Appendix. NOT FOR PUBLICATION

Appendix A: State UI Information

Information on state UI laws come from the *Significant Provisions of State Unemployment Insurance Laws*, published bi-annually by the US Dept of Labor, Employment and Training Administration. I consulted state laws and state employment agencies for more detailed information on benefit schedule variations[54].

Idaho

In Idaho, the fraction of highest quarter of earnings to compute the weekly benefit amount is $1/26$ for the whole period 1976 to 1984.

Maximum benefit amount

The maximum benefit amount in Idaho in January 1976 is $b_{max} = \$90$.
It was then increased seven times until December 1983:
\$99 for claims filed after 04jul1976
\$110 for claims filed after 01jul1977
\$116 for claims filed after 01jul1978
\$121 for claims filed after 01jul1979
\$132 for claims filed after 01jul1980
\$145 for claims filed after 01jul1981
\$159 for claims filed after 20jun1982.

Minimum benefit amount

The minimum benefit amount in Idaho in January 1976 is $b_{min} = \$17$.
It was then increased twice until December 1983:
\$36 for claims filed after 01jul1980
\$45 for claims filed after 01jan1984.

Duration of Benefits

Idaho has a special determination rule for potential duration described in table A1.

[54] CWBH has exhaustive information in Georgia on unemployment spells and wage records. But because of the parameters of the UI system in Georgia, the RK design was inoperable. $\tau_1 = 1/25$, $D_{max} = 26$, $\tau_2 = 1/4$ so that $D_{max} \cdot \frac{\tau_1}{\tau_1} > 4$ always larger than $\frac{bpw}{hqw}$ for all individuals on the left side of the benefit level kink. I don't have any observation with only kink in benefit level at the kink.

Table A1: Determination of Potential Duration 1st tier UI Idaho: 1976-1984

Ratio of bqw/hpw		UI Duration	
At Least...	Less Than...	before Jul 1st 1983	after Jul 1st 1983
1.25	1.50	10	
1.50	1.750	12	10
1.750	2.00	14	12
2.00	2.250	16	14
2.250	2.500	18	16
2.500	2.750	20	18
2.750	3.000	22	20
3.000	3.250	24	22
3.250	3.500	26	24
3.500	–	26	26

Louisiana

In Louisiana, the fraction of highest quarter of earnings to compute the weekly benefit amount is $1/25$ for the whole period 1979 to 1984.

Maximum benefit amount

The maximum benefit amount in Louisiana in January 1979 is $b_{max} = \$141$.
It was then increased four times until December 1983:
$149 for claims filed after 02sep1979
$164 for claims filed after 07sep1980
$183 for claims filed after 06sep1981
$205 for claims filed after 05sep1982

Minimum benefit amount

The minimum benefit amount in Louisiana from January 1979 until December 1983 is always $10.

Duration of Benefits

The fraction of base period earnings to determine the total amount of benefits payable for a given benefit year is 2/5. The maximum duration of benefits was set at 28 weeks. It was reduced to 26 weeks for claims filed after 03apr1983.

Missouri

In Missouri, the fraction of highest quarter of earnings to compute the weekly benefit amount is $1/20$ from the beginning of the period covered by the CWBh data (January 1978) until December 2nd, 1979 when it becomes .045.

Maximum benefit amount

The maximum benefit amount in Missouri in January 1978 is $b_{max} = \$85$.
It was then increased only once until December 1983:
\$105 for claims filed after02dec1979.

Minimum benefit amount

The minimum benefit amount in Missouri from January 1979 until December 1983 is always \$15.

Duration of Benefits

The fraction of base period earnings to determine the total amount of benefits payable for a given benefit year is 1/3. The maximum duration of benefits is 26 weeks for the whole period covered by the CWBH data.

New Mexico

In New Mexico, the fraction of highest quarter of earnings to compute the weekly benefit amount is $1/26$ for the whole period covered by the CWBh data (January 1980 to December 1983).

Maximum benefit amount

The maximum benefit amount in New Mexico in January 1980 is $b_{max} = \$106$.
It was then increased three times until December 1983:
\$105 for claims filed after02dec1979.
\$117 for claims filed after 01jan1981
\$130 for claims filed after 01jan1982
\$142 for claims filed after 01jan1983

Minimum benefit amount

The minimum benefit amount in New Mexico in January 1980 is \$22.
It was then increased to: \$24 for claims filed after 01jan1981
\$26 for claims filed after 01jan1982
\$29 for claims filed after 01jan1983

Duration of Benefits

The fraction of base period earnings to determine the total amount of benefits payable for a given benefit year is 3/5. The maximum duration of benefits is 26 weeks for the whole period covered by the CWBH data.

