Disability Benefit Take-Up and Local Labor Market Conditions*

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Abstract

Exploiting county-level variation in oil-producing areas from shocks to world oil and gas prices from 1970–2011, we study how local labor market conditions affect disability take-up. Our analysis extends well-known previous work using a similar research design by analyzing a different price shock, a larger and more representative set of labor markets, and a more recent period marked by skyrocketing disability payments. Our estimated elasticity for SSDI receipt with respect to earnings of −0.29 is surprisingly similar to that found in previous work. Our preferred estimate for the SSI elasticity of −0.16 is smaller than previous findings, but we show that most of the difference is explained by changes in the SSI program.

JEL Codes: J0, H0

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1 Introduction

Understanding the labor market disincentive effects of government-provided disability insurance benefits has long been of interest to economists. While larger benefit amounts lead workers to apply for benefits and exit the labor force, better labor market opportunities have a countervailing effect on such transitions (Parsons 1980; Haveman and Wolfe 1984; Bound 1989; and Bound and Waidmann 1992). Early work (e.g., Parsons 1980) finds that increases in the share of earnings replaced by disability benefits lower employment rates although the potential endogeneity of the simple replacement rates in these studies may complicate the interpretation of these early findings (Bound 1989). One strand of more recent work, examining arguably exogenous local labor demand shocks, finds that worse labor market conditions lead to increased levels of disability payments (Black, Daniel, and Sanders 2002; Autor and Duggan 2003; Autor, Dorn, and Hanson 2013; Sloane 2015). Other recent work uses random assignment of decision makers in the disability benefit determination process to estimate the labor supply impact of benefit awards (Maestas, Mullen, and Strand 2013; French and Song 2014).

This paper examines the extent to which US Social Security Disability Benefits (SSDI) and Supplemental Security Income (SSI) payments, the two largest Federal programs that provide assistance to people with disabilities, are affected by local labor market earnings levels.[1] To circumvent the potential reverse causality problem arising from the possibility that local areas with chronically poorer health likely will have both higher rates of disability claiming and lower earnings, we use variation in local earnings generated by shocks to oil and gas production due to

[1] In December 2013, 8.9 million former workers and 2 million dependent spouses and children received SSDI and 9 million people received federal SSI. Source: OASDI Beneficiaries by State and County, 2013; SSI Annual Statistical Report, 2013. Disability is defined by law as the "inability to engage in any substantial gainful activity by reason of any medically determinable physical or mental impairment". An applicant's education, age, and work experience are considered when determining whether she may still work.

exogenous movements in the world prices of oil and gas. Our research design is akin to the seminal article of Black, Daniel, and Sanders (2002), who use county-level movements in coal production in four states in the Appalachian region during the energy price boom and bust in the 1970s and 1980s, although our analysis differs from and extends their work in several key ways.

First, whereas Black et al. study the boom and bust cycle in energy prices in the 1970s and 1980s, our analysis covers the period from 1970 to 2011, which spans both the earlier cycles and more recent run up in energy prices during the 2000s. Second, the demographic composition of workers in the oil and gas industry is more representative of U.S. workers overall than are workers in the coal industry studied by Black et al.[2] Third, compared to the four eastern U.S. states from Black et al.'s study, the eleven states with the largest share of oil and gas employment used in our analysis are more broadly representative, as they include places from along the Gulf of Mexico to the central U.S. To the extent that disability benefits offer higher replacement rates to lower-earning workers (Autor and Duggan 2003), disability take-up for the typical American worker, who more closely resembles the oil and gas workers we study, may exhibit very different responses to earnings shocks that do the coal workers to shocks in their sector. Given these differences, our work provides new and timely evidence on the relationship between disability programs and local labor market conditions.[3]

We find that increases in county earnings lead to significant decreases in both county SSDI and SSI payments. Our baseline estimates of the elasticity of benefit payments to earnings for SSDI and SSI are −0.29 and −0.16, respectively, and are robust to numerous specifications of our

[2] Appendix Table 1 presents characteristics of all U.S. workers, workers in the oil and gas industry and workers in the coal industry in 1970, 1990, and 2010. The numbers for 1970 and 1990 are computed using the Public Use Micro Sample of the Decennial Censuses in those years. The 2010 numbers are computed using 2008–2012 American Community Survey 5-year data files (Ruggles et al. 2015).

[3] Our analysis differs from the work of Autor, Dorn, and Hanson (2013) and Sloane (2015) in several key ways, apart from the fact that we study a very different type of shock. In particular, whereas these papers study more states, our analysis is conducted at the level of the county, exploits annual-level variation, and studies a longer time frame than either of these studies.

instrumental variable. Despite the differences across the studies in the time periods and samples they analyze, our SSDI elasticity is comparable to Black et al.'s estimate. We show that our much smaller estimate for the SSI elasticity compared to Black et al. can be substantially accounted for by the federalization of the SSI program in 1974, which produced uniform program standards across states and led to an immediate 60% increase in total benefit payments.

