundation to the early 1930s, the Fed's balance sheet had consisted essentially of its gold reserve, which backed the currency (subject to a 40 percent reserve requirement) and private debt. Monetary policy consisted of managing the portfolio of private debt, either through discounting or, since the 1920s, through open-market purchases and sales of private debt. The debt was short-term, either commercial paper or bankers' acceptances, with typically 90 days or less to maturity.

Figure 6, panels A and B show the rates at which the Fed bought commercial paper and bankers' acceptances, compared with open-market rates. The Fed's rate in panel A is somewhat lower than the market rate because the latter pertains to paper of four to six months maturity, whereas the Fed purchased shorter maturities. Both panels in figure 6 show that, until 1932, the Fed's rate was close to the market rate; in other words, the Fed was active in the open market. After 1934, the Fed's rates are above market rates, indicating that the Fed had ceased to use interest rates for the conduct of monetary policy.

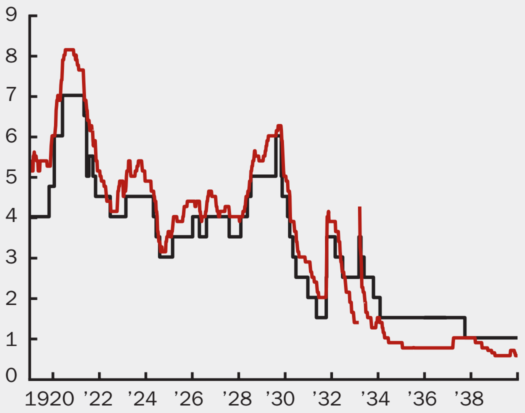
In the years that followed, the stance of monetary policy was dictated by actions of the Treasury. This can be seen in figure 7, which plots the sources of reserve funds—that is, the existing and potential sources of legal tender. Treasury currency (that is, currency issued directly by the Treasury) and Federal Reserve credit—the first two components—played no role in the 1930s, as they remained essentially constant. The Fed's portfolio during that period consisted of gold certificates (issued by the Treasury in 1934 in exchange for the Fed's gold reserve) and government bonds. Private debt had completely disappeared. The portfolio was kept constant throughout the period, with a few minor exceptions. The gold stock, the third component, was the main source of variation in the monetary base.

The Treasury did not immediately monetize the capital gain it had made on gold. The source of growth in the monetary base is to be found elsewhere. From 1934 on, persistent gold inflows into the United States account for the growth in the gold component. There were two reasons for the inflows. After the devaluation of 1934, foreigners bought dollars because they had become cheaper (and U.S. domestic prices had not adjusted fully). Later, gold inflows continued because increasing political instability in Europe induced long-term capital flows into the United States.

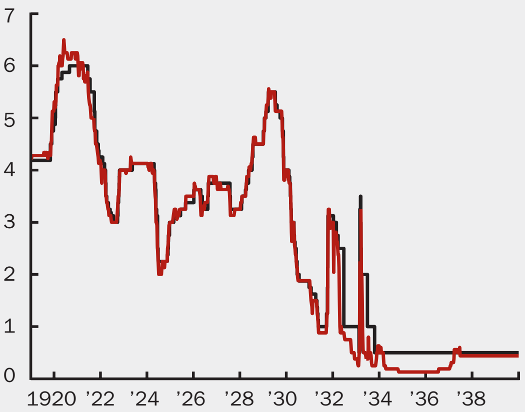
FIGURE 6

New York Reserve Bank (NYRB) rates and prevailing open-market rates, 1919–39

A. NYRB discount rate and open-market prevailing rate on 4–6 month prime commercial paper



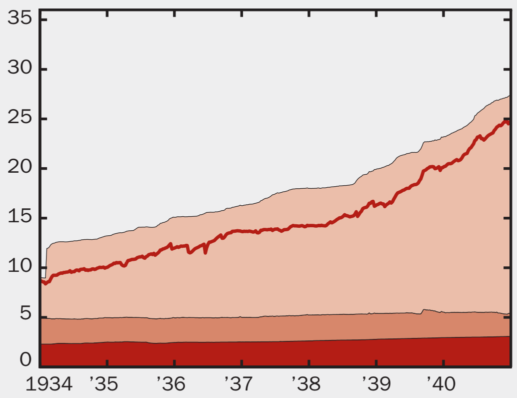
B. NYRB buying rate and open-market prevailing rate for 90-day bankers' acceptances



Note: Data are weekly over the period 1919–39.
Source: Board of Governors of the Federal Reserve System (1943), tables 115, 117, and 121, pp. 442–445, 452–459.

FIGURE 7

Monetary base and components of the supply of reserve funds, 1934–40



Note: Data are weekly over the period January 3, 1934–December 31, 1940.
Source: Board of Governors of the Federal Reserve System (1943), table 103, pp. 378–394.

When foreigners offered gold for sale, the Treasury issued gold certificates and deposited them at the Fed, increasing its account's balances. The Treasury then used the increase to pay for the gold. Thus, gold inflows translated one for one into increases in the monetary base. In other words, gold inflows were monetized. This accounts for the steady increase in the monetary base.

The growth in reserves

Then, as now, the U.S. banking system comprised a variety of banks depending on supervisory jurisdiction. Banks incorporated under federal law were all members of the Federal Reserve System and the Federal Deposit Insurance Corporation (FDIC). Banks incorporated under state law could be members of the Federal Reserve System and the FDIC, the FDIC only, or neither. Unincorporated banks could be members of the FDIC. In June 1936, member banks represented 70 percent of all bank deposits. In this section I focus on member banks' statistics, since they were more frequently collected by the Federal Reserve System and were directly affected by the System's changes in reserve requirements.[4]

Reserves (which can take the form of currency or balances at Federal Reserve Banks) were required by law since 1917. The quantity of required reserves depended on the total amount of demand deposits (net of deposits of other banks) or time deposits that a bank held, as well as its location (see table 2). Reserves above the required amount are excess reserves. If a bank located in a central reserve city held $1 in excess reserves, it could potentially increase its demand deposits by an additional $7.69.

For all member banks, reserves grew from $2,235 million in June 1933, near the trough of the

Great Depression, to $6,613 million by March 1937, on the eve of the 1937 recession, a 300 percent increase. Demand deposits during the same period grew only from $26,564 million to $41,114 million—a 150 percent increase. With constant reserve requirements, this meant that excess reserves grew considerably: In January 1934, they were estimated to be $827 million, but by March 1935, when the Fed began to be concerned, they had reached $2,200 million, or 48 percent of total reserves. By comparison, before the banking panics of 1931, excess reserves were typically 2 percent or 3 percent of total reserves (Board of Governors of the Federal Reserve System, 1943).

Why did banks hold such large reserves? Friedman and Schwartz (1963) propose a shift in banks' preferences for reserves as a consequence of the banking panics of 1931–33. This shift took place gradually over the period 1933–36, and subsided slowly only in the late 1930s and early 1940s as experience with the FDIC and general economic recovery made banks more comfortable holding lower levels of reserves.[5] In contrast, Frost (1971) has argued that banks' demand for reserves was stable throughout the period. With fixed costs of adjusting reserves, that demand for reserves behaves differently at low levels of interest rates than at higher levels. Below a certain threshold, the demand rises much more rapidly as rates fall. The reason is that, when short-term rates are low, it is less costly to hold large amounts of reserves than repeatedly to incur the fixed cost of adjusting. Frost's explanation for high reserves in the 1930s is solely the low level of interest rates.

Policy reaction (1935–37)

Beginning in March 1935, the Federal Open Market Committee (FOMC), which determines monetary policy and interest rates, became increasingly concerned with the growth in excess reserves. Its members feared that such reserves could ultimately lead to an uncontrolled credit expansion, once banks decided to increase their deposits. At that date, the Fed staff prepared a background memo titled "Excess reserves and Federal Reserve policy." The authors believed that increasing government debt supplied the bonds that led to reserve growth. But the memo concluded that neither past experience nor central bank theory gave any guidance for a policy response in the current circumstances (Meltzer, 2003, pp. 492–493).

Yet in spite of mounting concerns, it took the Fed over a year to take action. This was due partly to the uncertainty presented in the March 1935 memo and partly to the need to avoid antagonizing the Treasury. Concerns over potential inflation were balanced against concerns over the recovery and the federal government's desire for low interest rates when it was financing its debt.

By October 1935, excess reserves in the banking system exceeded the Fed's portfolio of government bonds, and the FOMC decided to analyze the distribution of excess reserves across banks to make sure that increases in requirements would not fall disproportionately on some banks. It also decided to coordinate policy with the Treasury. The ultimate result of this coordination was that the policy actions in 1936–37 took two forms: increases in reserve requirements by the Fed and sterilization of gold inflows by the Treasury (explained in more detail later).

Friedman and Schwartz (1963, p. 544) see monetary policy (that is, the increase in reserve requirements and, "no less important," the gold sterilization program) as "a factor that significantly intensified the severity of the decline and also probably caused it to occur earlier than otherwise."

Reserve requirements

The Banking Act of 1935, passed in August 1935, made important changes to the structure of the Federal Reserve.[6] One of the changes concerned reserve requirements. Since 1917, reserve requirements had been set in section 19 of the Federal Reserve Act at various levels depending on the location of the member bank.[7] The Board of Governors was now given the authority to change the reserve requirements "in order to prevent injurious credit expansion or contraction,"[8] but the requirements could be no lower than they had been since 1917 and no higher than twice those levels.

The purpose of increasing the reserve requirements was to pave the way for a return to the Fed's traditional policy tools, namely, rediscounting (buying privately issued debt at a discount to reflect the time to maturity) and open-market operations. The Fed thought that, as long as excess reserves were so large, it could have no effect on the banking sector's lending activities. Only if banks became borrowers again would the Fed be able to ease or tighten policy; until then, in the famous phrase of Marriner Eccles, Chairman of the Board of Governors, the Fed would be "pushing on a string" (Meltzer, 2003, p. 478).

Why were policymakers worried about inflation in 1936? The answer is twofold. First, there were objective signs of inflation. Wholesale prices, which had been stable in the early part of 1936, began to rise in late 1936 and early 1937. The annualized six-month change in wholesale prices rose steadily from 0.5 percent in August 1936 to 10.2 percent in March 1937. Retail prices as measured by the National Industrial Conference Board's (NICB) cost-of-living index did not rise as fast, but the 12-month change was nevertheless 5.4 percent by March 1937.[9]

TABLE 2

Member bank reserve requirements, 1917–41

Effective date	Percent of net demand deposits: Central reserve city	Reserve city	Country	Percent of time deposits: All
June 21, 1917	13	10	7	3
August 16, 1936	19.5	15	10.5	4.5
March 1, 1937	22.75	17.5	12.25	5.25
May 1, 1937	26	20	14	6
April 16, 1938	22.75	17.5	12	5

Source: Board of Governors of the Federal Reserve System (1943), table 107, p. 400.

The other answer is that some policymakers were still worried about repeating what they saw as the mistake of the 1920s. In that view, the Great Depression was partly a result of the speculative excesses of the 1920s, which the Fed had not done enough to prevent. Whether they saw incipient signs of a speculative boom developing (or whether they wanted to prevent such a boom from getting started in the first place), there was for some an inclination toward preemptive action. Although this view was perhaps not dominant in the FOMC, it nevertheless supported the move toward action in early 1937 and slowed the reversal of policy later on during the downturn.

That said, it should be emphasized that the Fed did not see the increase in reserve requirements as contractionary, and its public pronouncements insisted that the stance of policy had not changed. In the Fed's view, mopping up excess reserves through the increase in requirements should have had no effect. Recent authors such as Currie (1980), Calomiris and Wheelock (1998), and Telser (2001) have argued that the increase in reserve requirements did not cause the recession.

Table 2 shows the changes in reserve requirements. The first increase in reserve requirements was announced on July 14, 1936, and went into effect a month later. At the time, total reserves stood at $5.87 billion, split almost exactly between required reserves of $2.95 billion and excess reserves of $2.92 billion. Thus, an increase of 50 percent in reserve requirements could be easily met by banks with the excess reserves.

The second and third increases were announced on January 30, 1937: The first was to take effect on March 1; the second on May 1. At the time of the announcement, total reserves had increased to $6.78 billion, and required reserves were $4.62 billion, slightly higher than they had been after the first increase. The increase of 33 percent in requirements could again be met from excess reserves, leaving over $600 million in excess reserves.

Figure 8, panel A shows total and estimated required reserves at weekly reporting member banks. The four vertical dotted lines on this panel and on panels B and C mark the four changes in reserve requirements (three increases and one decrease).

Two things are apparent from figure 8. One is that, in the aggregate, the increase in reserve requirements did not reduce excess reserves to zero (see panel B): The estimated excess reserves on May 5, 1937, the first reporting date after the last increase, were $887 million—28 percent of what they were on August 12, 1936, before the increases began.

The second point to make is that the growth of total reserves paused during 1937, and then resumed, mirroring the behavior of the monetary base (see figure 8, panel A). The two lines in the graph thus summarize the two prongs of monetary policy: The lower line (required reserves) reflects the Fed's actions, while the upper line (total reserves) reflects Treasury's sterilization of gold inflows.

Gold sterilization

As explained previously, since 1934 the Treasury had let gold inflows increase the monetary base. Starting in December 1936, the Treasury changed its procedure. Instead of, in effect, converting gold inflows into the monetary base, it used proceeds from bond sales to pay for the gold that was brought to the Treasury at the price of $35 per ounce. As a result, the gold stock in the United States continued to increase but the monetary base remained roughly constant (see figure 9, p. 26). From December 1936 to February 1938, the gold stock increased 15 percent, but the monetary base grew by only 4 percent. The policy was halted in February 1938 and reversed over the ensuing months.

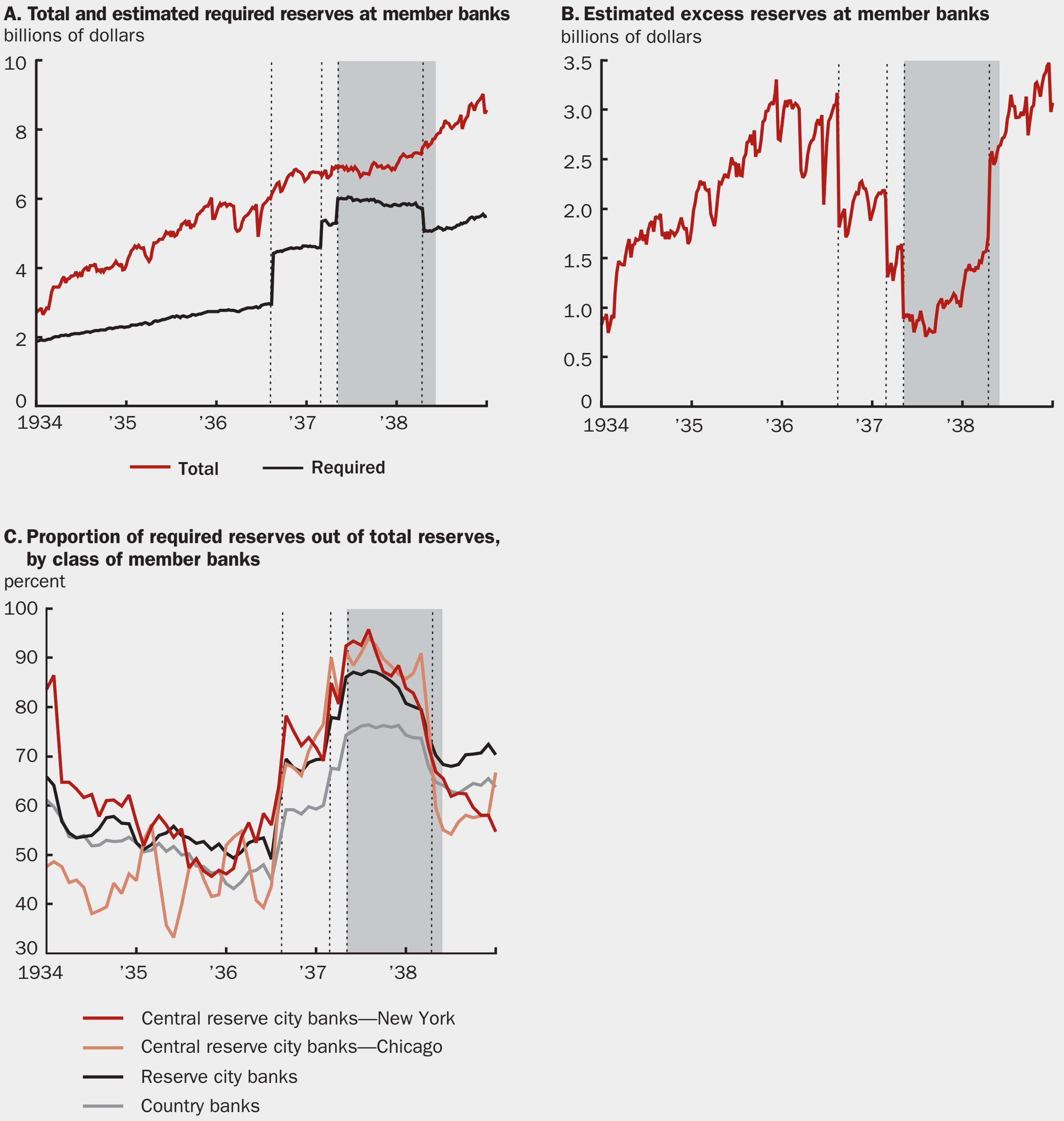
The response of banks

How did banks respond to the increase in reserve requirements? Figure 8, panel B plots excess reserves. It is apparent that, in the aggregate, excess reserves were sufficient to meet the new requirements, and panel B shows no sign of banks scrambling to keep their excess reserves at high levels, contrary to what is occasionally asserted.