Washington

In Washington, the weekly benefit amount is computed as a fraction of the average of the two highest quarters of earnings. The fraction to compute the weekly benefit amount is $1/25$ for the whole period covered by the CWBh data (June 1979 to December 1983).

Maximum benefit amount

The maximum benefit amount in Washington in June 1st, 1979 is $b_{max} = \$128$.
It was then increased to:
$137 for claims filed after 25jun1979
$150 for claims filed after 06jul1980
$163 for claims filed after 01jul1981
$178 for claims filed after 01jul1982
$185 for claims filed after 01jul1983

Minimum benefit amount

The minimum benefit amount in in Washington in June 1979 is always $17.
It was then increased to: $41 for claims filed after 06jul1980
$45 for claims filed after 01jul1981
$49 for claims filed after 01jul1982
$51 for claims filed after 01jul1983

Duration of Benefits

The fraction of base period earnings to determine the total amount of benefits payable for a given benefit year is 1/3. The maximum duration of benefits is 30 weeks for the whole period covered by the CWBH data.
Note that until February 26, 1983, the state of Washington provides for 13 weeks of State-funded additional benefits for individuals who have exhausted their regular and Federal-State Extended Benefits[55]. However, no additional benefit period was paid while a Federal program was in effect.

[55] The additional benefits correspond to an *ad hoc* program which is triggered on only if the Governor determines it necessary.

EB trigger dates

Information on national and state triggers and trigger dates comes from the weekly trigger notice reports of the Bureau of Labor Statistics. Note that in the weekly trigger notice reports, there are sometimes some slight adjustments ex-post because of lags in the computation of the IUR triggers. I therefore rely on ex post trigger notices where the starting and ending dates of each episodes of EB are indicated.

National Trigger Dates

Until the Omnibus Budget Reconciliation Act of 1981, (effective July 1st 1981), the EB system had two triggers. A national trigger and state specific triggers. During the period 1976 to 1981, the national trigger was on three times, from 2/23/1975 to 7/2/1977, from 8/28/1977 to 01/28/1978, and from 7/20/1980 to 1/24/1981, automatically triggering periods of EB in all US states.

Idaho Trigger Dates

During the period 1976 to 1984, and on top of national EB periods, the EB trigger for Idaho was on four times: from 4/30/1978 to 7/29/1978, from 2/25/79 to 6/6/1979, from 2/17/80 to 7/18/81, and finally from 10/18/81 to the end of the period covered by the CWBH data.

Louisiana Trigger Dates

During the period 1979 to 1984, and on top of national EB periods, the EB trigger for Louisiana was on three times: from 7/20/1980 to 1/24/1981, from 9/12/1981 to 10/23/1982, and finally from 1/23/83 to the end of the period covered by the CWBH data.

Missouri Trigger Dates

During the period 1978 to 1984, and on top of national EB periods, the EB trigger for Missouri was on twice: from 6/1/80 to 7/25/1981, and from 3/26/1982 to 6/19/82.

New Mexico Trigger Dates

During the period 1980 to 1984, and on top of national EB periods, the EB trigger for New Mexico was on only once from 8/29/82 to 11/27/82

Washington Trigger Dates

During the period 1979 to 1984, and on top of national EB periods, the EB trigger for Washington was on without interruption from 7/6/1980 to 7/2/83.