The remainder of the paper proceeds as follows. Section 2 lays out the empirical models we use for estimation while Section 3 discusses the construction of our county-level data. Section 4 presents our main results and robustness checks and Section 5 concludes.

2 Empirical Specification

To access the effects of local economic conditions on disability payments, we estimate the following equation:

$$\Delta_c y_{cst} = \beta_0 + \delta(\Delta_c E^*_{cst}) + \Delta_c \boldsymbol{x}_{cst}\boldsymbol{\beta}_1 + \boldsymbol{d}_{st}\boldsymbol{\beta}_2 + \varepsilon_{cst}, \quad (1)$$

where y_{cst} is the logarithm of real SSDI or SSI payments for county c in state s between year $t-1$ and t; $\boldsymbol{d}_{st}$ is a vector of state-year dummy variables for state s in year t to allow state-specific differences by year; $\boldsymbol{x}_{cst}$ is a vector of control variables; ε_{cst} is idiosyncratic factors of county c in year t; Δ_c indicates the difference between t and $t-1$ within the county.[4]

The regressor of interest, $\Delta_c E^*_{cst}$, is change in (log) earnings or employment of county c in year t. The variable is starred to denote the fact that we observe error-ridden versions of the

[4] We control for the logarithm of county population, the log growth of county population, whether the county is in a Metropolitan Statistical Area (MSA) for the 1990 Census, and the fraction of 1969 earnings from manufacturing. Population and MSA status are controlled as proxies for better access to health care and amenities such as public transportation, both of which are appealing to individuals with disabilities. If people with disabilities are over-represented in the migrants, then the population growth captures this mechanical impact on the disability payments. We use the fraction of 1969 earnings from manufacturing to control for any impact of industry structure on disability. This last variable is missing for some small counties due to disclosure concerns although our results are quite similar if we drop this variable and include these counties.

actual variable. Given the log-log specification, the parameter on this regressor, δ, is the elasticity of disability payments with respect to local economic conditions.

Some concerns arise in estimating δ using equation (1). One concern is that first difference estimators exacerbate the attenuation effects of measurement error (Bound, Brown, and Mathiowetz 2001). Also, as discussed in McKinnish (2007), year-to-year within-county variation often reflects transitory fluctuations that have little effect on long-term behavioral outcomes such as initiating SSDI benefits. Although permanent earnings is the relevant concept for the SSDI participation decision, the available earnings measure contains both permanent and transitory components. Another concern is that $\Delta_c E^*_{cst}$ is an endogenous regressor since changes in local economic conditions are likely correlated with changes in unobserved local factors that may affect disability participation. For example, if SSA field workers are more sympathetic to unemployed workers when local economy is suffering or if law firms spend more in disability-related advertisements in counties where economic conditions decline, OLS estimates of the relationship between disability payments and earnings will be biased downwards. Alternatively, more generous disability payments (less stringent screening criteria, higher replacement rates, etc.) may decrease labor force participation and bias the estimates in the opposite direction (Parsons 1980; Autor and Duggan 2003). The combined impact of these potential biases on the OLS estimate is therefore ambiguous.

We use two stage least squares (TSLS) to address both the measurement error and endogeneity concerns. We use variation in county-level earnings induced by exogenous shocks to international oil prices. As these prices surge (decline), both oil production as well as efforts to locate and extract new oilfields increases (decreases) leading to higher (lower) earnings in the affected counties. After accounting for state-year interactions, our instrumental variables strategy exploits differences across counties within the same state, as areas with higher baseline employment shares in the oil and gas industry experience larger earnings changes in response to oil price movements.

Our baseline instruments are constructed as the interaction between the oil price changes and the importance of oil and natural gas to a county, that is,

$$\Delta R_t \times Oil_c, \tag{2}$$

where ΔR_t measures the real oil price changes between periods t and $t-1$, and Oil_c is the initial importance of oil and gas industry in the county as measured by oil and gas employment share in a year before the sample period. Two lags of the instrument are also included in ΔR_t to account for the setup time of new wells and local businesses adjustment following international shocks. In addition, given the long disability benefit determination periods, disability benefit payments may take time to fully adjust to a given shock. The resulting TSLS model estimates equation (3), with the change in the local labor market measure replaced by the predicted value from the first stage equation

$$\Delta_c E^*_{cst} = \gamma_0 + \sum_{\tau=0}^{2} \alpha_\tau \Delta R_{t-\tau} \times Oil_c + \Delta_c \boldsymbol{x}_{cst} \boldsymbol{\gamma}_2 + \boldsymbol{d}_{st} \boldsymbol{\gamma}_3 + u_{cst}. \tag{3}$$

3 Data

The Bureau of Economic Analysis' Regional Economic Information System (REIS) provides annual county-level data on earnings and employment beginning in 1969. Earnings include wage and salary disbursements, other labor income, and proprietors' income. The Bureau of Labor Statistics (BLS) compiles REIS wage and salary disbursements using ES-202 filings collected as part of the state unemployment insurance program. The REIS data on county wages and employment are known to be measured with some error.[5] Dan Black kindly provided us with county SSDI payments in Decembers 1970–2001.[6] We collected more recent county SSDI payments from SSA's "OASDI Beneficiaries by State and County" Report (various years). County annual SSI received payments are from REIS. All dollar values are deflated by annual average CPI-U with the base year 2010.