The picture is somewhat different, however, if one looks at more disaggregated data. Figure 8, panel C shows the proportion of required reserves out of total reserves by class of member banks. Recall that member banks were classified according to their location. Banks in

FIGURE 8

Reserves of member banks, 1934–38



Notes: Data for panels A and B are weekly over the period January 3, 1934–December 28, 1938; data for panel C are monthly over the period January 1934–December 1938. The four vertical dotted lines in each panel mark the four changes in reserve requirements (three increases and one decrease) noted in table 2. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Source: Author’s calculations based on data from the Board of Governors of the Federal Reserve System (1943), tables 103 and 105, pp. 378–394, 396–398.

New York City and Chicago (the two central reserve cities) were considerably closer to their limit than banks in reserve cities and country banks.

A comparison of table 3 and table 4 confirms that the reaction of member banks in New York City was markedly different from that of the banking system overall. From June 1936 to June 1937, total deposits grew by 2.3 percent overall but only 0.3 percent in New York City member banks. Loans increased by 8.6 percent overall, and by 21.2 percent among New York City banks; but the latter banks reduced their holdings of government bonds by 23.8 percent and other securities

TABLE 3

All banks: Main assets and total deposits, 1936–38

Call dates	Loans	Investments: Total	Investments: U.S. government bonds	Investments: Other securities	Deposits
	(millions of dollars)				
June 1936	20,636	27,776	17,323	10,453	57,884
December 1936	21,359	28,086	17,587	10,499	60,619
June 1937	22,410	27,155	16,954	10,201	59,222
December 1937	22,065	26,362	16,610	9,752	58,494
June 1938	20,982	26,230	16,727	9,503	58,792
December 1938	21,261	27,570	17,953	9,617	61,319

Source: Board of Governors of the Federal Reserve System (1943), table 2.

TABLE 4

New York City member banks: Main assets and total deposits, 1936–38

Call dates	Loans	Investments: Total	Investments: U.S. government bonds	Investments: Other securities	Reserves	Deposits
	(millions of dollars)					
June 30, 1936	3,528	6,028	4,763	1,265	2,106	11,387
December 31, 1936	3,855	5,426	4,209	1,217	2,658	11,824
March 31, 1937	3,961	5,140	3,829	1,311	2,719	11,400
June 30, 1937	4,276	4,730	3,630	1,100	2,749	11,421
December 31, 1937	3,673	4,640	3,595	1,045	2,738	10,759
March 7, 1938	3,532	4,785	3,611	1,174	2,941	10,570
June 30, 1938	3,172	4,841	3,740	1,101	3,517	11,192
September 28, 1938	3,146	5,209	3,987	1,222	3,743	11,410
December 31, 1938	3,262	5,072	3,857	1,215	4,104	11,706

Source: Board of Governors of the Federal Reserve System (1943), table 23.

by 13.0 percent, while banks overall reduced these holdings 2.1 percent and 2.4 percent, respectively.

Although banks did not suddenly increase their reserve holdings, they did reduce the rate of growth of deposits. Figure 9 shows that demand deposits, which had been growing steadily at 20 percent per year since March 1933, grew more slowly starting in July 1936 and peaked in March 1937. Over the next 12 months, they fell by 6 percent, reached a low in December 1937, and began growing again after the end of the recession.

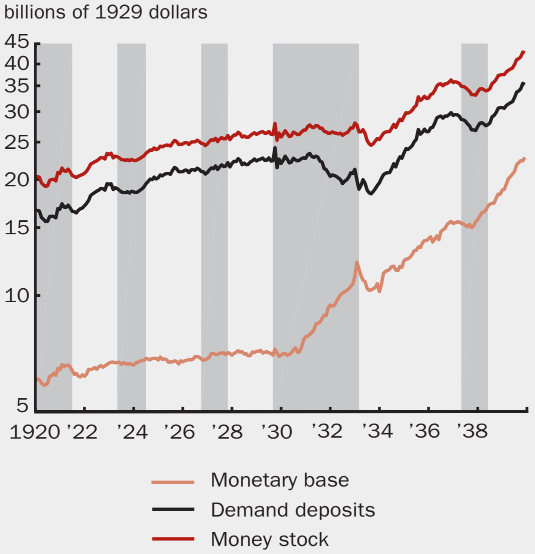
This describes the size of the banking sector from the liability side. Which assets shrank to meet this fall in liabilities? Figure 10 shows that member bank assets fell into three broad categories: reserves, investments (U.S. bonds and other securities), and loans. Reserves grew, as we saw. Loans were not affected much, although they grew more slowly than before. Among reporting member banks, the 12-month growth rate of loans peaks on August 4, 1937, at 19.1 percent, and the absolute level peaks on September 15, 1937, at $10.05 billion, 16 percent higher than a year before. Loans then fall 4.5 percent in the next three months as the recession deepens. The category that bore the brunt of the reduction was investments, particularly government debt, simply because those were the most liquid assets. This can be seen for weekly reporting member banks in table 5, which shows the composition of assets for the week following each change in reserve requirements. From the second to the third increase (March–May 1937), total assets fell by $1.3 billion: Loans actually increased by $0.5 billion, and most of the reduction came from U.S. bonds.

Looking at interest rates confirms that the impact of reserve requirements manifested itself on tradable securities rather than loans. Table 6 shows that rates charged by banks on loans were little affected (and in some locations fell) after the reserve requirements increased, while short-term commercial paper rates rose.

The impact of the second round of reserve requirement increases was felt immediately in the U.S. bond market. There is in fact a particular day, March 15, when U.S. long-term bonds, whose yields had remained very stable, went up by 2 basis points, prompting the Secretary of the Treasury to get on the phone and

FIGURE 9

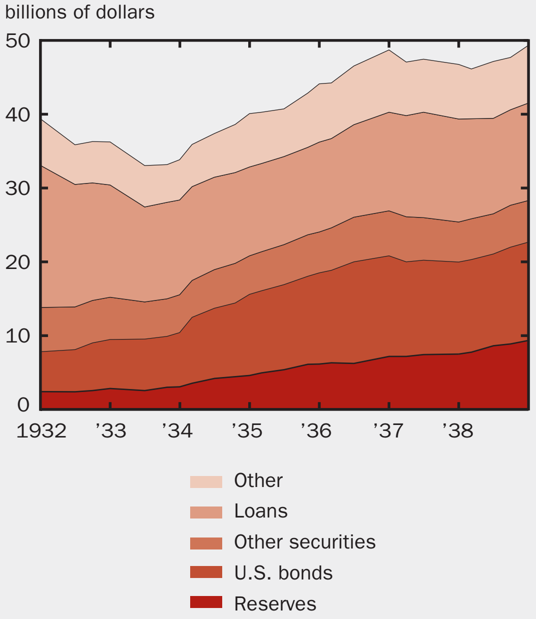
Monetary base, demand deposits at commercial banks, and money stock, 1919–39



Notes: Data are monthly over the period December 1919–December 1939 and adjusted for cost of living. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.
Source: Friedman and Schwartz (1963), appendix A.

FIGURE 10

Components of member banks' balance sheets, 1932–38



Note: Data are quarterly over the period 1932–38.
Source: Board of Governors of the Federal Reserve System (1943), table 18.

complain to Eccles, the Chairman of the Fed, that the Fed had bungled the increase in reserve requirements.

Figure 11, panels A and B show clearly how bond rates increased in March 1937, before the recession began. Figure 12, which shows that corporate issues declined sharply in March 1937, suggests a channel through which the increase in reserve requirements could have affected the economy, namely, by reducing the banking sector's demand for (government and) corporate liabilities. The fall in lending translated into higher interest rates and a lower volume of issues.

Prices

The behavior of prices (see figure 13) during the recession is broadly consistent with the notion that monetary policy was contractionary. Whichever indicator one uses, it is apparent that prices peaked in mid-1937, as the recession was under way (recall that there is some ambiguity as to the exact starting date, either May or August). The aggregate indexes normally used (the deflators for gross domestic product and personal consumption expenditures) are only available annually for this period. The monthly indexes such as the National Industrial Conference Board's cost-of-living index and the Wholesale Price Index, closely watched at the time for evidence of inflation, both peaked in September 1937; the Consumer Price Index did too, although it is not a very good measure for this period because data were not collected on a monthly basis, and missing data are interpolated.

What is rather puzzling is that the trend in prices, having turned deflationary during the recession, continued well after the end of the recession. The National Industrial Conference Board's index bottoms out in June 1939, having declined by 2.2 percent since the end of the recession, while wholesale prices, having fallen 3 percent, bottom out in August 1939 just before the outbreak of the European war sets off speculative buying.

Cole and Ohanian (1999) have used the recession of 1920–21, during which output fell sharply after a steep drop in prices and recovered strongly once the deflation ended, to highlight the puzzling nature of the slow recovery after the end of deflation in 1933. The recovery of 1938 adds to this puzzle: The economy rebounded as sharply as in 1921—an analogy noted by Friedman and Schwartz (1963) in spite of continued deflation (Steindl, 2007).

The following conclusions emerge from the foregoing discussion. First, monetary policy was as much

TABLE 5

Assets of weekly reporting member banks after reserve requirement changes, 1936–38

	Loans	Government bonds	Other securities	Reserves	Vault cash	Balances with domestic banks	Total assets
	millions of dollars						
August 19, 1936	8,369	10,564	3,323	4,884	373	2,288	32,315
March 3, 1937	9,054	10,303	3,318	5,291	398	2,055	33,677
May 5, 1937	9,533	9,499	3,208	5,307	337	1,797	32,362
April 20, 1938	8,585	9,156	3,068	5,980	330	2,188	31,938

Source: Board of Governors of the Federal Reserve System (1943), table 48.

TABLE 6

Short-term rates and lending rates, 1936–38

	1936		1937		1938	
	June	December	June	December	June	December
	percent					
Short-term open-market rates in New York City						
Prime commercial paper, 4–6 months	0.75	0.75	1.00	1.00	0.88	0.63
Prime bankers' acceptances, 90 days	0.13	0.19	0.47	0.44	0.44	0.44
Rates charged on customers' loans by banks in principal cities						
Total (19 cities)	2.71	2.58	2.57	2.52	2.56	2.60
New York City	1.71	1.74	1.73	1.70	1.70	1.70
North and East (7 cities)	3.02	2.94	2.79	2.72	2.70	2.95
South and West (11 cities)	3.51	3.14	3.29	3.23	3.31	3.31

Source: Board of Governors of the Federal Reserve System (1943), tables 2, 120, and 124.

a consequence of Fed actions as Treasury actions. The fall in the money stock was a direct consequence of the gold sterilization program. Second, the increases in reserve requirements began to bite only in early 1937. Finally, the channel through which monetary policy affected the banking system is not as straightforward as sometimes asserted. Lending did not begin to fall until the recession was under way. The reaction of the banking system was to liquidate securities holdings, primarily U.S. bonds, but also private sector securities. The impact on market rates is evident starting in March 1937.

Labor costs

Roose (1954) and other authors have cited alternative explanations to the monetary and fiscal policy stories. A number of these stories center on increased labor costs, which I now take up.

New Deal policies and the slow recovery

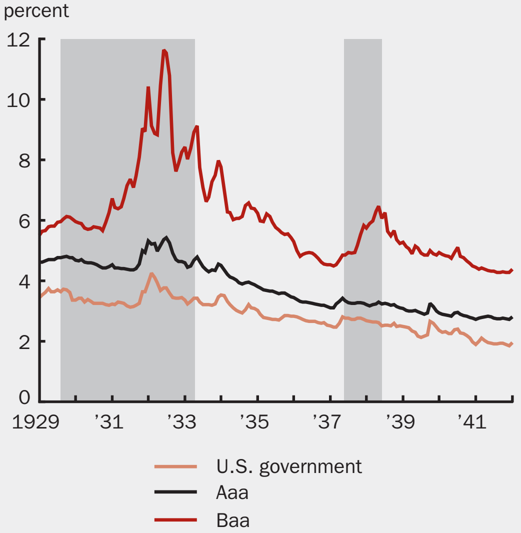
As Friedman and Schwartz (1963, p. 493) put it, "the most notable feature of the revival after 1933 was not its rapidity but its incompleteness": Unemployment remained high throughout the 1930s, the revival was erratic and uneven, and private investment (particularly construction) remained very low compared with the 1920s. They go on to note that prices rose much more than in earlier expansions despite the large resource gaps that remained. The reason was "almost surely the explicit measures to raise prices and wages undertaken with government encouragement and assistance. ... In the absence of the wage and price push, the period 1933–37 would have been characterized by a smaller rise in prices and a larger rise in output than actually occurred" (Friedman and Schwartz, 1963, pp. 498–499).

Cole and Ohanian (2004) develop a general equilibrium model to address this question. Specifically, they document that output, consumption, investment,

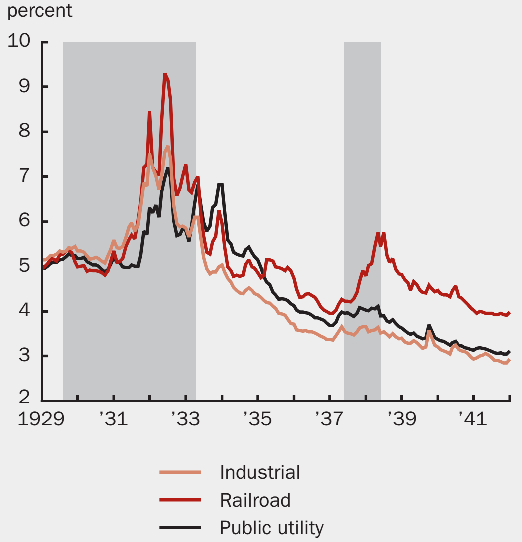
FIGURE 11

Rates on bonds, 1929–41

A. Rates on government bonds and corporate bonds, by rating



B. Rates on corporate bonds, by sector

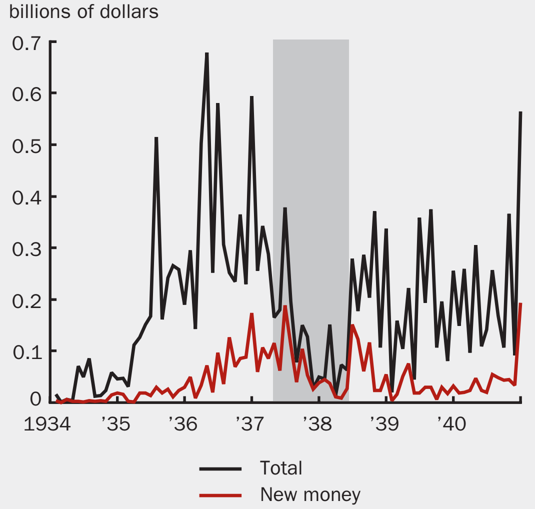


Notes: Data are monthly over the period January 1929–December 1941. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Source: Board of Governors of the Federal Reserve System (1943), table 128, pp. 468–471.