Table A2: DESCRIPTIVE STATISTICS FOR FULL CWBH SAMPLE

	Idaho			Louisiana			Missouri			New Mexico			Washington		
	Mean	s.d.	N	Mean	s.d.	N	Mean	s.d.	N	Mean	s.d.	N	Mean	s.d.	N
	Duration Outcomes (wks)														
Initial spell	13.9	12.4	33365	14	10.6	34077	12.2	10.9	28665	14	12.6	27004	17.6	15.4	41992
wks UI paid	11.7	10.7	33365	13.8	10.4	34077	12.5	11.3	28665	13.4	12.8	27004	16.2	14.8	41992
wks UI claim	15.8	12.2	33365	15.1	10.4	34077	15.4	11.8	28665	15.8	12.6	27004	18.9	15.4	41992
	Earnings and Benefits ($2010)														
bpw	25136	22164	33365	26993	19446	34077	23733	17334	28665	23334	17132	27004	31232	20380	41992
hqw	9827	16405	33365	9581	6441	34077	8211	5830	28665	8252	5382	27004	8982	5321	41992
wba	262.4	86.3	33365	304.8	117.1	34077	225	51.4	28665	230	69.5	27004	286.7	94.7	41992
potential duration Tier I	20	5.5	33365	25	4.4	34077	22.1	5.2	28665	25.7	1	27004	27	4.2	41992
	Covariates														
age	30.2	12.7	33361	34.6	12.7	33850	34.8	12.7	28651	33.7	11.4	26924	34.2	11.9	41955
male	.666	.471	33361	.683	.465	33624	.609	.488	28663	.651	.477	27002	.627	.484	41972
educ. (yrs)	12	2.2	17774	11.4	2.7	31272	11.3	2.2	1867	11.7	2.5	26482	12.4	2.4	41702
dependents	2	1.6	18781	2	1.6	17325	2	1.6	21746	2.2	1.7	25534	1.7	1.5	28834
censored	.165	.362	33365	.128	.323	34077	.151	.382	28665	.162	.336	27004	.107	.289	41992

Notes: The initial spell, as defined in Spiegelman et al. [1992], starts at the date the claim is filed and ends when there is a gap of at least two weeks in the receipt of UI benefits. The duration of paid UI corresponds to the number of weeks a claimant receives unemployment compensation. The duration of a UI claim is the number of weeks a claimant is observed in the administrative data for a given unemployment spell. bpw is the base period earnings, and hqw is the highest quarter of earnings. wba is the weekly benefit amount of UI. Potential duration Tier I is the potential duration of the regular state UI program. In Missouri, information on education level is almost always unavailable.

Appendix B: Additional Results, Figures and Tables

RKD in Double-Difference

One main issue with the identifying assumptions of the RK design concerns the functional dependence between the forcing variable and the outcome of interest. It could be that the relationship between the forcing variable and the outcome is either kinked or quadratic. Then estimates are likely to be picking up this functional dependence between y and w_1.

A simple way to understand the issue is to remember the basic intuition behind the RK design. The model that I am interested in is $y = f(b, w_1, \varepsilon)$, where I want to get an estimate of f_1'. In this model, we have: $\frac{dy}{dw_1} = f_1' \frac{\partial b}{\partial w_1} + f_2' + f_3' \frac{\partial \varepsilon}{\partial w_1}$. The RKD assumes that f_2' and f_3' are the same on both sides of the kink (smoothness assumptions). Then, it follows that

$$\frac{\underset{k^+,k^-}{\Delta} \frac{dy}{dw_1}}{\underset{k^+,k^-}{\Delta} \frac{\partial b}{\partial w_1}}$$

identifies f_1', because $\underset{k^+,k^-}{\Delta} f_2' = 0$ and $\underset{k^+,k^-}{\Delta} f_3' = 0$.
If the assumption of smoothness in the functional dependence between the forcing variable and the outcome is violated, meaning that $\underset{k^+,k^-}{\Delta} f_2' \neq 0$ then, identification is not possible in the standard RKD. But if we have two sets of observations A and B for which we are willing to assume that $\underset{k^+,k^-}{\Delta} f_2'$ is the same, and for these two groups

$$\underset{k^+,k^-}{\Delta} \frac{\partial b}{\partial w_1}$$

is different, then f_1' is identified by α_{DD}, where:

$$\alpha_{DD} = \frac{\underset{A,B}{\Delta} \underset{k^+,k^-}{\Delta} \frac{dy}{dw_1}}{\underset{A,B}{\Delta} \underset{k^+,k^-}{\Delta} \frac{\partial b}{\partial w_1}} \tag{7}$$

Such an identification strategy is reminiscent of double-difference strategies. In practice it consists in comparing the change in slope at point k in the relationship between the outcome and the forcing variable for two identical groups of observations, but one of the two groups is subject to a kink in the schedule of b at k, and the other group is not.

To implement this strategy, the idea is to use the presence of variations in the maximum benefit amount over time, that shift the position of the kink across the distribution of the forcing variable (as shown in figure 1). The problem though is that, taken separately, each variation in max_b is too small to give enough statistical power to detect changes in slopes because the bandwidths are too small, and as previously pointed out, the drawback of the RKD is to be quite demanding in terms of bandwidth size. The idea therefore is to compare periods that are further away in time.