The TSLS part of our analysis focuses on the effect of energy supply shocks in oil and natural gas producing counties in eleven "oil and gas states", spanning 1970 to 2011. We define the oil and gas states as those contiguous states where the mining industry comprises at least 1% of total employment in the 1974 County Business Patterns (CBP) and a substantial share of the mining establishments are related to oil and gas.[7] We construct a measure of the initial importance of oil and gas industry in a county, Oil_c, which is the share of employment in the oil and gas industry computed from CBP data.

Prior to 1974, we only observe one-digit total mining employment for each county in CBP. As the first large oil price shock occurs in 1974 and our sample begins in 1970, we would prefer to determine the initial oil and gas importance in some year prior to 1974, if not prior to 1970. The tradeoff, however, is that with an earlier year we can only determine mining employment as opposed to oil and gas employment at the county-level. As shown in Appendix Table 2, over half of mining establishments in the oil and gas states, with the exception of Utah, are in the oil and

[5] Some employers with establishments in multiple counties may only report wages and employment ES-202 information at the state-level. These reports are allocated back to counties based on their industry level distribution by county among employers reporting at the county-level, generating some measurement error. In addition, components of other labor income and proprietors' income such as pension plan contributions, health and life insurance contributions, and private worker's compensation contributions are only collected at the state-level and also use an allocation rule to determine county-level totals.
[6] County-level SSDI payment data was not published in 1981.
[7] These eleven states are Colorado, Kansas, Louisiana, Mississippi, Montana, New Mexico, North Dakota, Oklahoma, Texas, Utah, and Wyoming. We exclude West Virginia since less than half of total mining employment is in the oil and gas industry which affects the construction of our 1967 instrument.

gas industry with little change in these shares between 1967 and 1974. Thus, we construct the county-level initial oil and gas importance measure, Oil_c, using the 1967 CBP and we present results using the 1974 CBP to compute Oil_c as robustness check.[8]

Figure 1 depicts the geographic distribution of the oil and gas industry by county within the eleven oil and gas states using CBP data from 1974. This figure illustrates the tremendous variation in the importance of oil and gas across counties in these states.

We obtain data on energy prices from the Energy Information Administration's Annual Energy Review.[9] From 1969 to the 1990s, there were at least three large exogenous shocks to the world oil supply: the 1973–74 OPEC oil embargo following the Yom Kippur War; the period from the end of 1979 to early 1981, following the overthrowing of the Shah of Iran and the start of the Iran-Iraq War; and the 1990–91 First Persian Gulf War. The late 1990s and 2000s saw booming oil demand from newly industrialized countries and stagnant world crude oil production, with occasional exogenous events such as the U.S. attack on Iraq and the turmoil in Nigeria (Hamilton 2009a and 2009b). During the same period, the advances in hydraulic fracturing and horizontal drilling technologies made shale gas commercially viable. These events and trends affected both the prices of oil and natural gas, as well as the employment in these industries in the United States, as Figure 2 illustrates.

Real oil prices doubled between 1973 and 1974, were stable for several years, then tripled over a three-year period. Prices fell sharply over the next five years to levels in the mid-1980s

[8] Because of the risk of disclosing firm specific information, exact employment numbers for two-digit industries are not available at the county level. However, the CBP provides county-level information on both the number of firms in each two-digit industry and the number of firms that fall into a specific firm-size category (e.g., 20 to 49 employees) for these industries. By weighting the number of firms in a firm-size category by the midpoint of the number of employees in that category, we create an estimate of the number of employees in each two-digit industry at the county level. We then create county-level estimated employment shares by industry as the ratio of the estimated industry employment to the estimated total county employment, where the total county employment is also estimated using the firm-size methodology.

[9] Oil prices are the US average first purchase price per barrel; natural gas prices are the wellhead price per thousand cubic feet. National oil and gas industry employment are from CBP.

that were slightly lower than those of the mid-1970s. After hovering around the same level for more than a decade, real oil prices started to climb up again in the early 2000s, reaching the all-time high in mid-2008, then collapsed in 2009. Real natural gas prices followed a very similar pattern to that for oil prices: a six-fold increase between 1970 and the early 1980s, then a decline of more than half over the next six years, and later a similar run-up and collapse in the 2000s. Figure 2 shows that national employment in the oil/natural gas industry closely tracked the movement in prices, with the main difference that national employment was not kept artificially flat during various periods, as was true for oil prices in two periods in the 1970s because of policy decisions.[10] As part of our robustness checks, we use national oil and gas employment changes as the measure of ΔR_t in place of oil price changes to construct our instruments.