FIGURE 12

New corporate security issues, total and proceeds proposed as new money, 1934–40



Notes: Data are monthly over the period January 1934–December 1940. The shaded area indicates an official period of recession as identified by the National Bureau of Economic Research.

Source: Board of Governors of the Federal Reserve System (1943), table 138, pp. 491–492.

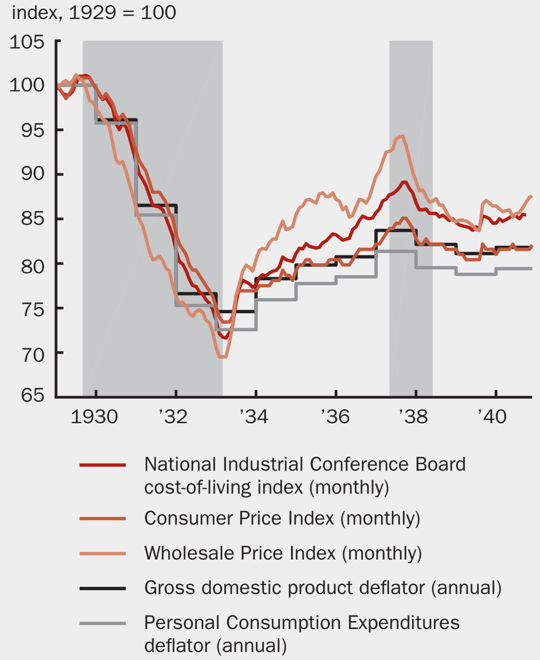
and hours worked remained far below trend from 1934 to 1939, while total factor productivity was back to trend in 1936 and wages were 10 percent to 20 percent above trend. They also document the history of the National Industrial Recovery Act (NIRA), which gave industries protection from antitrust legislation as long as firms raised prices and shared their profits with workers in the form of higher wages. The NIRA was struck down by the Supreme Court in May 1935, but the National Labor Relations Act (NLRA), or Wagner Act, was passed in July 1935 to increase the bargaining power of unions; furthermore, according to Cole and Ohanian (2004), the Roosevelt administration did not enforce antitrust laws, allowing firms to continue to collude in raising prices. The combined effect of firm collusion and worker bargaining power is shown in their equilibrium model to result in lower output than in a competitive version of the same economy. Starting from 1934 conditions and assuming the same changes in total factor productivity (that is, changes not due to changes in inputs) as in the data, the competitive version of the economy returns to trend by 1936 or so, while the cartelized version of the economy, calibrated so that wages are 20 percent above their normal level, displays lower consumption, investment, and employment, along with higher wages.

New Deal policies and the 1937 recession

The Wagner Act, which is still in force today, immediately generated legal challenges. Several cases involving that act were taken up by the Supreme Court in January 1937.

FIGURE 13

Various measures of prices, 1929–41

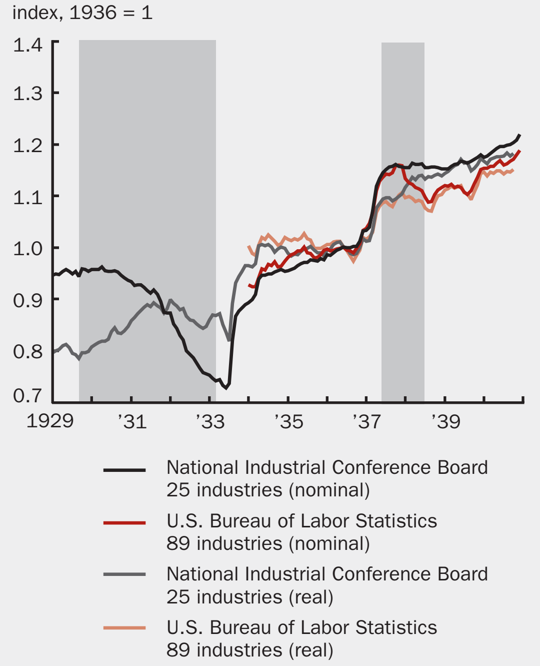


Note: The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Sources: Beney (1936); National Industrial Conference Board (1938, 1939, 1940); U.S. Bureau of Labor Statistics from National Bureau of Economic Research, Macrohistory Database, series m04169a; and U.S. Bureau of Labor Statistics and U.S. Bureau of Economic Analysis from Haver Analytics.

FIGURE 14

Indexes of nominal and real wages, 1929–40



Notes: Data are monthly over the period January 1929–December 1940. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Source: U.S. Bureau of Economic Analysis, *Survey of Current Business*, 1934–41, various issues.

In an earlier version of their work, Cole and Ohanian (2001, p. 49) commented that “while more work is required to assess the 1937–38 downturn, our theory raises the possibility that an increase in labor bargaining power may have been an important contributing factor to the downturn of 1937–38.” In recent testimony to the Senate, Ohanian (2009) went further to assert: “Wages jumped in many industries shortly after the NLRA was upheld by the Supreme Court in 1937, and our research shows that these higher wages played a significant role in the 1937–38 economic contraction.”

The behavior of wages

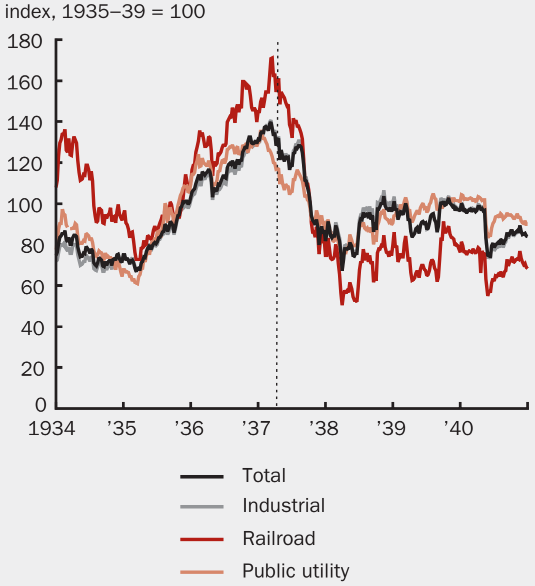
Figure 14 shows the behavior of wages, as measured by the U.S. Bureau of Labor Statistics (BLS) and the National Industrial Conference Board, both in nominal terms and deflated by the monthly cost-of-living index computed by the National Industrial Conference Board.

Nominal average hourly earnings as measured by the BLS had been close to constant from January to August 1936, oscillating between \$0.571 and \$0.575. They fell to \$0.569 in September and then began to rise. By April 1937, they had reached \$0.638; they peaked at \$0.667 in November of that year, 17 percent higher than in September 1936. But most of that increase (70 percent) had taken place before the Supreme Court handed down its decision on April 12, 1937. The picture is not much different if we look at the National Industrial Conference Board’s index: According to this measure, 67 percent of the rise had occurred before the decision.

Stock prices (shown in figure 15) do not support the notion that the release of the Supreme Court decision suddenly shifted bargaining power within existing collusive arrangements from firms to workers. If that had been the case, then the net present value of the firms’ share of the collusive rents should have fallen upon receipt of the news. In fact, April 12, 1937, was a quiet day on stock markets, and the next day the *Wall Street Journal* commented that the decision “caused little more than a ripple in the markets,” an initial sell-off in steels having been recovered before day’s end, and trading before and after the decision “hardly

FIGURE 15

Stock prices, 1934–40

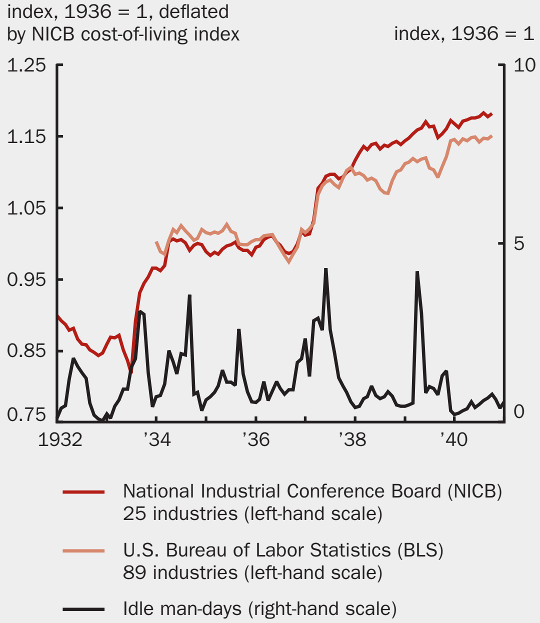


Notes: Data are weekly over the period January 3, 1934–December 25, 1940. The dashed vertical line marks April 13, 1937, the day after the Supreme Court decision on the Wagner Act was released.

Source: Board of Governors of the Federal Reserve System (1943), table 134.

FIGURE 16

Nominal wages and man-days lost to strikes, 1932–40



Notes: Data are monthly over the period January 1932–December 1940 (the BLS series starts in January 1934). Factory hourly average earnings are measured on the left-hand scale, and idle man-days due to strikes are measured on the right-hand scale.

Sources: National Bureau of Economic Research, Macrohistory Database, series m08257; and U.S. Bureau of Economic Analysis, *Survey of Current Business*, 1934–41, various issues.

better than dull.” The paper went on to assert that “had the decisions ... gone the other way, there is little doubt that the share market would have responded with a show of active buying; but as the reverse was true, the stock market’s following managed to be philosophic about it” (Dow Jones and Company, 1937b, p. 37).

The commentary, as well as a quotation by Henry Ford the following day that “we thought the Wagner Act was the law right along” (Dow Jones and Company, 1937a, p. 2), suggests that the decision had been correctly anticipated all along. Surprisingly, however, the stock market had been rising steadily over the previous two years. From May 1935, when the NIRA was struck down, to April 1937, it shot up 70 percent (see figure 15).

Why did wages rise in 1936–37?

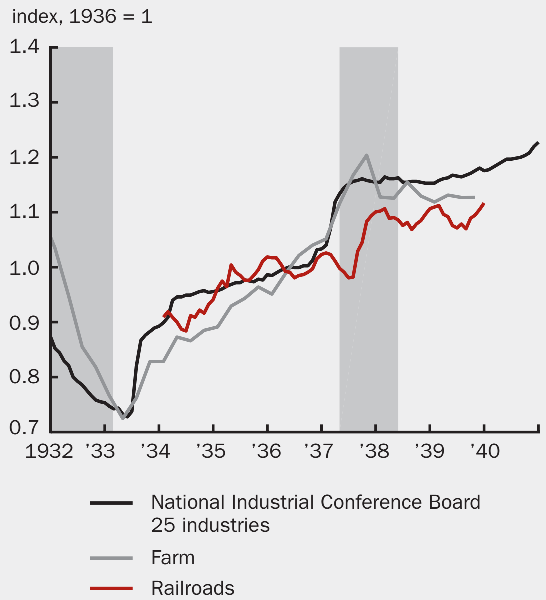
There is little doubt that the rise in wages was linked to the Wagner Act and the resulting increase in labor union activity. Figure 16 compares the level of wages with the number of man-days lost to strikes. While strikes were recurrent throughout the period, a sustained increase in strikes is noticeable from the end of 1936 to a peak in June 1937 of 5 million man-days—four times the decade’s average.

To test the claim that the Wagner Act caused wage increases in manufacturing, I looked at wages in railroads and farming, sectors to which the act did not apply. Figure 17 shows data collected by the National Industrial Conference Board for wage rates in 25 industries, all wage earners in Class I railroads, and farm labor. Wages did not rise for railroad employees when they were rising in industry: In fact, railroad wages fell during the same period. However, farm wages follow the same pattern as industrial wages throughout the whole period.

I have looked at industry data to see if industries that saw a greater increase in wages also saw a larger drop in employment. To do this, I regressed the percentage change in employment from July 1937 to June 1938 (the peak and trough of total employment) on the percentage change in average hourly earnings from September 1936 to November 1937 (the trough and peak of nominal wages in figure 14, p. 29). The results for the 31 industries are shown in figure 18. The relationship is negative and significantly different from 0

FIGURE 17

Average hourly earnings in 25 industries and in railroads and wage rates of farm labor, 1932–40

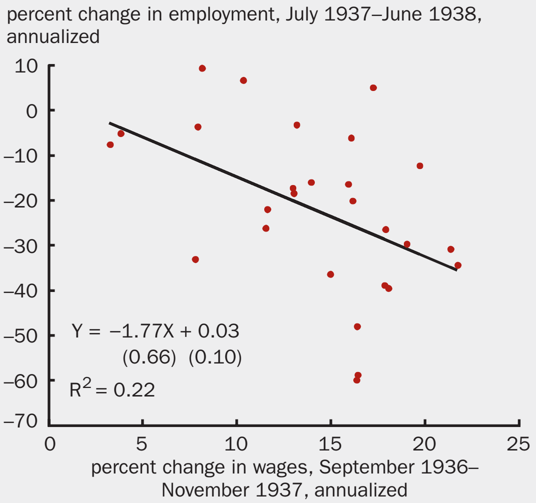


Notes: The average hourly earnings data for the 25 industries and the railroads are monthly and the monthly wage rates of farm labor data are quarterly, seasonally adjusted, over the period 1932–40. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research.

Sources: National Industrial Conference Board (1938, 1939, 1940).

FIGURE 18

Impact of wage increases on employment across 31 industries



Source: Author's calculations based on data from the U.S. Bureau of Economic Analysis, *Survey of Current Business*, 1936–38, various issues.

at the 1 percent confidence level. The coefficient is large: A 1 percent increase in wages leads to a 1.8 percent fall in employment, but it is imprecisely estimated.

More problematic is the fact that these results are not robust to slight changes in the dates at which the changes are measured. For example, just changing the end date for wage increases from November 1937 to June 1937 reduces the R^2 statistic from 22 percent to 7 percent; and the estimated coefficient is three times smaller and not significantly different from 0 at the 5 percent confidence level.

Quantitative assessment

I have described the main explanations proposed for the recession of 1937. To assess which one (or more) of these explanations is the most plausible, it is necessary to go beyond theoretical plausibility and pure issues of timing. Ideally, one would construct a well-specified economic model that encompasses the competing explanations, estimate the parameters of the model using actual data, and then carry out experiments in the model. This is a difficult task, if only because there is not an agreed-upon model to use.

It is nevertheless possible to make a quantitative assessment, using the techniques of vector autoregression (VAR) analysis. The basic idea behind VAR analysis is to construct a statistical model of the relations between any number of variables of interest. The variables are all interrelated, both over time (one variable's current value affects another variable's future value) and within a single time period. That is because, typically, all of the variables are determined simultaneously by the economy. Prices do not explain quantities any more than quantities explain prices: Both are determined jointly by supply and demand relationships. In a dynamic setting, where variables evolve over time, an additional complication is that the future may influence the past. Suppose that it is known with certainty that a certain event will take place a year from now; individuals will plan ahead accordingly and alter their decisions today. The future is embedded in the present to the extent that it is anticipated. Only surprises may reveal to us what the effect of a particular variable might be.