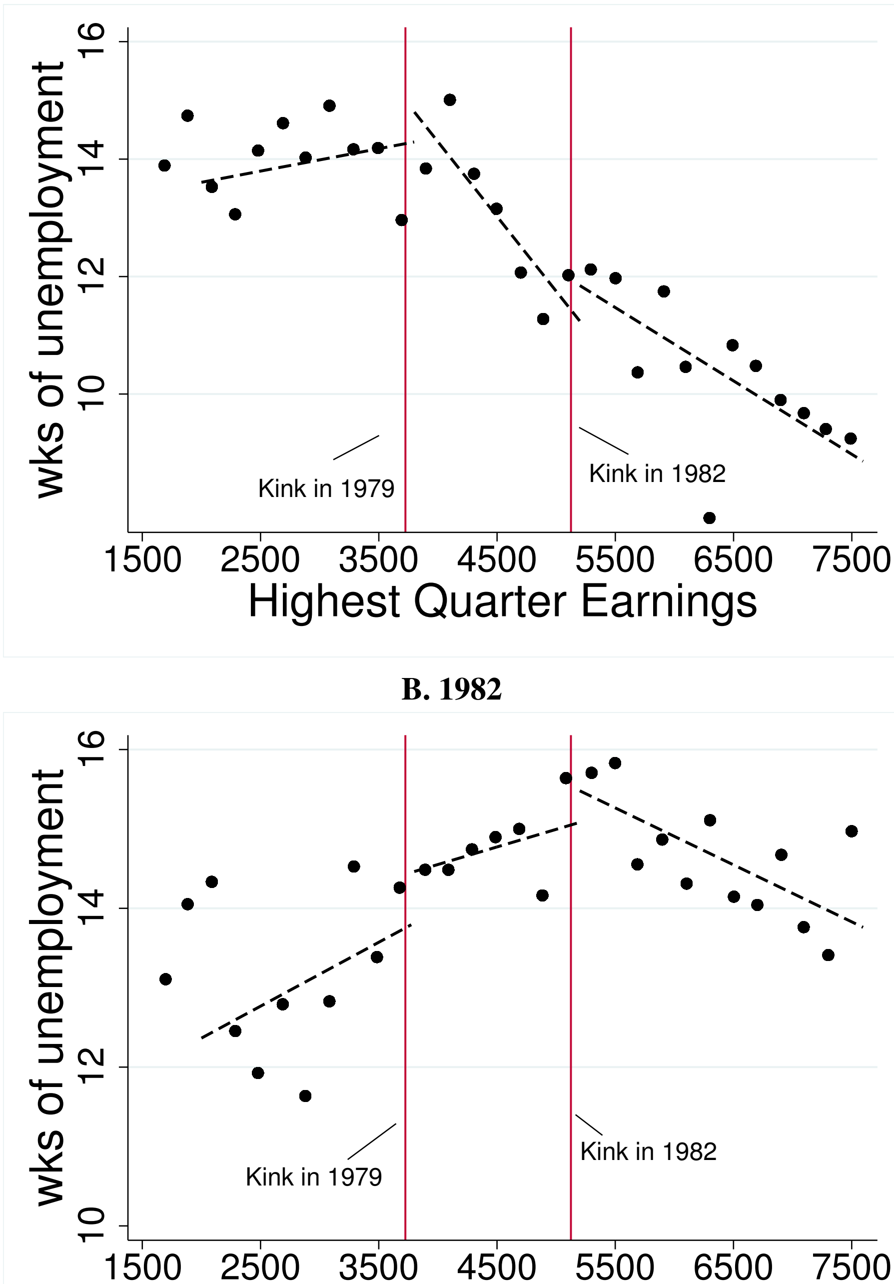
The obvious drawback of this option is that the identifying assumption is less likely to hold as one compares periods that are further away in time. In particular, one may worry about the high inflation rates during this period. It is important to note here that the maximum benefit amount increased in Louisiana a lot faster than inflation (40% between September 1979 and Sept 1982 and total inflation was less than 20% during that period), so that there is a clear and important change in the schedule in *real* terms [56]. Figure B1 shows the relationship between the duration of paid unemployment and the forcing variable in 1979 and 1982. Interestingly, there is a kink in this relationship in 1979 at the level of the 1979-kink in the schedule, and this kink disappears in 1982, when a new kink appears right at the level of the 1982-kink. Furthermore, in the interval between the 1979 and 1982 kinks, there is a change in slope in the relationship between the duration of unemployment and the forcing variable. This evidence is strongly supportive of the validity of the RK design.

[56] To further alleviate this concern, I also control for quadratic in *real* highest quarter of earnings in the DD-RKD specifications and find similar results.