4 Results

Table 1 presents our results from estimating equation (1) with both SSDI (Panel A) and SSI (Panel B) as outcomes using counties in the eleven oil and gas states.[11] Our OLS estimates, presented in column (1), show no significant relationship between annual county-level changes in earnings and the analogous changes in disability payments. However, as we discussed above, there are reasons to believe that these estimates will be inconsistent because the regressor $\Delta_c E^*_{cst}$ may be both being measured with error and endogenous in (1).

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[10] Direct federal control of crude oil prices lasted for almost a decade, from August 1971 to January 1981. In addition, since the early 1900s, oil production in the United States has been overseen by various state regulatory boards, such as the Oklahoma Corporations Commission, the Louisiana Conservation Commission and, most importantly, the Texas Railroad Commission. Although the specific language outlining each board's functions and objectives differ from state to state, these agencies set limits on level of extraction and exploration in their particular states so as to stabilize price, and prevent over-exploitation of oil reserves.

[11] The sample sizes for SSDI payment regressions are larger than for SSI payment regressions because the REIS contains missing payment data for counties in years when these payments are below a nominal $50,000. Thus, these missing values occur more frequently in earlier sample years and in smaller counties. In addition, we only present results from using a balanced panel for each payment outcome, although the results are quite comparable when we allow for an unbalanced panel. These results are available from authors.

We next present TSLS estimates of equation (1) in which we instrument for the change in log earnings using variation in oil prices and the geographic distribution of oil and gas production as shown in equation (3). The F-statistic testing the joint significance of the excluded instruments in the first stage equation, shown below each point estimate, greatly exceeds the conventional threshold of ten.[12] Over the entire sample period, which runs from 1970 through 2011, our TSLS estimates shown in column (2) of Table 1 are −0.29 and −0.16 for SSDI and SSI, respectively, and are highly statistically significant. Thus, local disability payments increase significantly when local earnings decline.[13]

Using earnings as the regressor of interest captures both the impact of employment and wage changes on disability benefit take-up. However, a decline in earnings due to lost employment (e.g., a plant closing) might have a different effect on the decision to apply for disability benefits than lower real wage growth, especially if lack of employment is a requirement for a successful disability benefit award. In the last two columns of Table 1, we estimate models identical to (1), except that $\Delta_c E^*_{cst}$ in these regressions is the change in log employment in order to focus on the extensive work margin. Our TSLS benefit payment elasticity estimates for employment (column (4)) are more than twice as large as the earnings elasticities (column (2)), indicating that a 10% reduction in employment leads to a much larger disability response than a 10% reduction in earnings. Throughout the remainder of the paper we only present results that use the earnings regressor as proportional difference between the

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[12] The corresponding first stage estimates are found in Appendix Table 3.

[13] We also estimated the baseline TSLS model separately for the two periods conventionally treated in the literature as being times of distinct oil "boom" (1973–1981) and oil "bust" (1982–1986). For SSDI receipt, the estimated elasticity with respect to earnings of −.547 (.132) during the boom and −.135 (.124) during the bust are quite different, which suggest that the propensity to participate in the program during bad labor market conditions versus to leave or not apply during good conditions might be asymmetric. For SSI, the estimated effects of − .072 (.149) and −.151 (.05) during the boom and bust are not statistically different, which suggests that for this program the effects might not be asymmetric.

estimates from the earnings and employment regressors choices is roughly the same as is found in Table 1.[14]

Table 2 analyzes the robustness of our results to a number of additional specification variants. The energy price variation that identifies our results comes from two eras: the energy price boom and bust cycle in the 1970s and 1980s as well as more recent sharp price movements that occurred in the 2000s. The first two columns of Table 2 split the sample along these time dimensions with the results for 1970–1993 shown in column (1) and the findings for 1994–2011 shown in column (2). For SSDI, the estimate for the boom and bust cycle is very similar to that for the full sample period. The result for the latter period is about 30% smaller than the point estimate for the full sample although both point estimates fall well within the 95% confidence interval for the other point estimate. For SSI, the boom and bust cycle estimate is slightly larger than the full sample estimate while the latter period estimate is not statistically different than zero.

The third column of Table 2 constructs Oil_c based on the 1974 oil and gas employment shares rather than the 1967 mining employment shares. Recall that 1974 is the earliest year in which the CBP data allows us to compute county-level employment shares for the oil and gas industry as opposed to the entire mining industry. Although the measure of Oil_c constructed using the 1974 data is based on information from after the start of the sample period we study, the resulting estimates are comparable to our baseline findings.

The models shown in next column in Table 2 use the contemporaneous oil shock as an instrument without also including two lags of the oil shock as in the baseline results. Since local businesses need time to respond to international price shocks, and the disability determination

[14] These results are available from the authors upon request.

process also takes many months to complete, especially given the time it may take to successfully appeal a claim that is initially denied, the total impact of an oil price shock in a given year will likely take multiple years to be realized. Indeed, in column (4) of Table 2, we find smaller estimates when we exclude lagged oil shocks from our instrument set, especially for SSDI payments.

We have noted previously that oil price movements were constrained because of various institutional features prior to the early 1980s. Given the infrequent adjustment of oil prices during this era, movements in oil and gas industry employment likely better represented producers' desires to extract and search for oil than oil price movements. We therefore use national oil and gas industry employment to construct the instrument ΔR_t rather than oil prices. In the final column of Table 2, we find somewhat larger responses of SSDI and SSI payments when using this alternative instrument.