VAR analysis is in some ways a generalization of standard regression analysis but acknowledges that the left-hand-side variable in one relation is the right-hand-side variable in another. Therefore all regressions are computed simultaneously. Each regression has a residual, which represents the effect of unexpected or unexplained variations. For example, we may specify that output is a function of past values of money growth,

TABLE 7

Variance–covariance matrix of the residuals of the vector autoregression

	M1	Surplus	AHE	CP rate	WPI	IP
M1	1.000					
Surplus	−0.026	1.000				
AHE	−0.203	0.087	1.000			
CP rate	−0.454	−0.041	0.041	1.000		
WPI	0.070	−0.064	−0.405	0.059	1.000	
IP	0.355	−0.032	−0.475	−0.163	0.327	1.000

Note: M1 is the money stock; surplus is the federal surplus; AHE is an index of real average hourly earnings in industries, adjusted for changes in output per man-hour; CP rate is the interest rate on commercial paper; WPI is the Wholesale Price Index; and IP is industrial production.
Sources: Author's calculations based on data from Friedman and Schwartz (1963); Board of Governors of the Federal Reserve System (1943), tables 115, 117, 121, and 150; Board of Governors of the Federal Reserve System, G.17 statistical release, various issues; and National Bureau of Economic Research, Macrohistory Database, series 08142 and 01300.

fiscal surplus, wages, and other variables of interest. In each time period, we will have the predicted value of output based on the past histories of these variables, and the actual value will differ to some extent: That is the error term, or innovation. Statistical theory tells us that a system of variables can be represented as the sum of the responses to current and past innovations. If the innovations are properly identified, it becomes possible to say, for example, that output responds in a certain way to an unexpected change in money growth or fiscal policy.

The problem is to identify the innovations to each variable. If we allow that, say, monetary policy can be affected within the current period by fiscal policy, then the error term in the money equation will combine innovations to money as well as innovations to fiscal policy. In that case, output responses to this innovation will be responses to *both* monetary policy and fiscal policy, and statistics are of no use in disentangling the two. In general, one must make identifying assumptions guided by economic theory to interpret a VAR.

The VAR I run includes a small set of variables to describe the state of the economy: industrial production (IP), the rate on commercial paper to measure short-term interest rates, and the Wholesale Price Index (WPI) to represent prices (Sims, 1980; and Burbidge and Harrison, 1985). Furthermore, it includes one variable for each competing explanation: the money stock (M1), the federal surplus, and an index of real average hourly earnings in industries (AHE), adjusted for changes in output per man-hour.[10] The data are monthly and extend from January 1919 to December 1941.[11]

The main identifying assumption is that money does not respond within the month to other contemporary variables. Thus, innovations to the regression of money growth on other variables represent only innovations to money growth itself. This is a common identification assumption for post-World War II data (Christiano, Eichenbaum, and Evans, 1999). I also assume that surpluses do not react within the month to anything but money growth, and that the wage does not respond to anything but money growth and the fiscal variable.

As it turns out, this particular ordering matters little. Table 7 shows the variance–covariance matrix of the residuals from the VAR. Note how it is close to diagonal in the first three elements, which means that the residuals of the monetary, fiscal, and wage equations are uncorrelated. Therefore, the residuals of the regressions can be interpreted as fundamental innovations, that is, exogenous and unpredictable changes in the monetary, fiscal, and wage conditions.[12]

Table 8 shows that the percentage of forecast error at various horizons is attributable to the innovations in each of the variables. The three factors together explain about half of the unpredictable movements in industrial production, but wages alone explain little. The other half is attributable to innovations to the other variables (interest rate, prices, and industrial production itself) to which I do not try to attach any particular meaning.

To understand how much each factor (monetary policy, fiscal policy, and wages) contributed to the recession of 1937, I examine a historical decomposition. The method is as follows. Using the estimated statistical relationships between the variables and data up through December 1935 only, I predict what industrial production will be in all succeeding months. Then, for each of the three factors, I compute the effect of its innovations on all the variables in each of the succeeding months from January 1936 to the end of the sample in December 1941. An innovation to money growth affects future money growth, but also all other variables. The effect of the sequence of innovations that I have computed from January 1936

TABLE 8

Variance decomposition of industrial production at various horizons

Months	Standard error	M1	Surplus	AHE	CP rate	WPI	IP
		(-------	-------	percent	-------	-------	-------)
6	0.097	15.5	4.8	16.1	4.3	6.8	52.6
12	0.136	30.3	7.9	11.2	2.5	5.0	43.0
18	0.165	39.1	8.5	7.7	2.7	3.5	38.4
24	0.187	41.2	8.3	6.9	3.2	2.8	37.7
30	0.203	40.4	8.0	6.8	3.3	2.5	38.9
36	0.216	39.2	8.1	6.7	3.4	2.2	40.4

Note: M1 is the money stock; surplus is the federal surplus; AHE is an index of real average hourly earnings in industries, adjusted for changes in output per man-hour; CP rate is the interest rate on commercial paper; WPI is the Wholesale Price Index; and IP is industrial production.
Sources: Author's calculations based on data from Friedman and Schwartz (1963); Board of Governors of the Federal Reserve System (1943), tables 115, 117, 121, and 150; Board of Governors of the Federal Reserve System, G.17 statistical release, various issues; and National Bureau of Economic Research, Macrohistory Database, series 08142 and 01300.

onward can be traced out for each factor and added to the baseline projection of industrial production, representing how much that factor explains.

The results are shown in figure 19. The black line represents the actual path of industrial output over the period. The baseline represents the forecast for industrial output based on information up through December 1935. This forecast essentially sees industrial output growing at a 4.5 percent trend. To the baseline I add successively the effect of innovations to the money stock (M1), to the surplus, and to wages.

The figure supports the following conclusions. First, the effect of changes in wages starting in early 1937 was to depress output, but by a small amount.[13] Moreover, the effect remains negative until late 1941, which is not surprising, since, as we saw, wages remained relatively high. Second, fiscal shocks and monetary shocks between them do a good job of accounting for the recession, with the following nuances. If monetary and fiscal shocks had been the only forces at play, the economy should have peaked in late 1936. Also, the monetary shock alone does not explain the full depth of the recession; the fiscal and monetary shocks explain the economy's turning point in mid-1938, but not the full extent of the recovery.

This indicates that other forces at work in the economy sustained the expansion from late 1936 to mid-1937, in spite of the contractionary impact of monetary policy and fiscal policy. Likewise, other forces contributed to the recovery, even as monetary policy and fiscal policy turned expansionary in mid-1938.

The results from the VAR need to be interpreted with caution. In particular, they do not disprove the importance of wages. In the Cole and Ohanian (2004) story, the change in workers' bargaining power is not a temporary shock to an otherwise stationary system, but a shift from one steady-state equilibrium to another. The VAR is not designed to capture such changes. The results do suggest, however, that as a quantitative matter the monetary and fiscal shocks are sufficient to account for the general pattern of the recession. Furthermore, the extent to which these factors fail to reproduce the data is by predicting an earlier and more prolonged downturn; in other words, the factor that is missing is an expansionary one, not a contractionary one like wages.

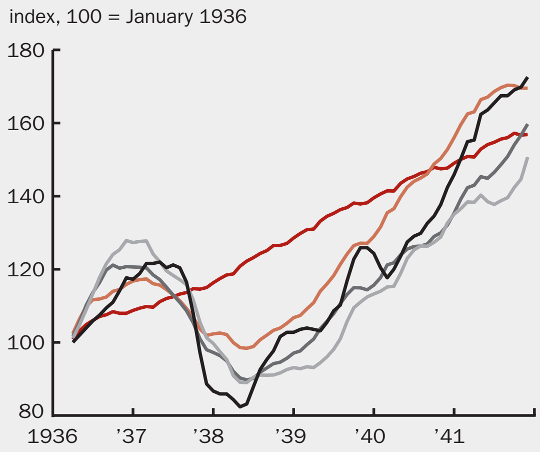
Conclusion

The recession of 1937 has been cited as a cautionary tale about the dangers of premature policy tightening on the way out of a deep downturn. In contrast, some authors have downplayed the role of monetary policy suggested by Friedman and Schwartz (1963). In particular, Cole and Ohanian (1999) dismiss the role of reserve requirements in the 1937 recession for two reasons. One is timing: "we would expect to see output fall shortly after" the changes in reserve requirements; but, they write, industrial production peaked in August 1937, 12 months after the first change (Cole and Ohanian, 1999, p. 10). The other is that interest rates did not increase: Commercial loan rates remained in the same range, and rates on corporate bonds "were roughly unchanged between 1936 and 1938" (Cole and Ohanian, 1999, p. 10).

I have shown in this article that tightened monetary policy consisted in the joint action of increased reserve requirements that were staggered from August 1936 to May 1937 and gold sterilization that started in December 1936. Gold sterilization turned the growth rate of money negative, and banks responded to increased reserve requirements by curtailing the financing of firms, with visible effects on interest rates. Industrial production peaked only a few months later, in May 1937.

FIGURE 19

Historical decomposition of industrial output, 1936–41



Actual
Baseline
Baseline + M1
Baseline + M1 + surplus
Baseline + M1 + surplus + wages

Notes: Data are monthly over the period January 1936–December 1941. M1 is the money stock; surplus is the federal surplus.
Sources: Author’s calculations based on data from Friedman and Schwartz (1963); Board of Governors of the Federal Reserve System (1943), tables 115, 117, 121, and 150; Board of Governors of the Federal Reserve System, G.17 statistical release, various issues; and National Bureau of Economic Research, Macrohistory Database, series 08142 and 01300.

Moreover, monetary policy went into reverse: The New York Fed lowered its discount rate from 1.5 percent to 1 percent on September 27, 1937; the gold sterilization program ended in February 1938 and was reversed from February to April; and reserve requirements were lowered on April 16, 1938, just as the federal government announced a large increase in spending. The recession ended in June 1938.

An alternative story would rely on increased labor costs due to the effects of the Wagner Act. The act was passed in 1935, but labor activism built up over the ensuing two years and resulted in 10 percent wage increases over a short period in early 1937. Although it has no particular timing advantage over the monetary and fiscal policy explanations, the labor costs story would plausibly account for the onset of the recession but not for the recovery, since the wage increases were not reversed.

Finally, a simple VAR shows that monetary and fiscal factors account fairly well for the pattern of industrial production and, in particular, for the depth of the recession, although other factors are needed to explain why the economy did not contract earlier and why it rebounded so strongly. Wages cannot account for much of the downturn. Naturally, there are limits to the persuasiveness of an essentially statistical exercise. But, in the absence of a full-fledged economic model, this exercise suggests no additional explanation may be needed.

NOTES

[1]See Blinder, (2009), Krugman (2009), and Romer (2009).

[2]Author's calculations based on data from the U.S. Bureau of Economy Analysis, *National Income and Product Accounts of the United States*.

[3]These were computed as the marginal tax rate weighted by the number of returns in each bracket above $4,000, using numbers in U.S. Department of the Treasury (1938, p. 88; 1940, pp. 119, 193). Barro and Sahasakul (1983) find much lower average marginal tax rates because the number of filers was only 20 percent of all households, and they assign a zero marginal tax rate to the other 80 percent.

[4]Member banks made quarterly reports, whereas nonmember banks reported twice a year. Furthermore, some member banks, representing 82 percent of all member banks by assets, reported statistics on a weekly basis.

[5]The Federal Deposit Insurance Corporation, which began operations in January 1934, levied a premium on participating banks based on total deposits, and insured deposits up to $5,000 per depositor. In 1936, insured deposits amounted to $22,230 million, representing 68 percent of the deposits of participating banks and 47 percent of all bank deposits (Board of Governors of the Federal Reserve System, 1943, p. 401).

[6]Banking Act of 1935 (49 Stat. 706).

[7]There were two central reserve cities, namely, New York and Chicago, and 60 reserve cities (see the list in Board of Governors of the Federal Reserve System, 1943, p. 401).

[8]Banking Act of 1935 (49 Stat. 706).

[9]Beney (1936) and National Industrial Conference Board (1938).

[10]The data come from the NBER Macrohistory Database, series 08142 and 01300. Wages are deflated by the Wholesale Price Index.

[11]The VAR is monthly; all variables are in logs except the surplus, which can be negative. A time trend and seasonal dummies are added because the surplus is not seasonally adjusted. Lag length, chosen to minimize the Akaike information criterion, is 3. As an alternative, I have also used (the log of) man-days idle due to strikes instead of wages.

[12]There is some negative correlation between wages adjusted for labor productivity and money. An alternative specification in which average hourly earnings are not adjusted for productivity yields a nearly diagonal matrix, and the results of the historical decomposition are quite similar.

[13]The impulse response function of wages on output is negative at first but turns positive after ten months. This suggests that the shock identified as a wage shock is more complex than a shock to workers' bargaining power and probably includes a productivity component. If man-days idled by strikes is used instead of wages, the impulse response function is consistently negative, but of smaller magnitude: The percentage of IP variance explained by innovations to man-days at the three-year horizon is 2.5 percent instead of 6.5 percent for average hourly earnings.

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Employment growth: Cyclical movements or structural change?

Ellen R. Rissman

Introduction and summary

The Federal Reserve, in its policy analysis, must carefully weigh incoming data and evaluate likely future outcomes before determining how best to obtain its twin goals of employment growing at potential and price stability. It is tempting to regard high or rising unemployment as a sign of a weak economy. And, normally, a weak economy is one with little inflationary pressure and, therefore, room for expansionary monetary policy to stimulate growth. But unemployment is influenced by more than simply aggregate conditions. In a dynamic economy that responds to changing opportunities, some industries are shrinking while others are growing. Labor must flow from declining industries to expanding ones. This adjustment takes time. It takes time for employees in declining sectors to learn about new opportunities in other industries, acquire necessary skills, apply for job openings, and potentially relocate. And during this period of adjustment, the unemployment rate rises as waning industries lay off workers. Thus, the unemployment rate may increase or decrease, even though the aggregate state of the economy remains stable, simply because the labor market adjusts to shifting patterns of production.

For policymakers, it is essential to decipher what portion of a rising unemployment rate is due to a cyclical slowdown in which many sectors of the economy are simultaneously affected, as opposed to a structural realignment in production in which particular sectors of the economy are affected. The two factors ideally should result in different policy responses. If unemployment is rising because of a weak economy, the textbook response is for the Fed to take a more accommodative policy stance. If, instead, the unemployment rate is rising because of underlying compositional shifts in employment, an easing of monetary policy may discourage declining industries from contracting by keeping them marginally profitable, impeding the adjustment process. Furthermore, this policy may also encourage inflation as employers across a broad spectrum of industries compete for scarce labor resources. Thus, comprehending the underlying sources of movements in the unemployment rate is more than just a theoretical exercise: It has practical implications for monetary policy.

As a first step toward evaluating the role of structural change, I need to be able to measure it. Lilien (1982) suggests a dispersion measure that is a weighted average of squared deviations of industry employment growth rates from aggregate employment growth. Abraham and Katz (1986) argue that Lilien's measure does not properly account for cyclical shifts in employment across industries, instead conflating cyclical variation with structural change. When aggregate economic conditions are weak, certain sectors are affected more than others because demand for their products is more cyclically sensitive, but as soon as economic conditions improve, these sectors will also recover more quickly. The Lilien measure more accurately captures both cyclical variation in employment responses and structural changes in the composition of employment across industries, making it impossible to disentangle the importance of the two effects on the measure of dispersion.

The sectoral shifts hypothesis has been revisited more recently by Phelan and Trejos (2000) and Bloom, Floetotto, and Jaimovich (2009). Phelan and Trejos (2000) calibrate a job creation/job destruction model to data from the U.S. labor market to suggest that permanent changes in sectoral composition can precipitate aggregate economic downturns. Bloom, Floetotto, and Jaimovich (2009) examine the effect of what they term "uncertainty shocks" on business cycle dynamics,

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arguing that increases in uncertainty lead to a decline in economic activity in affected industries, followed by a rebound. Increasing uncertainty, in their view, causes firms to be more cautious in their hiring and investment decisions and impedes the reallocation of capital across sectors. Thus, structural change and recessions are simultaneous events, implying that distinguishing structural change from cyclical downturns is problematic.

As noted by Bloom, Floetotto, and Jaimovich (2009), structural realignment (in other words, sectoral reallocation) may be concurrent with economic downturns. Businesses on the brink of downsizing or disappearing altogether may find that they are tipped over the edge during a recession. To the extent that whole industries are affected, the downturn will then occur at the same time as sectoral reallocation. Recessions are followed by expansions, whereas sectoral reallocation tends to have a long-term impact on the composition of employment. Therefore, shifts in production that are cyclical in nature tend to be transitory, but those that are the result of structural realignment are more long lasting.

Previous studies, including Loungani, Rush, and Tave (1990) and Rissman (1993), have employed a variety of techniques to distinguish between sectoral shifts that are driven by structural change and those that are driven by cyclical swings. Loungani, Rush, and Tave (1990), for example, suggest that stock market prices efficiently reflect the future stream of business profits. They employ measures based on stock prices to create a dispersion measure that reflects structural shifts rather than short-term cyclical fluctuations. In Rissman (1993), I note that structural change is long lasting, whereas cyclical swings are of a shorter duration. I use this observation to distinguish between compositional shifts in employment that are due to cyclical fluctuations, which are short term, and those that are due to structural realignment, which are long term. Rissman's (1993) measure cannot be produced in real time because current changes in employment patterns may be either temporary or permanent. Thus, this measure offers little guidance for policymakers who need to make decisions based on current information. In contrast, the Loungani, Rush, and Tave (1990) measure has the benefit of being based on stock price data that are available at high frequency. However, stock prices are noisy, and it may be difficult to disentangle the persistence of shocks from them. In particular, a given decline in a stock price may be a reflection of short-run factors or may instead be interpreted as a small permanent decline in an industry's fortunes. Having a supplementary employment-based measure that does not require the use of leading data, in contrast to Rissman (1993), would provide a useful benchmark.

This problem of optimally inferring the current state has been widely studied in economics and in related statistical literature. Stock and Watson (1989) employ the Kalman filter to create an index of coincident economic indicators. They formally operationalize the idea that the business cycle "refers to co-movements in different forms of economic activity, not just fluctuations in GNP [gross national product]."[1] Stock and Watson (1989) examine several different economic time series, including employment, and try to extract a common factor. I use the same approach here to identify a common factor in the labor market based on how it affects employment in different industries. This common factor is permitted to have different loadings in each industry, giving some context to the notion that some sectors are more cyclically sensitive than others. This framework has the added benefit of creating a common factor that can be interpreted as a measure of the employment cycle, focusing only on the industry cross section of employment growth. This is particularly relevant, since it is widely thought that the labor market typically lags the business cycle. Thus, a measure of the business cycle based only on cross-sectional employment growth helps clarify the relationship between more traditional measures of the cycle, such as real gross domestic product (GDP) growth, and employment growth. This measure of the cycle may help shed light on the phenomenon of the jobless recoveries that we have experienced during the two most recent expansions following the contractions ending in 1991:Q1 and 2001:Q4. Furthermore, the model is based upon quarterly data, giving policymakers a more timely tool for evaluating the relative importance of cyclical and structural factors to the labor market than other measures. There is little reason why the model cannot be estimated on a monthly basis as well. Finally, the model provides some insight into the sources and magnitude of structural change in the economy.

To summarize the results, most industries exhibit cyclical employment growth, which accounts for the majority of the variation in employment in those industries. However, structural shifts are also important and account for most of the variation in employment growth in the finance, insurance, and real estate (FIRE) sector and in the government sector. Perhaps not surprisingly, given the well-chronicled declines in the housing market, the construction industry has undergone a structural reduction in employment after a notably long period of structural expansion. Recent structural employment declines in finance, insurance, and real estate are particularly large when compared with past episodes. Careful measurement of structural change suggests that sectoral reallocation may have been on the rise in the past few quarters. However, structural realignment cannot account for much of the recent increase we have observed in the unemployment rate.

In the next section, I examine employment growth for nine industries comprising most of total nonfarm employment. Then, I introduce the estimation framework. I present my results using this framework. Finally, I develop a measure of sectoral reallocation and investigate its impact on the unemployment rate.

Industry employment growth

The U.S. Bureau of Labor Statistics collects detailed industry employment data for workers on nonfarm payrolls. Over the years the industry classification system has changed to reflect the changing industrial composition of the economy. Because of this, it is difficult to compare earlier industry data, which were collected using the Standard Industrial Classification (SIC) System, with more recent industry data, which were collected using the North American Industry Classification System (NAICS). For example, nine new service sectors and 250 new service industries are recognized in the NAICS data, but they are not in the SIC data. The problem of comparability over time is less of an issue with the broadest industry aggregates. Earlier estimates of sectoral reallocation were computed using SIC data. To facilitate comparison with earlier work, the NAICS data were converted as closely as possible to be consistent with SIC classifications.

Figure 1 shows annualized quarter-to-quarter employment growth from 1950 through the second quarter of 2009 for the following nine sectors: construction; durable manufacturing; nondurable manufacturing; transportation and utilities; wholesale trade; retail trade; finance, insurance, and real estate; services; and government.[2] Business cycle contractions, as determined by the National Bureau of Economic Research (NBER), have been shaded for reference. The figure also shows the average annual industry employment growth rate over this period.

Given the current focus on the housing market as the source of some of our economic problems, it is interesting to examine employment in the construction sector. Employment growth in construction is highly volatile and, not surprisingly, quite cyclical as well. Construction employment growth appears to decline in advance of business cycle peaks and reaches its bottom at or just past the trough of a recession. Although employment growth in construction was above average during the most recent expansion, which peaked in December 2007, the strong employment growth does not appear abnormally large in comparison with earlier recoveries. Nonetheless, the most recent quarters show a very strong drop in construction employment, surpassing even the large declines of the mid-1970s. It is an open question as to what part of this observed decline in construction is structural in nature and what part is cyclical (and will therefore rebound when aggregate conditions improve).

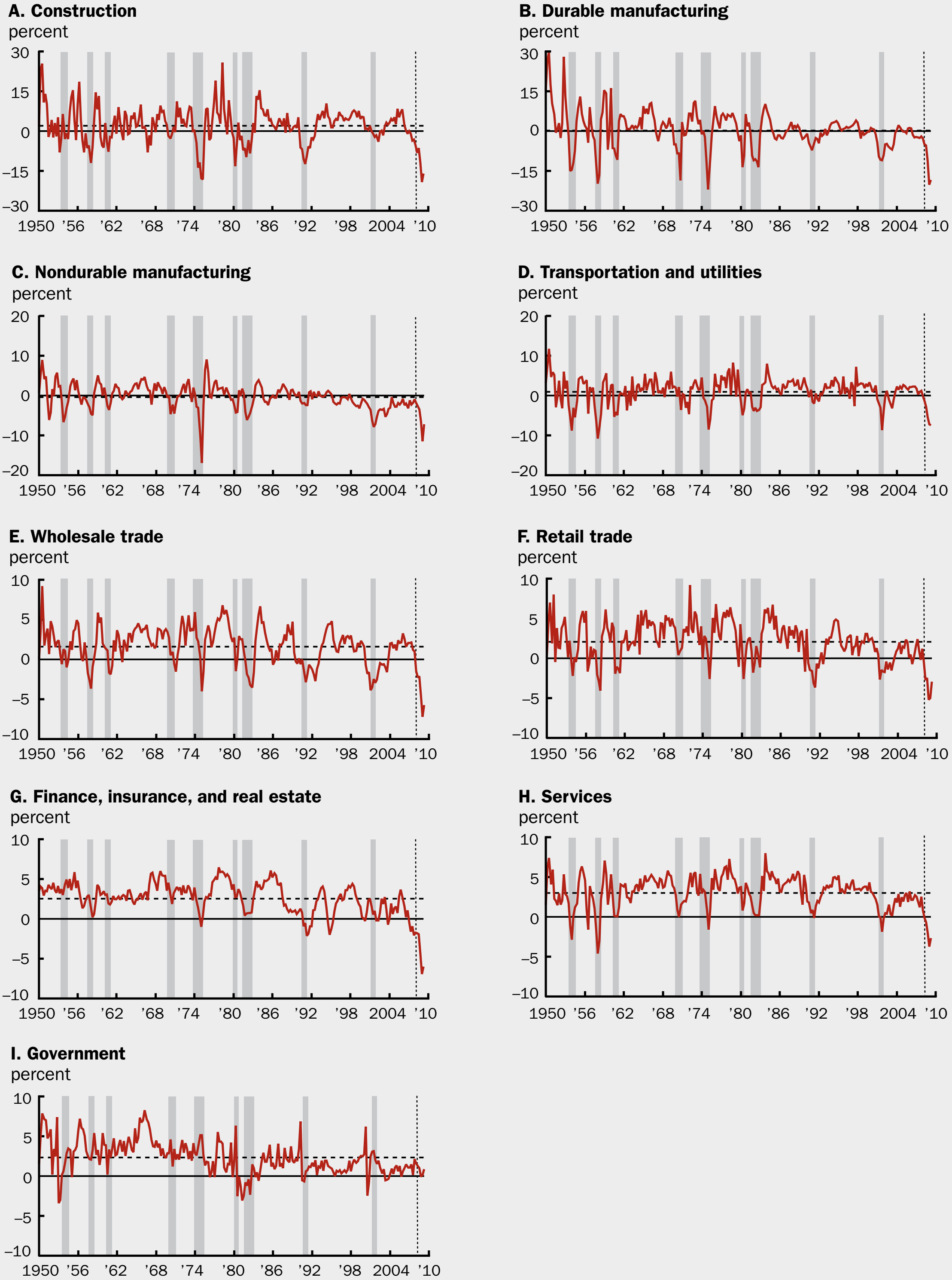
The finance, insurance, and real estate sector tells a somewhat different story. Like most industries, FIRE experiences reduced employment growth during recessions. Yet, while FIRE's employment growth has dipped below average during recessions, historically, employment in this sector has very rarely declined. The steep drop in employment in the early 1990s seems to be the harbinger of a change in employment growth in this sector, with average employment growth falling below the 3 percent growth of earlier decades. Furthermore, the steep job losses of the past few quarters are unprecedented in the past 60 years. The key question is whether the sharp employment declines are cyclical, with employment likely to rebound as the economy moves into the expansionary phase of the business cycle, or structural and, therefore, likely to linger. Later, I will show that employment growth in this industry tends to be highly persistent, suggesting that these declines are likely to last for quite a while. Yet, these job losses in FIRE may not transfer directly into increased unemployment. Since workers in FIRE may have skills that are more easily transferred to other areas, they may be more likely to find employment in expanding sectors; therefore, the adjustment out of this sector may not involve much of an increase in the unemployment rate.

The services sector is also interesting to consider. At one time, this sector was thought to be the engine of employment growth, as can be seen by the high average employment growth rates since the 1950s. Yet, more recently, employment growth here has been weak as well. And employment growth in services over the past couple of quarters is the lowest it has been since the late 1950s.

Taken as a whole, these data suggest several important facts. First, average growth rates differ across industries, with some sectors of the economy barely growing at all, such as durable and nondurable manufacturing, and others exhibiting more robust growth, such as FIRE and services. Second, some industries are far more volatile than others. Construction, durable and nondurable manufacturing, and transportation and utilities have wide swings in employment growth compared with the other industries. Third, unsurprisingly, employment growth is highly cyclical, dropping during contractions and rising during expansions. However, some industries appear more cyclically sensitive than others. Focusing on the period since the onset

FIGURE 1

Employment growth: Selected industries, 1950:Q1–2009:Q2



Notes: These are quarterly annualized growth rates calculated on an SIC (Standard Industry Classification system) conformable basis. The dashed horizontal line in each panel is the average annual industry employment growth rate. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line in each panel indicates the most recent business cycle peak.

Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

of the current recession in the fourth quarter of 2007, employment has declined precipitously in most industries. If most of the recent declines in employment growth are cyclical, then employment growth should rebound and return to normal as the economy moves into the expansionary phase of the business cycle. However, a portion of the recent declines in employment growth may be the result of other factors such as structural realignment in the economy. If this is indeed the case, then it may indicate that some industries will likely experience more permanent reductions in employment or employment growth. An accurate assessment of whether employment data are driven by the business cycle or structural change is important for formulating policy and for projecting the future path of employment growth.

Table 1 shows the same employment growth data for the entire sample in the first row and divided into ten-year increments in the subsequent rows.[3] Construction employment has averaged 2.0 percent annualized quarterly growth over the entire sample period. However, over the past decade the average quarterly growth in construction employment has been –0.42 percent. Durable and nondurable manufacturing have experienced large declines in employment over the past decade, with job losses or stagnant growth since the late 1970s. Employment growth has been weak for the past decade in transportation and utilities, as well as in wholesale and retail trades. In fact, all sectors have exhibited weaker average employment growth over the past decade than they have averaged over the past 60 years.[4]

A model of industry employment growth

The discussion in the previous section suggests that industry employment growth, in addition to having a long-term average, can be described by two additional components: a cyclical component and an idiosyncratic component that reflects other noncyclical factors. Let

1) $g_{it} = a_i + C_{it} + X_{it}$,

where g_{it} is employment growth in sector i at time t, $i = 1, \ldots, I$, and $t = 1, \ldots, T$; a_i is average employment growth in the industry; C_{it} is the cyclical portion of industry employment growth (and it varies across time and industry); and X_{it} is the idiosyncratic part of industry employment growth (and it also varies across time and industry). This construction is similar to the problem analyzed by Stock and Watson (1989), in which they noted that individual aggregate time series depend upon a common cyclical component and an idiosyncratic component.

TABLE 1

Average annualized quarterly employment growth, in total and by decade

	Construction	Durable manufacturing	Nondurable manufacturing	Transportation and utilities	Wholesale trade	Retail trade	Finance, insurance, and real estate	Services	Government
	(- - - - - - - -	- - - - - - - -	- - - - - - - -	- - - - - - - -	percent	- - - - - - - -	- - - - - - - -	- - - - - - - -	- - - - - - - -)
Total	2.00	0.37	−0.40	1.00	1.63	2.06	2.53	3.00	2.29
2000s	−0.42	−3.82	−3.41	−0.29	−0.41	−0.18	0.18	1.16	1.04
1990s	2.42	−0.04	−0.78	1.76	1.20	1.44	1.56	3.24	1.28
1980s	1.63	−0.90	−0.25	1.32	1.57	2.52	2.97	3.72	1.13
1970s	2.64	0.98	0.09	1.30	2.97	3.33	3.59	3.62	2.65
1960s	2.08	2.50	1.23	1.30	2.42	3.03	3.38	3.70	4.15
1950s	3.54	3.31	0.58	0.52	1.93	2.08	3.37	2.45	3.41

Note: Data are seasonally adjusted.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

As currently specified, equation 1 cannot be estimated because there is no way to distinguish between the cyclical and idiosyncratic components. To address this issue, I assume that the cycle is a common component affecting all industries. However, the cycle may have a differential impact across sectors. Specifically,

$$2)\quad C_{it} = b_i^1 C_t + b_i^2 C_{t-1},$$

where b_i^1 and b_i^2 are parameters indicating the sensitivity of the i-th sector to current and lagged values of the business cycle. Furthermore, it is assumed that the cycle itself follows a second-order autoregressive process with:

$$3)\quad C_t = \phi_1 C_{t-1} + \phi_2 C_{t-2} + u_t.$$

Here u_t is independent and identically normally distributed with unit variance. The ϕ_1 and ϕ_2 are unknown parameters that are to be estimated. Setting $\sigma_u^2 = 1$ determines the scale of the business cycle. For example, an alternative estimate of the cycle $C_t^* = \delta C_t$ would result in estimates of the b_i values scaled by 1/δ. Two sets of estimates are possible, both C_t and $-C_t$, depending upon the initial values of the parameters. For ease of interpretation, it is assumed that the business cycle has a positive impact on durable manufacturing employment growth.

The idiosyncratic component of industry employment growth X_{it} is assumed to follow an AR(1) process. Specifically,

$$4)\quad X_{it} = \gamma_i X_{it-1} + \varepsilon_{it},$$

where γ_i is a sector-specific parameter that indicates the degree of persistence of sectoral shocks. It is assumed that the ε_{it} values are uncorrelated over time and across industries. Note that $E(\varepsilon_{it}) = 0$ and $E(\varepsilon_{it}^2) = \sigma_i^2$ for all i, t. Furthermore, the ε_{it} values are assumed to be uncorrelated with the cyclical shock u_t for all i, t. This specification allows for a common unobserved cycle that has a differential impact across industries. It also permits structural change to occur through the idiosyncratic component X_{it}. Thus, changes in an industry's employment growth are due to either cyclical factors or factors that are specific to that particular industry.