Our study, which uses nearly twenty years more recent data, analyses a sample of workers, whose education is higher, who work in a different industry and live in different parts of the country compared to Black et al. How do our estimates compare with those in the earlier study? Our baseline estimates of the elasticity of benefit payments to earnings are −0.29 and −0.16 for SSDI and SSI, respectively. Black et al's corresponding baseline estimate for the SSDI elasticity of −0.345 is quite comparable to ours. However, their baseline estimate of the SSI elasticity of −0.713 is much larger than our finding. On the one hand, across a number of robustness checks, their baseline result for SSI is at the larger end of their range of estimates which spans −0.4 to −0.7.[15] On the other hand, a number of our robustness checks for SSI are closer to −0.2 including

[15] Black et al.'s SSDI estimates have a much tighter range of −0.3 to −0.4.

our estimates in which we restrict the sample to the 1970–1993 energy price boom and bust period.

The close similarity between our SSDI estimates and those from Black et al. suggests that the differences in the SSI results across the two studies are unlikely to be simply due to the aforementioned observable sample differences.[16] We consider instead the possibility that differences in the SSI estimates between the studies might be due to changes in the SSI program that may have coincided in different ways with the timing of the prices shocks in the two studies.

The most obvious candidate programmatic change was the establishment of the federal SSI program in 1974 intended to bring uniform eligibility standards and a national minimum benefit to programs that had previously varied greatly across states along these dimensions. Within one year, total beneficiaries in the federal program increased nearly 25% relative to participation in the state-level SSI antecedents, while total benefit payments increased by almost 60% (Social Security Administration 1975).[17] However, given heterogeneity in the administration of these programs across different states, the increase in benefit payments varied greatly across states and, quite plausibly, across counties within a given state.

The federalization of SSI in 1974 may have generated outlier observations in terms of county-level SSI benefit growth because of the rapid increase in total SSI payments. We thus re-examine the SSI results in both in ours and the Black et al. samples, after dropping the 1973 to 1974 first difference observations, which were directly affected by the change in the program

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[16] Recent work in the treatment effects literature compares estimates across settings by accounting for differences in observable characteristics with constant sub-group treatment effects (e.g., Hotz, Imbens, and Mortimer 2005; Angrist and Fernandez-Val 2013) or by estimating the distribution of treatment effects (e.g., Brinch, Mogstad, and Wiswall, Forthcoming). These papers focus on binary treatments as opposed to the continuous endogenous regressor that we use in our analysis and do not easily extend to the current context.

[17] The state-level antecedents to the SSI program were Old Age Assistance, Aid to the Blind, and Aid to the Permanently and Totally Disabled.

structure.[18] The first column of Table 3 shows our baseline estimate of TSLS estimate of the SSI elasticity with respect to earnings of −0.17 found in Table 1.[19] When we drop the 1973–74 first difference observations, our resulting estimate of −0.20, shown in column (2) of Table 3, is essentially unchanged from our baseline finding. Column (3) of Table 3 shows Black et al.'s original estimate of −0.713 which is found in column (2) of Table 3 in their paper. Our attempt at replicating their result, aided greatly by data provided by Dan Black, yields an estimate of −0.746 as shown in column (4) of Table 3.[20] When we omit first difference observations for 1973–74, the point estimate drops nearly in half to −0.382. While this estimate is still larger than our baseline SSI estimate, accounting for the 1974 federalization of SSI explains the majority of the difference between our SSI finding and Black et al's.

5 Conclusion

We find that worsening labor market conditions in a county cause disability benefit receipt there to increase. Our estimates of the elasticity of benefit payments to earnings during 1970–2011 for SSDI and SSI are −0.29 and −0.16, respectively. Earlier work finds a much larger SSI elasticity, but we show that much of the difference can be accounted for by the federalization of the SSI program in 1974.

Our finding that agents substantially substitute disability benefits for earnings when labor market circumstances worsen shows that this striking result, which was previously documented by Black et al., is not simply a feature of that earlier study's focus on a sample of less-educated

[18] Prior to 1974, the SSI measure used in the analysis contains payments to the state-level SSI antecedents.
[19] Following Black et al., we do not weight the analysis in Table 3.
[20] One likely reason we do not exactly match the results from Black et al. in our replication is that the REIS data is subject to periodic revisions which means that we do not necessarily have exactly the same data on county-level SSI payments, earnings, population, etc. that was used in the prior study.

men in rural Appalachian coal areas in the time period before 1990. Instead, we find a similar relationship in a more nationally representative set of states, among more-educated persons, and during a much more recent time period. This somewhat surprising result suggests that the response of disability take-up to market conditions is similar across areas and disability programs, despite different targeted population for SSDI and SSI programs.