Estimation is accomplished using the Kalman filter, details of which are discussed in box 1. The state vector $\underline{z}_t$ is given by $\underline{z}_t = [C_t, C_{t-1}, C_{t-2}, X_{1t}, X_{2t}, \ldots, X_{It}]'$. The Kalman filter algorithm enables estimates of the state vector $\underline{z}_t$ and the underlying parameters to be estimated. These parameters include the values for a_i, b_i^1, b_i^2, γ_i, σ_i, and ϕ_1 and ϕ_2. The shocks u_t and ε_{it} can also be obtained for $i = 1, \ldots, I$ and $t = 1, \ldots, T$. The Kalman filter is a way of optimally updating the underlying state vector as new information becomes available each quarter. A Kalman smoothing algorithm is used to optimally backcast for final estimates of the state vector and model parameters.

Estimation results

The estimate of the cycle $\hat{C}_t$ obtained from the Kalman filter exercise is shown in figure 2.[5] The 2× standard error bands are also shown. These standard error bands indicate whether the estimate is significantly different from zero. Defining a recession as the period during which the estimated employment cycle is significantly below zero, the estimate indicates that we are currently in the midst of a deep recession. The cyclical point estimate in 2009:Q1 measures the recession to be the most severe since 1950. However, because of parameter uncertainty, this point estimate is not significantly worse than earlier recessions in a statistical sense. The estimate for 2009:Q2 indicates that aggregate employment continues to deteriorate, albeit at a slower pace.

Employment failed to rebound as quickly as other sectors of the economy during the two most recent recoveries following the NBER-dated recessions of 1990–91 and 2001. This lack of improvement in the labor market, termed the "jobless recovery," drew commentary from both the popular press and economists. As computed here, the employment-based measure of the cycles indicates that the contractions lasted seven and eleven quarters, respectively—significantly longer than the length of the NBER's contractionary periods of three and four quarters, respectively—indicating that the labor market experienced a delayed recovery relative to other measures of economic activity that the NBER's Business Cycle Dating Committee examines in determining business cycle peaks and troughs. Shortly after the 2001 recession, Groshen and Potter (2003) suggested that the abnormally slow recovery was the result of sectoral reallocation (in other words, structural factors) rather than cyclical factors. The evidence provided here shows that the slow growth in employment was likely attributable to weak cyclical activity.[6] Using a similar methodology, Aaronson, Rissman, and Sullivan (2004) reach a similar conclusion. Furthermore, findings presented in the next section regarding the role of X_{it} appear to show that sectoral shocks do not play a major role in accounting for unemployment. Recall that the employment cycle is defined by co-movement in employment growth rates across many industries simultaneously. As such, the model interprets the lengthy employment contraction during these two episodes as broad-based; that is, a wide spectrum of industries are negatively affected,

BOX 1

The Kalman filter

The Kalman filter is a statistical technique that is useful in estimating the parameters of the model specified in equations 1–4 (pp. 44–45). In addition, the Kalman filter enables the estimation of the processes u_t and ε_{it} and the construction of the unobserved cyclical variable C_t and the idiosyncratic components X_{it}. The Kalman filter consists of a state equation and a measurement equation. The state equation describes the evolution of the possibly unobserved variable(s) of interest, $\underline{z}_t$, while the measurement equation relates observables $\underline{g}_t$ to the state. The vector $\underline{g}_t$ is related to the $m \times 1$ state vector, $\underline{z}_t$, via the measurement equation:

$$\text{B1)}\quad \underline{g}_t = B\underline{z}_t + D\underline{\eta}_t + H\underline{w}_t,$$

where $t = 1, \ldots, T$; B is an $N \times m$ matrix; $\underline{\eta}_t$ is an $N \times 1$ vector of serially uncorrelated disturbances with mean zero and covariance matrix I_N; and $\underline{w}_t$ is a vector of exogenous (possibly predetermined) variables with H and D being conformable matrices.

In general, the elements of $\underline{z}_t$ are not observable. In fact, it is this very attribute that makes the Kalman filter so useful to economists. Although the $\underline{z}_t$ elements are unknown, they are assumed to be generated by a first-order Markov process as follows:

$$\text{B2)}\quad \underline{z}_t = A\underline{z}_{t-1} + F\underline{u}_t + G\underline{w}_t,$$

for $t = 1, \ldots, T$, where A is an $m \times m$ matrix, F is an $m \times p$ matrix, and $\underline{u}_t$ is a $p \times 1$ vector of serially uncorrelated disturbances with mean zero and covariance matrix I_g. This equation is referred to as the state or transition equation.

The definition of the state vector $\underline{z}_t$ for any particular model is determined by construction. In fact, the same model can have more than one state-space representation. The elements of the state vector may or may not have a substantive interpretation. Technically, the aim of the state-space formulation is to set up a vector $\underline{z}_t$ in such a way that it contains all the relevant information about the system at time t and that it does so by having as small a number of elements as possible. Furthermore, the state vector should be defined so as to have zero correlation between the disturbances of the measurement and transition equations, $\underline{u}_t$ and $\underline{\eta}_t$.

The Kalman filter refers to a two-step recursive algorithm for optimally forecasting the state vector $\underline{z}_t$, given information available through time $t-1$, conditional on known matrices B, D, H, A, F, G. The first step is the prediction step and involves forecasting $\underline{z}_t$ on the basis of $\underline{z}_{t-1}$. The second step is the updating step and involves updating the estimate of the unobserved state vector $\underline{z}_t$ on the basis of new information that becomes available in period t. The results from the Kalman filtering algorithm can then be used to obtain estimates of the parameters and the state vector $\underline{z}_t$ by employing traditional maximum likelihood techniques.[1]

The model of employment growth proposed here can be put into state-space form, defining the state vector $\underline{z}_t = [C_t, C_{t-1}, C_{t-2}, X_{1t}, X_{2t}, \ldots, X_{It}]'$. The Kalman filter technique is a way to optimally infer information about the parameters of interest and, in particular, the state vector $\underline{z}_t$, which in this case is simply the unobserved cycle, C_t, and its two lags and the unobserved structural components X_{it}. The cycle, as constructed here, represents that portion of industry employment growth that is common across the industries while allowing the cycle to differ in its impact on industry employment growth in terms of timing and magnitude through the parameters b_i^1 and b_i^2. The model is very much in the spirit of Burns and Mitchell's (1946) idea of cycles entailing co-movement, but the estimation technique permits the data to determine which movements are common and which are idiosyncratic.[2]

[1]The interested reader may obtain further details in Harvey (1989) and Hamilton (1994).

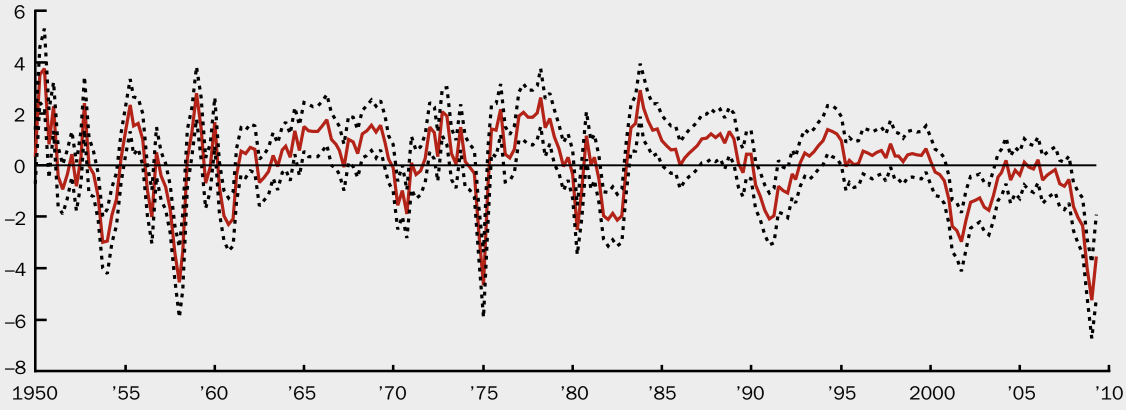
[2]Stock and Watson (1989) employ the Kalman filter in constructing leading and current economic indicators.

and the contraction is not concentrated in only a few industries, as would be the case if sectoral reallocation were the underlying cause of low aggregate employment growth.

Table 2 provides parameter estimates with associated standard errors. Focus on the coefficient estimates of the $\hat{b}_i^1$ values (second column): All sectors of the economy are affected by cyclical variation, as constructed here. However, the degree of cyclical sensitivity varies across industries, with durable manufacturing employment being the most contemporaneously cyclically sensitive, followed by construction. The estimated intercept term $\hat{a}_i$ (first column) is not significantly different from zero in construction, durable manufacturing, nondurable manufacturing, and transportation and utilities. The estimated parameter $\hat{\gamma}_i$ (fourth column) gives the degree of persistence of the idiosyncratic component. There is a great deal of variation in the persistence of these idiosyncratic shocks ε_i, with finance, insurance, and real estate exhibiting the most persistence.

FIGURE 2

Estimated employment cycle, 1950:Q1–2009:Q2



Note: The dashed lines indicate the 2× standard error bands, indicating whether the estimate is significantly different from zero.
Source: Author’s calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

TABLE 2

Parameter estimates, 1950:Q1–2009:Q2

	$\hat{a}_i$	$\hat{b}_i^1$	$\hat{b}_i^2$	$\hat{\gamma}_i$	$\hat{\sigma}_i$
Construction	1.8435 (1.1963)	1.8695*** (0.3407)	1.5357*** (0.4730)	0.4240*** (0.0741)	20.1902*** (1.8060)
Durable manufacturing	0.2714 (1.5350)	3.7417*** (0.3549)	0.8463 (0.6397)	0.612*** (0.0552)	9.9197*** (0.9951)
Nondurable manufacturing	−0.4657 (0.6295)	1.5231*** (0.1537)	0.4054 (0.2385)	0.6461*** (0.0479)	1.9574*** (0.2371)
Transportation and utilities	0.9105 (0.5921)	1.2185*** (0.2354)	0.7769** (0.3036)	0.0933 (0.0883)	3.8068*** (0.4611)
Wholesale trade	1.5546*** (0.4448)	0.8004*** (0.1135)	0.6365*** (0.1888)	0.5516*** (0.0673)	1.2072*** (0.1134)
Retail trade	1.9921*** (0.4377)	1.2430*** (0.1589)	0.2449 (0.2255)	0.1727* (0.0818)	1.7845*** (0.2004)
Finance, insurance, and real estate	2.3220*** (0.6162)	0.2073* (0.0913)	0.2340*** (0.0862)	0.8978*** (0.0361)	0.7583*** (0.0786)
Services	2.9343*** (0.4061)	1.0738*** (0.0994)	0.3221 (0.1661)	0.1728 (0.1119)	0.4891*** (0.0804)
Government	2.2712*** (0.3532)	0.0890 (0.1280)	0.1438 (0.1031)	0.5748*** (0.0639)	2.9139*** (0.2494)

*Significant at the 5 percent level.
**Significant at the 2 percent level.
***Significant at the 1 percent level.
Note: Standard errors are in parentheses.
Source: Author’s calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

BOX 2

Calculating the variance

Rewriting the model as a vector AR(1) process, define

B3) $\underline{y}_t = [g_{1t}, g_{2t}, \ldots, g_{It}, C_t, C_{t-1}, C_{t-2}, X_{1t}, X_{2t}, \ldots, X_{It}]'.$

Then

B4) $\underline{y}_t = \Pi \underline{y}_{t-1} + \underline{\nu}_t,$

which has a variance

B5) $\Omega = \Pi \Omega \Pi' + \Sigma.$

This can be solved as:

B6) $\text{vec}(\Omega) = [I - \Pi \otimes \Pi]^{-1} \text{vec}(\Sigma),$

where $\otimes$ is the Kronecker product of Π with itself and vec(x) is the vector constructed by stacking the columns of an $n \times m$ matrix into a single column vector. The matrix Π is given by

B7) $$\Pi = \begin{bmatrix} 0_{I\times I} & B_{I\times 3} & \Gamma_{I\times I} \\ 0_{3\times I} & A_{3\times 3} & 0_{3\times I} \\ 0_{I\times I} & 0_{I\times 3} & \Gamma_{I\times I} \end{bmatrix},$$

and the submatrices are given by

B8) $$B = \begin{bmatrix} b_1^1 & b_1^2 & 0 \\ b_2^1 & b_2^2 & 0 \\ \vdots & \vdots & \vdots \\ b_I^1 & b_I^2 & 0 \end{bmatrix},$$

B9) $$A = \begin{bmatrix} \phi_1 & \phi_2 & 0 \\ 1 & 0 & 0 \\ 0 & 1 & 0 \end{bmatrix},$$

and

B10) $$\Gamma = \begin{bmatrix} \gamma_1 & 0 & \ldots & 0 \\ 0 & \gamma_2 & & 0 \\ \vdots & & \ddots & \\ 0 & \ldots & 0 & \gamma_I \end{bmatrix}.$$

The error term $\underline{\nu}_t$ is given by

B11) $\underline{\nu}_t = [\varepsilon_{1t}, \varepsilon_{2t}, \ldots, \varepsilon_{It}, u_t, 0, 0, \varepsilon_{1t}, \varepsilon_{2t}, \ldots, \varepsilon_{It}]'.$

Shocks to both services and transportation and utilities are not statistically persistent. Furthermore, variation in these shocks differs across industries, reflecting in part the variation in employment growth noted in figure 1 (p. 43). Shocks to the idiosyncratic portion of industry employment growth are more variable in construction, durable manufacturing, and transportation and utilities than in other sectors of the economy (fifth column).

Using the model, it is straightforward to calculate the portion of the variation in an industry's employment growth that is attributable to cyclical activity and that which is attributable to industry-specific factors. Details of the calculations are found in box 2, and the results are presented in table 3. As noted previously, some industries exhibit much more variation in employment growth than others. Construction and durable manufacturing are the two most volatile sectors of the economy, exhibiting large swings in employment growth. By comparison, the variance of employment growth in nondurable manufacturing and transportation and utilities is about one-fifth that of the most volatile industries, and the least volatile sectors have about one-tenth the variance. The model attributes this volatility to either cyclical variation or the idiosyncratic structural component. Within construction, for example, about half the total variance in employment growth stems from the structural component and half is the result of cyclical variation. The cyclical component accounts for most of the variation in employment growth in durable manufacturing, nondurable manufacturing, transportation and utilities, wholesale trade, retail trade, and services. In contrast, the structural component carries the most weight in two sectors—FIRE and government.

In addition to examining the estimated cycle, it is also useful to consider the idiosyncratic portion of employment growth. Figure 3 shows the idiosyncratic component X_{it} for each of the nine industries from 1950:Q1 through 2009:Q2. Positive values suggest that employment growth is stronger in these industries than explained by either normal cyclical variation C_{it} or long-term averages a_i. Note that the scale differs

TABLE 3

Effect of cyclical and structural components on variation, 1950:Q1–2009:Q2

	Total variance	Fraction of total variance due to C	Fraction of total variance due to X_i
Construction	46.9245	0.4754	0.5246
Durable manufacturing	58.7445	0.7300	0.2700
Nondurable manufacturing	10.8692	0.6909	0.3091
Transportation and utilities	11.5470	0.6674	0.3326
Wholesale trade	5.7097	0.6961	0.3039
Retail trade	6.3833	0.7119	0.2881
Finance, insurance, and real estate	4.2831	0.0874	0.9126
Services	4.4125	0.8857	0.1143
Government	4.4569	0.0236	0.9764

Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

from one industry to the next. Upon closer inspection of the construction sector (figure 3, panel A), the estimates suggest that employment growth in this industry was higher than could be explained from the business cycle or sectoral trends over most of the 1990s through the first half of 2006, when the trend abruptly reversed, reflecting the unfolding crisis in the housing market. The sharp drop in X_{it} shows that construction employment seems to be taking a bigger hit in the current episode than can be explained based on the usual prior cyclical patterns for this sector. Perhaps even more noteworthy is the recent experience in finance, insurance, and real estate (figure 3, panel G) that shows a marked decline in recent years, suggesting this sector is in the midst of a restructuring that is unexplained by either the normal cyclical pattern or long-term trends. How this downsizing of FIRE affects the unemployment rate is an open question.

As table 1 (p. 44) suggested, the parameters of the model may change over time. A test of parameter stability can be done using a likelihood ratio test. The test statistic compares the log likelihood of the model estimated using the full sample, from 1950:Q1 through 2009:Q2, with the sum of the log likelihoods from the model estimated on two smaller samples—the 1950:Q1–1983:Q4 period and the 1984:Q1–2009:Q2 period. The resulting test statistic is distributed $X^2(46)$, and its value is 498.22, rejecting the hypothesis that at normal confidence levels the parameter vector is the same for the two smaller sample periods.

Table 4 presents parameter estimates from the 1984:Q1–2009:Q2 sample period. In comparing the estimates found in table 2 (p. 47) and table 4, there is some evidence of "The Great Moderation,"[7] with most of the coefficients on the contemporaneous estimate of the cycle, $\hat{b}_i^1$, being smaller in magnitude for the 1984:Q1–2009:Q2 sample period than for the entire sample. For example, in the full sample a one standard deviation increase in the cycle increased durable manufacturing employment growth by 3.7 percent per annum, whereas in the 1984:Q1–2009:Q2 sample, the impact was a much smaller 1.2 percent (see second row, second column of tables 2 and 4, respectively). Furthermore, generally, estimates of the variance of the idiosyncratic shocks in each industry, $\hat{\sigma}_i$, are much smaller for the 1984:Q1–2009:Q2 sample, with the exception of finance, insurance, and real estate (compare the fifth column in tables 2 and 4). For example, the estimate of the standard deviation in the shock to construction is 20.2 for the entire sample, but a much smaller 4.3 for the 1984:Q1–2009:Q2 sample. There is also evidence that for the 1984:Q1–2009:Q2 sample, industry shocks are more persistent, as can be seen by comparing the estimated $\hat{\gamma}_i$ values for the entire sample and those for the 1984:Q1–2009:Q2 sample, with government being a notable exception (see the fourth column in tables 2 and 4). Nonetheless, the interpretation of the results seems to hold. In particular, when estimated on the 1984:Q1–2009:Q2 sample, X_{it} in construction shows the run-up in construction employment starting in the mid-1990s and the abrupt decline in 2006 that cannot be explained by the typical cyclical patterns of the past. The estimated X_{it} values are shown in figure 4 for the two samples.

Sectoral reallocation

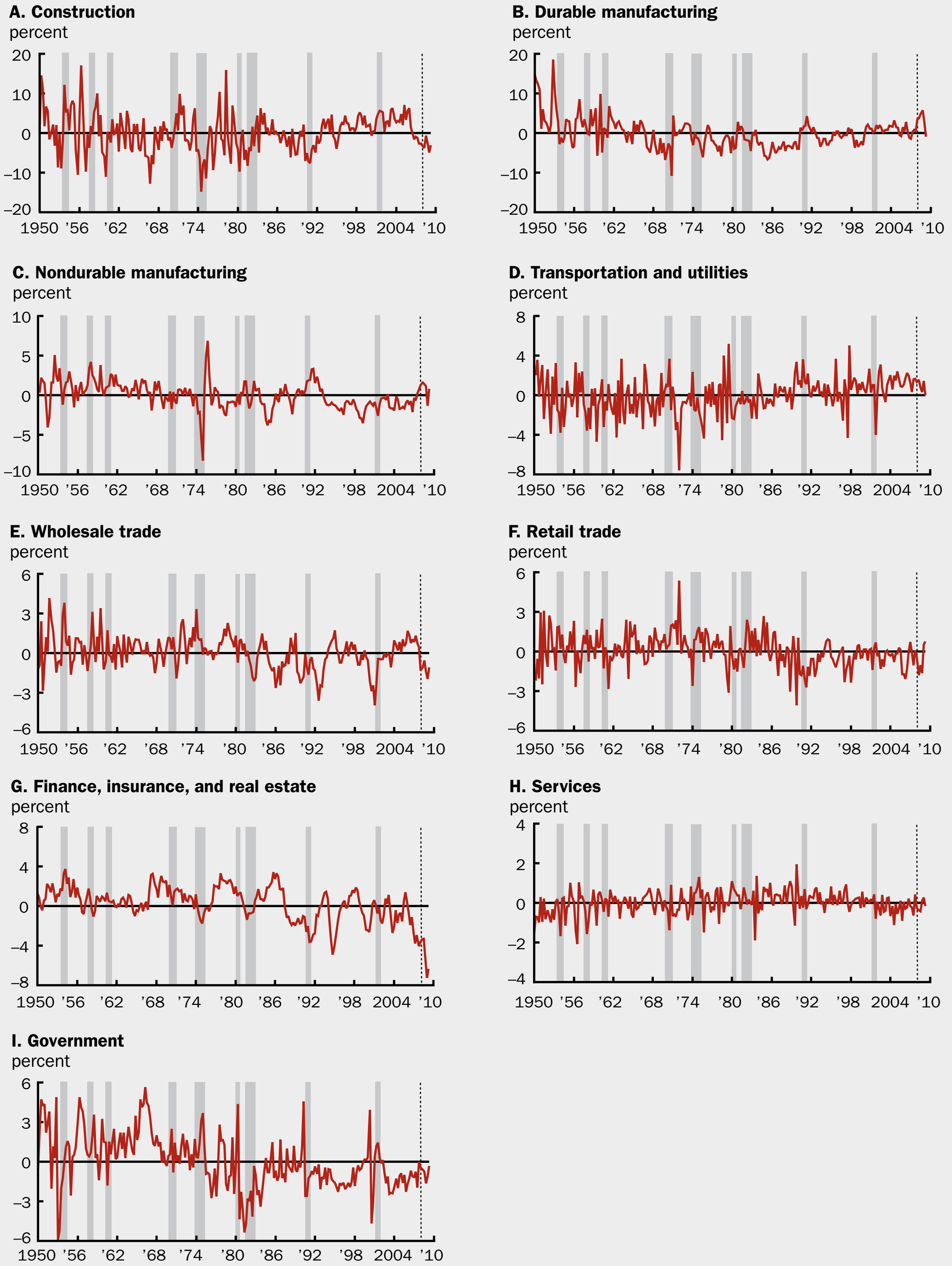
In his original paper, Lilien (1982) presented a dispersion measure as a way to quantify the degree of sectoral reallocation occurring in the economy at any given time. His measure is given by

$$5)\quad \sigma_{Lt} \equiv \left[\sum_i s_{it}(g_{it} - g_t)^2\right]^{1/2},$$

where s_{it} is industry i's employment share at time t; g_{it} is employment growth in i at time t; and g_t is total employment growth at time t. Abraham and Katz (1986) demonstrate that this dispersion measure will increase even if no sectoral reallocation is present, simply because some industries are more cyclically sensitive than others.

FIGURE 3

Noncyclical employment growth: Selected industries, 1950:Q1–2009:Q2



Notes: The panels show the estimated X_{it} values. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

TABLE 4

Parameter estimates, 1984:Q1–2009:Q2

	$\hat{a}_i$	$\hat{b}_i^1$	$\hat{b}_i^2$	$\hat{\gamma}_i$	$\hat{\sigma}_i$
Construction	1.1703	2.0239***	0.1505	0.7837***	4.3028***
	(5.8890)	(0.4471)	(0.7041)	(0.1066)	(1.5117)
Durable manufacturing	–2.0404	1.2233***	0.9706***	0.7809***	1.4940***
	(5.8171)	(0.3697)	(0.3687)	(0.1028)	(0.3755)
Nondurable manufacturing	–1.7455	0.6885***	0.3827	0.7261***	0.8580***
	(2.8657)	(0.2296)	(0.2630)	(0.1025)	(0.2103)
Transportation and utilities	0.9551	0.6666*	0.4383	0.0793	2.0082***
	(2.8641)	(0.3072)	(0.3680)	(0.1213)	(0.3299)
Wholesale trade	0.5980	0.7366***	0.2876	0.7551***	0.6228***
	(2.6710)	(0.1784)	(0.2258)	(0.0807)	(0.1714)
Retail trade	1.0153	0.7983***	0.1708	0.3190*	1.0826***
	(2.5684)	(0.2299)	(0.3379)	(0.1487)	(0.2504)
Finance, insurance, and real estate	1.1042	0.2289	0.2021	0.8818***	0.9104***
	(1.5869)	(0.1927)	(0.1860)	(0.0883)	(0.2466)
Services	2.5218	0.6012***	0.2775	0.0339	0.3434***
	(2.2840)	(0.1699)	(0.1683)	(0.2188)	(0.0882)
Government	1.3070***	–0.1437	0.2893	0.2061*	1.4475***
	(0.4682)	(0.2637)	(0.2858)	(0.1003)	(0.2520)

*Significant at the 5 percent level.
**Significant at the 2 percent level.
***Significant at the 1 percent level.
Note: Standard errors are in parentheses.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

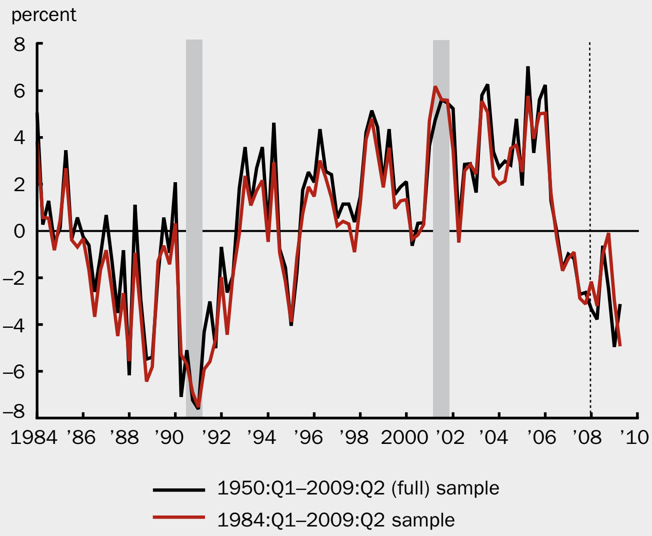
Keep in mind the Abraham and Katz (1986) criticism that Lilien's (1982) dispersion measure reflects cyclical movements: The framework presented previously provides a way to eliminate the impact of the cycle on employment shares, industry employment growth, and aggregate employment growth so as to create a dispersion measure that is purged of cyclical variation. This measure is given by

$$6)\quad \tilde{\sigma}_t \equiv \left[\sum_i \tilde{s}_{it}(\tilde{g}_{it} - \tilde{g}_t)^2\right]^{1/2},$$

where $\tilde{x}$ indicates that the variable x is purged of the cycle. To create the purged series, first, let $\tilde{g}_{it} = X_{it}$. Then, assuming that the cycle was zero in some reference year, taken here to be 1964, it is simple to calculate $\tilde{e}_{it}, \tilde{e}_t, \tilde{s}_{it}$, and $\tilde{g}_t$, where $\tilde{e}_{it}$ is noncyclical employment in industry i at time t and $\tilde{e}_t$ is total noncyclical employment at time t. Figure 5 shows the results of these calculations. The red line is Lilien's (1982) measure as given in equation 5, and the black line is calculated as in equation 6. The noncyclical measure of dispersion is far less volatile than the original measure, as Abraham and Katz (1986) argued. Nonetheless, there has been a modest uptick in this measure of structural realignment over the past couple of quarters. Figure 6 shows the noncyclical measure in panel A and another measure that is based only on the shocks ε_{it} in panel B. In this figure you can see the recent uptick more clearly. The most recent quarter shows a decline in these dispersion measures, reflecting industry shocks that are smaller in magnitude than those of the previous few quarters. However, while it suggests a potential role for industrial realignment in explaining recent increases in unemployment, this simple summary measure may not be too informative in explaining recent changes in the unemployment rate. To put it more succinctly, structural realignment in and of itself may have little impact on the unemployment rate. Workers laid off in one sector may be readily absorbed into other industries, particularly if real wages adjust to encourage the flow of workers from declining industries to expanding ones.

FIGURE 4

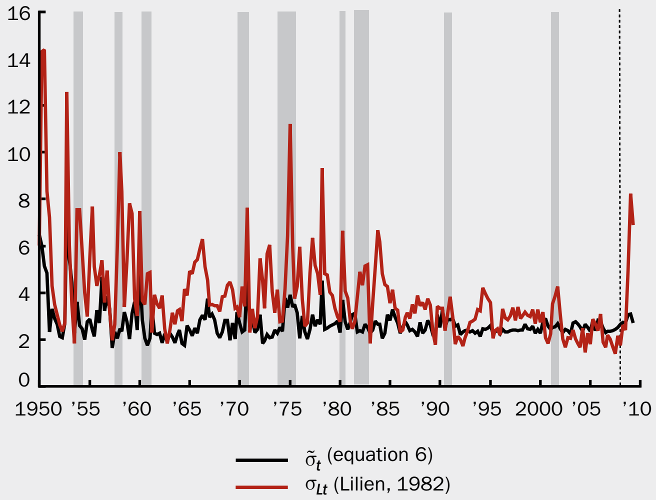
Estimated idiosyncratic component in construction: Full sample versus 1984:Q1–2009:Q2 sample



Notes: See the text for further details on the idiosyncratic component (X_{it}) of industry employment growth, which is estimated on the two samples. The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

FIGURE 5

Dispersion measures, 1950:Q1–2009:Q2



Note: The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

In order to determine whether the structural component of employment growth plays a role in unemployment dynamics, I ran regressions of the following form:

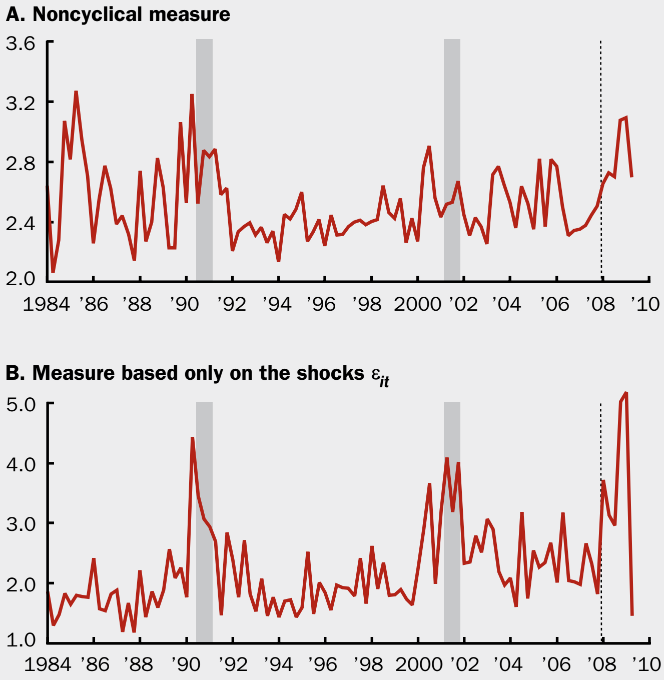
$$7) \quad \Delta ur_t = \alpha(L)\Delta ur_{t-1} + \delta(L)Cycle_t + \lambda(L)\Sigma_t + cW_t + \upsilon_t,$$

where $\alpha(L)$, $\delta(L)$, and $\lambda(L)$ are polynomials in the lag operator L; Δur_t is the change in the unemployment rate at time t; $Cycle_t$ is a measure of the cycle at time t; Σ_t is a measure of sectoral reallocation at time t, including the constructed dispersion measures or, more broadly, the individual estimated X_{it} and ε_{it} values; and W_t is other variables that potentially influence changes in the unemployment rate. The variable υ_t is a random shock assumed to be independent and identically normally distributed.