Our analysis also raises important issues for future research. By focusing on changes in earnings and disability payments measured at the county level, our research design cannot separate disability take-up decisions based on individual earnings fluctuations from those due to movements in market-wide earnings levels. Separately identifying these two potential mechanisms can sharpen our understanding of how earnings changes affect disability take-up. Also, we find that the SSI response for the more recent time period is not significantly different than zero. This result is particularly puzzling as we still find significant responsiveness of SSDI benefits during the latter period. Additional work is needed to explain the divergence of the estimated elasticities between the two disability programs.

6 References

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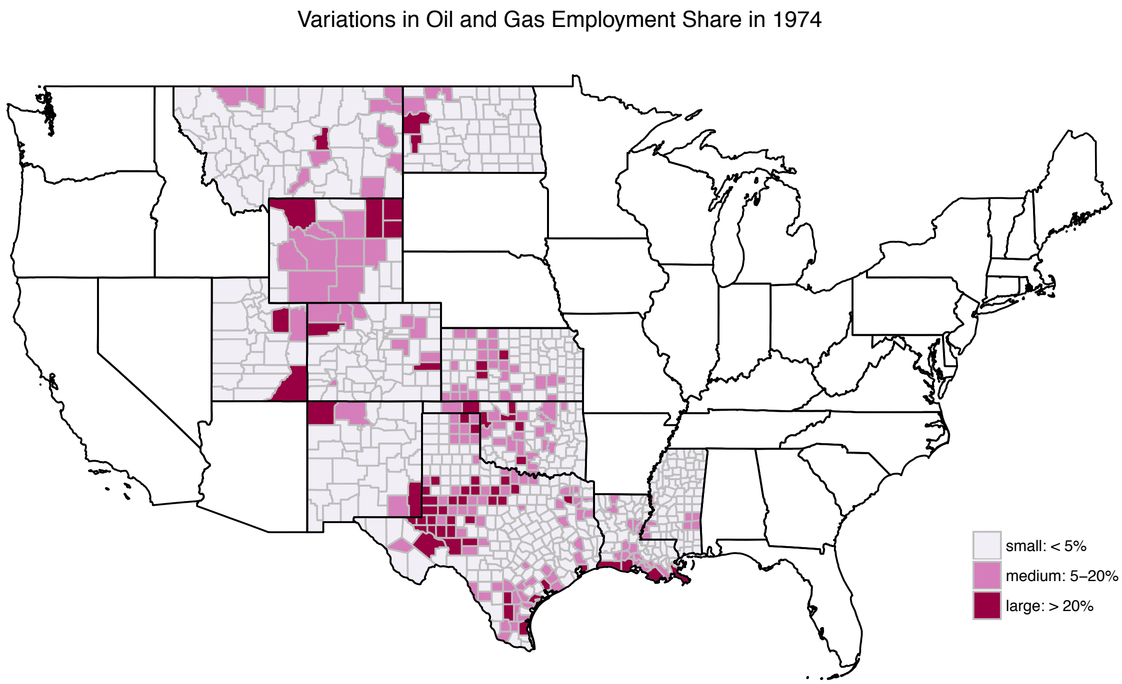
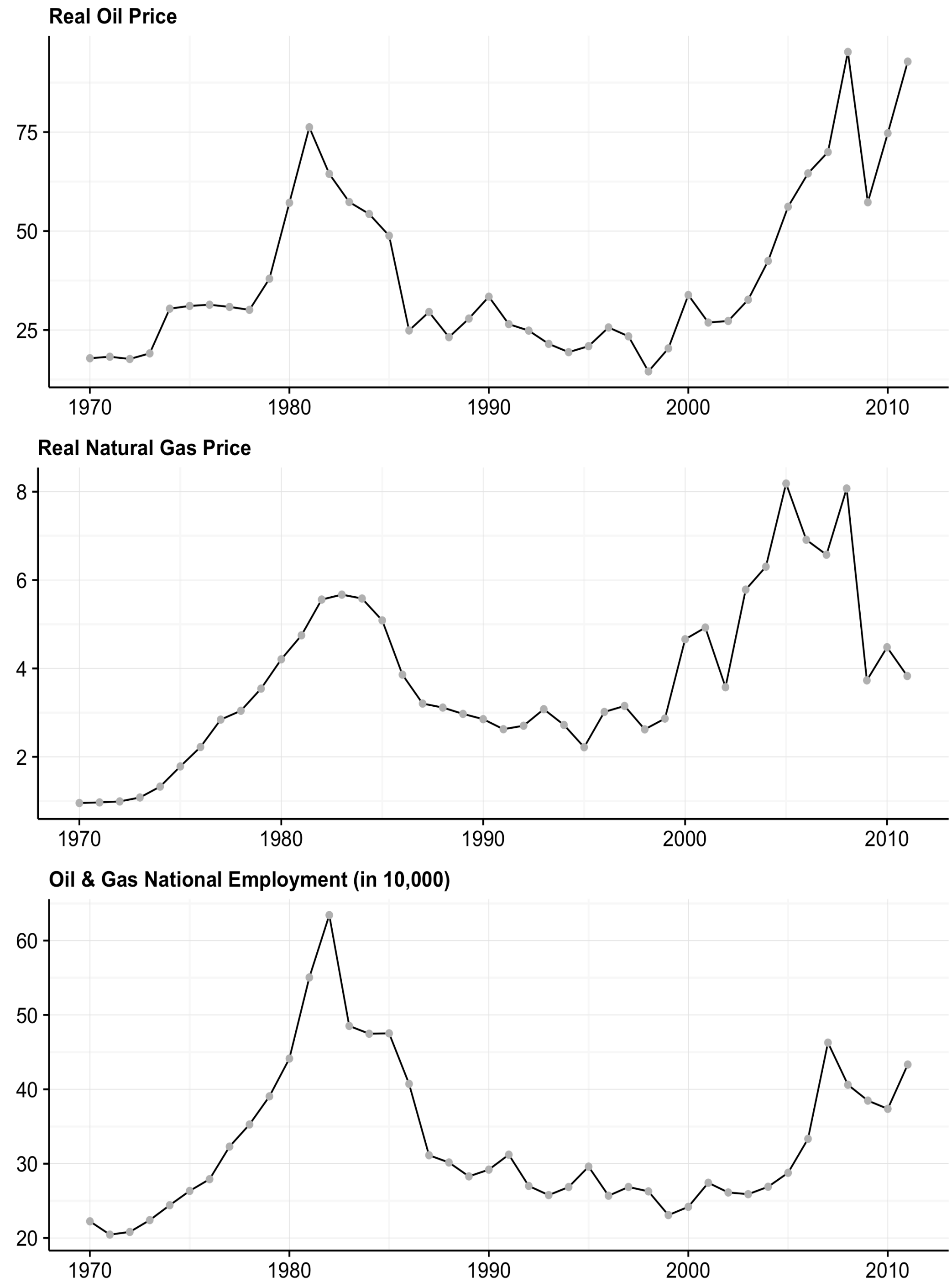
Figure 1: Share of County Employment in Oil and Gas Industry in 1974