Two separate measures of the cycle were examined, namely, deviations of real GDP growth from its long-term average $(gGDP_t - \overline{gGDP})$ and $\hat{C}_t$. Several different measures of Σ_t were considered, including the two noncyclical measures computed as in equation 6, as well as the estimated X_{it} values and the ε_{it} values individually. Regression results are shown in table 5. Three lags of changes in the unemployment rate are included in each regression, as is a demographic variable that is calculated as the change in the female labor force participation rate of white women aged 20 and above. (Other demographic variables that reflected changes in the age, race, and sex composition of the labor force were also investigated but were statistically insignificant and are not reported in these results.) Of the two cyclical variables considered, the measure of the employment cycle $\hat{C}_t$ performed better than deviations of real GDP growth from its long-term average, in that those regressions had higher $\bar{R}^2$ values. Generally, the two dispersion measures of sectoral reallocation did poorly in explaining changes to the unemployment rate. The third and fourth columns examine the impact of adding dispersion measures of sectoral reallocation to the regressions. These

FIGURE 6

Noncyclical measures of sectoral reallocation, 1984:Q1–2009:Q2



Note: The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

dispersion measures are statistically significant, but enter with the opposite sign anticipated by the sectoral reallocation hypothesis; that is, increasing reallocation, as measured here, tends to reduce the unemployment rate.[8] The last two regressions omit the cyclical variable, $\hat{C}_t$, and include the two dispersion measures. Only in the results of the sixth column, in which the cyclical variable is omitted, does dispersion enter significantly positive. The weak results suggest that sectoral reallocation as measured here may be positively associated with changes in the unemployment rate. However, once cyclical effects are properly accounted for, the impact disappears or changes sign.

One possibility is that these dispersion measures, being summary statistics, are not very good at capturing the effects of reallocation in the labor market. The dispersion measure treats all employment shifts of the same magnitude as identical, regardless of the industry. This ignores the possibility that human capital may differ across industries, suggesting that unemployment responses should differ across sectors as well. Specifically, some industries may require industry-specific human

TABLE 5

Regression results: Dependent variable is Δur_t, 1984:Q1–2009:Q2 sample

	1	2	3	4	5	6
3 lags Δur_{t-1}	Yes	Yes	Yes	Yes	Yes	Yes
Current and two lags of $gGDP_t - \overline{gGDP}$	−0.0433*** (0.0074)	—	—	—	—	—
Current and two lags of $\hat{C}_t$	—	−0.1674*** (0.0264)	−0.1771*** (0.0264)	−0.1980*** (0.0291)	—	—
Change in female participation rate	0.1126 (0.0749)	0.0690 (0.0652)	0.1240 (0.0696)	0.1270 (0.0686)	0.0639 (0.0950)	0.0362 (0.0911)
$\tilde{\sigma}$ based on $\hat{X}_{it}$	—	—	−0.0131* (0.0065)	—	0.0073 (0.0077)	—
$\tilde{\sigma}$ based on $\hat{\varepsilon}_{it}$	—	—	—	−0.0175* (0.0077)	—	0.0182* (0.0079)
$\overline{R}^2$	0.6714	0.7904	0.7970	0.7993	0.5987	0.6162

*Significant at the 5 percent level.
**Significant at the 2 percent level.
***Significant at the 1 percent level.
Notes: Estimating over the full sample did not materially change the results. The full sample was estimated from 1954:Q2 through 2009:Q2, since the female labor force participation rate data are not available prior to 1954:Q2. The estimate of the employment cycle employed in the analysis is from the 1950:Q1–2009:Q2 Kalman filter exercise. Standard errors are in parentheses.
Sources: Author's calculations based on data from the U.S. Bureau of Labor Statistics and U.S. Bureau of Economic Analysis from Haver Analytics.

TABLE 6

Effect of idiosyncratic components and shocks on changes in the unemployment rate, 1954:Q2–2009:Q2

	X_{it}			ε_{it}		
	Coefficient and standard error	$\bar{R}^2$	Coefficient and standard error	Coefficient and standard error	$\bar{R}^2$	Coefficient and standard error
Construction	−0.0111*** (0.0025)	0.7910	−0.0117*** (0.0027)	−0.0093*** (0.0027)	0.7836	−0.0093*** (0.0027)
Durable manufacturing	−0.0174*** (0.0044)	0.7876	−0.0191*** (0.0045)	−0.0156** (0.0050)	0.7818	−0.0102 (0.0062)
Nondurable manufacturing	−0.0022 (0.0076)	0.7718	−0.0013 (0.0082)	0.0010 (0.0105)	0.7718	0.0074 (0.0129)
Transportation and utilities	−0.0146* (0.0065)	0.7770	−0.0041 (0.0070)	−0.0119 (0.0062)	0.7756	−0.0043 (0.0064)
Wholesale trade	0.0161 (0.0104)	0.7743	0.0180 (0.0103)	0.0159 (0.0125)	0.7735	0.0166 (0.0126)
Retail trade	0.0242** (0.0098)	0.7781	0.0108 (0.0103)	0.0212* (0.0099)	0.7766	0.0132 (0.0111)
Finance, insurance, and real estate	0.0001 (0.0074)	0.7718	0.0031 (0.0072)	0.0171 (0.0139)	0.7734	0.0201 (0.0130)
Services	0.0596*** (0.0212)	0.7799	0.0228 (0.0252)	0.0508*** (0.0192)	0.7790	0.0448 (0.0243)
Government	0.0126* (0.0062)	0.7761	0.0146** (0.0059)	0.0224*** (0.0076)	0.7806	0.0227*** (0.0072)
			$\bar{R}^2 = 0.8179$			$\bar{R}^2 = 0.8064$

*Significant at the 5 percent level.
**Significant at the 2 percent level.
***Significant at the 1 percent level.
Notes: Dependent variable is Δur_t. Also included in the regressions are three lags of the dependent variable, one current and two lags of the estimated employment cycle, and changes in the labor force participation rate of white women aged 20 and above. See the text for further details.
Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

capital. Sectoral reallocation away from those industries will take time and cost more for those who have become displaced. To examine this possibility, I have entered the idiosyncratic components both individually and together. The results are found in tables 6 and 7, which differ only in their sample periods. Table 6 provides results for the period from 1954:Q2 through 2009:Q2, and table 7 provides results from 1984:Q1 through 2009:Q2.[9]

The first two columns of table 6 examine the effect of including each idiosyncratic component separately in a regression having both cyclical and demographic variables. The $\bar{R}^2$ values are reported from each of these regressions in the second column. The sectors of the economy in which the idiosyncratic component of employment growth is statistically significant are construction, durable manufacturing, transportation and utilities, retail trade, services, and government. The signs of these effects are also interesting to consider. Specifically, as noncyclical employment grows above trend in construction, durable manufacturing, and transportation and utilities, it reduces the unemployment rate. However, it has the opposite effect in retail trade, services, and government, in that shifts toward these industries tend to raise the unemployment rate. The third column reports the coefficients from a single regression in which all idiosyncratic industry components are included, in addition to current and lagged employment cycle and demographic variables. Noncyclical shifts in construction, durable manufacturing, and government are still statistically significant, entering with the same sign as in the single variable regressions. However, transportation and utilities, retail trade, and services are no longer statistically significant.

TABLE 7

Effect of idiosyncratic components and shocks on changes in the unemployment rate, 1984:Q1–2009:Q2

	X_{it}			ε_{it}		
	Coefficient and standard error	$\bar{R}^2$	Coefficient and standard error	Coefficient and standard error	$\bar{R}^2$	Coefficient and standard error
Construction	−0.0101* (0.0045)	0.7990	−0.0129* (0.0063)	−0.0085 (0.0055)	0.7935	−0.0125* (0.0059)
Durable manufacturing	−0.0025 (0.0070)	0.7884	−0.0054 (0.0087)	−0.0128 (0.0110)	0.7912	−0.0123 (0.0132)
Nondurable manufacturing	0.0130 (0.0101)	0.7919	0.0031 (0.0145)	0.0152 (0.0175)	0.7898	0.0139 (0.0208)
Transportation and utilities	−0.0254*** (0.0090)	0.8048	−0.0272** (0.0108)	−0.0246*** (0.0086)	0.8050	−0.0243** (0.0094)
Wholesale trade	0.0108 (0.0113)	0.7987	−0.0001 (0.0137)	0.0007 (0.0162)	0.7881	−0.0024 (0.0172)
Retail trade	0.0252* (0.0117)	0.7981	0.0206 (0.0143)	0.0239 (0.0121)	0.7966	0.0221 (0.0146)
Finance, insurance, and real estate	−0.0048 (0.0070)	0.7892	−0.0043 (0.0079)	0.0130 (0.0139)	0.7901	0.0107 (0.0137)
Services	0.0025 (0.0305)	0.7882	0.0102 (0.0376)	0.0003 (0.0274)	0.7881	0.0149 (0.0346)
Government	0.0087 (0.0099)	0.7997	0.0019 (0.0099)	0.0005 (0.0101)	0.7881	−0.0023 (0.0099)
			$\bar{R}^2$ = 0.8153			$\bar{R}^2$ = 0.8104

*Significant at the 5 percent level.
**Significant at the 2 percent level.
***Significant at the 1 percent level.

Notes: Dependent variable is Δur_t. Also Included in the regressions are three lags of the dependent variable, current and two lags of the estimated employment cycle, and changes in the labor force participation rate of white women aged 20 and above. See the text for further details.

Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

The fourth, fifth, and sixth columns of table 6 repeat the regression exercise but instead employ idiosyncratic shocks ε_{it} as explanatory variables. The results are consistent with the results using X_{it}. Shocks to construction and durable manufacturing tend to reduce unemployment, whereas shocks to retail trade, services, and government tend to raise unemployment (fourth column). The transportation and utilities industry does not meet the 5 percent significance criterion. However, its marginal significance level is close to 10 percent. Table 7 reestimates the equations of the preceding table, but with the 1984:Q1–2009:Q2 sample period. Most of the results disappear for this sample period.

To obtain estimates of the effect of sectoral reallocation on the unemployment rate, I assume that the economy was in equilibrium in 2007:Q4, with an unemployment rate of 4.8 percent. Furthermore, I assume that the cycle is set equal to its expected value from 2007:Q4 through 2009:Q2. In this analysis, that implies that $C_t = 0$. I also assume that there are no demographic changes in the female labor force participation rate over this period.

Table 8 provides estimates of the effect of X_{it} on the civilian unemployment rate as estimated from the equation used in the third column of table 7, using the 1984:Q1–2009:Q2 sample period. The first column gives the estimated total effect of the X_{it} on the unemployment rate, given the assumptions in the preceding paragraph. The impact of sectoral reallocation in this model is negligible. The remaining columns compute the impact on the equilibrium unemployment rate of having idiosyncratic employment growth shocks in the specified industry given by the estimated shocks. For example, although equilibrium employment remained

TABLE 8

Estimated impact of idiosyncratic industry employment growth on unemployment, 1984:Q1–2009:Q2

	All X_{it}	Construction	Durable manufacturing	Nondurable manufacturing	Transportation and utilities	Wholesale trade	Retail trade	Finance, insurance, and real estate	Services	Government
	(- - - -	- - - -	- - - -	- - - -	percent	- - - -	- - - -	- - - -	- - - -	- - - -)
2007:Q4	4.80	4.80	4.80	4.80	4.80	4.80	4.80	4.80	4.80	4.80
2008:Q1	4.79	4.84	4.78	4.80	4.76	4.80	4.79	4.82	4.80	4.80
2008:Q2	4.76	4.89	4.76	4.81	4.72	4.80	4.75	4.83	4.79	4.80
2008:Q3	4.71	4.91	4.73	4.81	4.69	4.80	4.73	4.85	4.79	4.80
2008:Q4	4.69	4.94	4.70	4.82	4.68	4.80	4.69	4.87	4.79	4.79
2009:Q1	4.73	5.01	4.68	4.81	4.64	4.80	4.70	4.90	4.79	4.79
2009:Q2	4.82	5.06	4.68	4.82	4.63	4.80	4.71	4.94	4.79	4.79

Notes: Results are for the regression in table 7, third column. Calculations assume that the equilibrium unemployment rate was equal to its value of 4.8 percent in 2007:Q4 and that the employment cycle is in equilibrium from 2007:Q4 through 2009:Q2, so that $X_{it} = 0$ from 2007:Q4 through 2009:Q2. The first column calculates the unemployment rate that would have occurred had the industry idiosyncratic components been as estimated from 2007:Q4 through 2009:Q2 and the employment cycle been in equilibrium. The second through tenth columns reflect the impact of the idiosyncratic components in each of the individual industries. For example, in the second column the estimated impact of idiosyncratic shifts in construction on the unemployment rate in 2009:Q2, assuming all other industry components to be as given by the estimated X_{it} values, is to raise the unemployment rate by 0.24 percentage points (calculated by subtracting the value in the last row, first column, from the value in the last row, second column, 5.06 – 4.82). Results differ if the unemployment rate regression is estimated using the entire sample period.

Source: Author's calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

largely unchanged, by 2009:Q2 the shocks to construction raised the unemployment rate by approximately 25 basis points (see the notes in table 8). This rise was offset by declines elsewhere.

As a whole, these models suggest that idiosyncratic shifts in industry employment growth account for very little of the observed increase in the unemployment rate over the past several quarters. On its own, this would imply that there is room for accommodative policy as a response to the current increase in unemployment, but bringing to bear additional evidence on dispersion would help us gain a better sense of whether the conclusions implied by the empirical model discussed here are robust. There is a great deal of uncertainty surrounding the estimates presented here. As noted before, the parameters of the state-space model appear to differ between the 1950:Q1–1983:Q4 period and the 1984:Q1–2009:Q2 period. Because of parameter and model uncertainty, these estimates of the impact of sectoral reallocation on the unemployment rate must be viewed somewhat skeptically. To underscore this fact, results of the same exercise that estimate the unemployment equation using the full sample suggest a decline in unemployment since 2008:Q1 attributable to sectoral reallocation.

Conclusion

The labor market appears to have a cycle that is well described by co-movements in employment growth. The estimate of the employment cycle that results from my model seems to agree with anecdotal evidence about jobless recoveries. The model also does a good job of capturing turning points in the business cycle, suggesting that it may be a useful tool for understanding labor market dynamics and may help in predicting future employment. The idiosyncratic component that the methodology yields may also provide some additional insight into the impact of structural realignment on changes in the unemployment rate. Structural change favoring construction, durable manufacturing, and transportation and utilities seems to be associated with decreasing unemployment; this suggests that there may be some impediments to displaced workers in these sectors finding jobs in other industries. Even with the downsizing of finance, insurance, and real estate, the overall impact on the unemployment rate is not statistically significant. One possibility is that employees from finance, insurance, and real estate are better able to find alternative employment in other sectors of the economy because the skills they possess are more readily transferable to employment in other industries. Conversely, employees in construction, durable manufacturing, and transportation and utilities may be less readily absorbed into other sectors.

NOTES

[1]Stock and Watson (1989), p. 353.

[2]The services sector includes information services, professional and business services, education and health services, leisure and hospitality, and other services. Mining has been omitted from the analysis for two reasons. First, because of the incidence of strikes, employment growth in this industry is quite volatile. Second, mining accounts for a small fraction of total employment.

[3]Averages for the current decade are based on data through 2009:Q2.

[4]The only exception, unreported here, is the mining sector.

[5]The hat symbol (^) indicates an estimate.

[6]There is another notable discrepancy when comparing the NBER business cycle recession dates with those estimated here. The two NBER recessions in the early and mid-1970s were longer by two and three quarters, respectively, than those proposed here. Instead, the employment-based measure of the cycle shows a labor market that was quick to return to more normal activity during those times.

[7]The Great Moderation is a term used to describe the period usually thought to have begun in 1984 and lasting through the present, during which many economic time series exhibited less volatility than in previous years. The validity of this concept as a permanent shift has been called into question by the recent financial crisis.

[8]The coefficients reported here are for contemporaneous measures of dispersion. Including a number of leads and lags did not substantively change the results. Altering the specification so that the dispersion measure was in changes or log changes had no bearing on the results either.

[9]The full sample period is slightly shortened by starting in 1954:Q2 because earlier data for female labor force participation were not available.

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