Sources: Share of county employment in oil and gas industry is calculated based on data from CBP.

Figure 2: Effect of International Oil Supply Shocks on Prices and Employment



Sources: Prices are from the Energy Information Administration's Annual Energy Review. Oil prices are the US average first purchase price per barrel; natural gas prices are the wellhead price per thousand cubic feet. National oil and gas industry employment are from CBP.

Table 1: The Effect of Local Economic Performance on Disability Payments (1970–2011)

Method:	OLS	TSLS	OLS	TSLS
Economic Condition:	Δ Log Earnings	Δ Log Earnings	Δ Log Employment	Δ Log Employment
	(1)	(2)	(3)	(4)
		A. SSDI		
Economic Condition	.004	−.293	.079	−.699
	(.014)	(.069)	(.065)	(.131)
F-Statistic for Excluded Instruments		25.0		27.0
N	28,329	28,329	28,329	28,329
		B. SSI		
Economic Condition	−.019	−.160	−.001	−.360
	(.012)	(.079)	(.030)	(.168)
F-Statistic for Excluded Instruments		27.1		25.4
N	26,154	26,154	26,154	26,154

Notes: Each point estimate in the table represents results from a different regression. Standard errors in parentheses account for arbitrary forms of clustering within counties. All regressions control for: state*year fixed effects, MSA indicator (1990), log population, change in log population, and fraction of earnings from manufacturing (1969). SSDI balanced sample only includes counties with non-missing SSDI payments for all years except 1981, when all counties are missing. SSI balanced sample only includes counties with non-missing SSI payment for all years. All regressions are weighted by county population in the previous period.

Table 2: Robustness of Two Stage Least Squares Estimates

Specification:	1967 Instrument: 1970–1993 (1)	1967 Instrument: 1994–2011 (2)	1974 Instrument (3)	No Lags of Instrument (4)	Employment Instrument (5)
			A. SSDI		
Δ Log Earnings	−.287	−.200	−.272	−.147	−.414
	(.123)	(.091)	(.088)	(.126)	(.068)
F-Statistic for Excluded Instruments	36.9	12.8	17.6	67.3	23.9
N	15,225	13,104	28,329	28,329	28,329
			B. SSI		
Δ Log Earnings	−.203	.042	−.118	−.107	−.374
	(.087)	(.167)	(.068)	(.080)	(.082)
F-Statistic for Excluded Instruments	41.5	13.0	17.7	79	23.4
N	14,940	11,214	26,154	26,154	26,154

Notes: Each point estimate in the table represents results from a different regression. Standard errors in parentheses account for arbitrary forms of clustering within counties. All regressions control for: state*year fixed effects, MSA indicator (1990), log population, change in log population, and fraction of earnings from manufacturing (1969). SSDI balanced sample only includes counties with non-missing SSDI payments for all years except 1981, when all counties are missing. SSI balanced sample only includes counties with non-missing SSI payment for all years. All regressions are weighted by county population in the previous period.

Table 3: Impact of 1974 SSI Federalization on SSI TSLS Estimate

States/Shocks:	Oil/Gas	Oil/Gas	Coal	Coal	Coal
Years:	1970–1993	1970–1993	1970–1993	1970–1993	1970–1993
Method:	TSLS	TSLS	TSLS	TSLS	TSLS
Instruments:	Employment	Employment	Reserves	Reserves	Reserves
Include 1973–74?:	Yes	No	Yes	Yes	No
	(1)	(2)	(3)	(4)	(5)
Δ Log Earnings	−.165	−.195	−.713	−.746	−.382
	(.107)	(.108)	(.134)	(.140)	(.098)
F-Statistic for Excluded Instruments	41	36.4	26.5	24.3	22.7
N	14,940	14,318	7,904	7,904	7,577

Notes: Each point estimate in the table represents results from a different regression in which the outcome is the change in log county SSI payments. Standard errors in parentheses account for arbitrary forms of clustering within counties. All regressions control for: state*year fixed effects, MSA indicator (1990), log population, change in log population, and fraction of earnings from manufacturing (1969). SSI balanced sample for oil/gas regressions only includes counties with non-missing SSI payment for all years. Column (3) is taken from column (2) of Table 3 of Black, Daniel, and Sanders (2002). All regressions are unweighted.

Appendix Table 1: Characteristics of Workers by Industry

	All Workers			Oil & Gas Workers			Coal Workers		
	1970	1990	2010	1970	1990	2010	1970	1990	2010
Age									
25–29	14	16	11	15	15	14	11	8	10
30–39	24	32	22	26	39	25	21	39	21
40–49	26	25	26	29	24	24	30	31	23
50 and over	36	27	41	31	22	37	38	22	46
Education									
Less than high school	41	14	7	38	15	11	72	24	10
High school	34	34	34	30	35	44	23	49	59
Some college	12	28	24	13	26	22	4	20	21
College and above	14	24	35	18	25	24	1	7	10
Sex									
Male	62	54	52	88	82	85	98	94	94
Female	38	46	48	12	18	15	2	6	6
Race									
White	89	85	80	97	90	88	96	97	95
Black	10	9	9	2	4	4	3	2	2
Other	1	6	11	<1	6	8	<1	1	3
Occupation									
Managerial and Professional	22	27	34	24	27	27	4	8	10
Technical, Sales, and Administrative	26	30	28	18	20	15	4	7	7
Precision Production, Craft, and Repairers	14	12	10	48	34	41	67	55	57
Operatives and Laborers	22	16	12	9	18	15	24	28	22
Service	12	12	14	1	1	1	1	2	4
Farming, Forestry, and Fishing	4	3	3	<1	<1	<1	<1	<1	<1

Notes: Authors' compilations from the 1970, 1990 Public Use Micro Sample of the Decennial Censuses, and 2008–2012 ACS 5-year data files. Census's industry coding for oil and natural gas industry only include the extraction sector. Sample consists of individuals of age 25 and above who have worked positive weeks during the last year.

Appendix Table 2: Oil and Gas States

	Share of 1974 CBP State Employment in Mining	CBP Oil and Gas Share of all mining establishments in:	
		1974	1967
Wyoming	15.8%	77%	82%
New Mexico	7.8%	82%	76%
Louisiana	5.9%	88%	91%
Montana	4.3%	66%	68%
Oklahoma	3.9%	90%	92%
Utah	3.7%	47%	32%
Texas	3.2%	88%	92%
Colorado	2.2%	58%	50%
Kansas	1.8%	85%	87%
North Dakota	1.2%	65%	61%
Mississippi	1.1%	81%	83%

Notes: CBP stands for County Business Patterns.

Appendix Table 3: Selected First Stage Estimates of Impact on Earnings

States/Shocks:	Oil/Gas	Oil/Gas	Oil/Gas
Years:	1970–2011	1970–1993	1994–2011
Weighted?:	Yes	Yes	Yes
	(1)	(2)	(3)
		A. SSDI	
Instrument	.277	.315	.260
	(.033)	(.034)	(.049)
Instrument lagged once	.180	.197	.124
	(.028)	(.037)	(.028)
Instrument lagged twice	.043	−.152	.135
	(.019)	(.023)	(.027)
F-Statistic for Excluded Instruments	25.0	36.9	12.8
N	28,329	15,225	13,104
		B. SSI	
Instrument	.302	.357	.281
	(.034)	(.034)	(.052)
Instrument lagged once	.168	.197	.128
	(.027)	(.034)	(.030)
Instrument lagged twice	.027	−.177	.150
	(.020)	(.026)	(.030)
F-Statistic for Excluded Instruments	27.1	41.6	13.0
N	26,154	14,940	11,214

Notes: Standard errors in parentheses account for arbitrary forms of clustering within counties. All regressions control for: state*year fixed effects, MSA indicator (1990), log population, change in log population, and fraction of earnings from manufacturing (1969). SSDI balanced sample only includes counties with non-missing SSDI payments for all years except 1981, when all counties are missing. SSI balanced sample only includes counties with non-missing SSI payment for all years. All regressions are weighted by county population in the previous period.