Figure B1: RKD IN DOUBLE-DIFFERENCE USING VARIATIONS IN THE MAXIMUM BENEFIT LEVEL, LOUISIANA, 1979 VS 1982

A. 1979

B. 1982



Notes: The graph shows the average value of the duration of paid unemployment in each bin of the forcing variable in 1979 (panel A) and 1982 (panel B). The maximum benefit amount has been increased by more than 40% during the period, shifting the position of the kink in the schedule across the distribution of the forcing variable, as shown by the two red bars indicating the position of the kink for the two periods. The change in slope between the two periods in the interval between the two kinks is indicative of an effect of *b* on *y*, and can be used to identify the average treatment effect of *b* in a double-difference RKD. See text for details.

Table B1: DOUBLE-DIFFERENCE RKD ESTIMATES OF THE EFFECT OF BENEFIT LEVEL USING VARIATIONS IN THE MAXIMUM BENEFIT LEVEL, LOUISIANA, 1979 VS 1982

	(1)	(2)	(3)	(4)	(5)	(6)
	Duration of Initial Spell	Duration UI Claimed	Duration UI Paid	Duration of Initial Spell	Duration UI Claimed	Duration UI Paid
	A. 1979 Kink			**B. 1982 Kink**		
α_{DD}	.064	.088	.051	.065	.069	.05
	(.035)	(.035)	(.035)	(.034)	(.034)	(.034)
h_-	2500	2500	2500	1400	1400	1400
h_+	1400	1400	1400	2500	2500	2500
Opt. Poly	1	1	1	1	1	1
N	6495	6495	6495	4744	4744	4744

Notes: The table reports the results of the implementation of a Double-Difference RKD using variations in the maximum benefit amount over time, as described in the previous subsection. α_{DD} is the Double-Difference RKD estimate of the average treatment effect of benefit level as described in equation (7). It consists in comparing the change in slope at point k in the relationship between the outcome and the forcing variable for two identical groups of observations, but one of the two groups is subject to a kink in the schedule of b at k, and the other group is not. Standard errors for the estimates of α_{DD} are in parentheses. There are two sets of DD-RKD estimates, one for each kink. For the 1979-kink, I compare the change in slope in the duration of unemployment spells at the level of the 1979-kink in the forcing variable for the unemployed in 1979 (who had a schedule of benefit kinked at that point) against the unemployed in 1982 (who had a continuous schedule of benefits at that point). For the 1982-kink, I compare the change in slope in the duration of unemployment spells at the level of the 1982-kink in the forcing variable for the unemployed in 1982 (who had a schedule of benefit kinked at that point) against the unemployed in 1979 (who had a continuous schedule of benefits at that point). h_- and h_+ are the sizes of the lower and upper bandwidth. The optimal polynomial order is chosen based on the minimization of the AIC.

Placebo forcing variable

Another way to test for the existence of a kinked or quadratic functional dependence between earnings and unemployment duration is to use a placebo forcing variable. The placebo needs to be a good proxy for lifetime earnings, but must not be too correlated with the highest quarter of earnings that determines the benefit level. Table B2 explores the robustness of the RKD results by using the post unemployment wage as a placebo forcing variable instead of the pre-unemployment highest quarter of earnings. The post unemployment wage used is the wage for the first quarter of full employment after an unemployment spell. Post unemployment wages are available only for spells starting after September 1979 in Louisiana. Post unemployment wages are correlated with lifetime earnings but are not too much correlated with the highest quarter of earnings that determines the benefit level. Therefore, this table explores to what extent the baseline results are driven by some functional dependence between earnings and unemployment duration and shows that we cannot detect any effect in these placebo specifications using post unemployment wages as a forcing variable.

Table B2: ROBUSTNESS: RKD ESTIMATES OF THE EFFECT OF BENEFIT LEVEL USING POST UNEMPLOYMENT WAGE AS THE FORCING VARIABLE, LOUISIANA

	(1) Duration of Initial Spell	(2) Duration UI Claimed	(3) Duration UI Paid
	Sep 79-Sep 80		
α	-.024	-.022	-.02
	(.046)	(.045)	(.045)
Opt. Poly	1	1	1
	Sep 80-Sep 81		
α	-.025	-.019	-.019
	(.026)	(.026)	(.026)
Opt. Poly	1	1	1
	Sep 81-Sep 82		
α	.026	.031	.019
	(.034)	(.033)	(.033)
Opt. Poly	1	1	1
	Sep 82-Dec 83		
α	.01	.009	.005
	(.024)	(.024)	(.023)
Opt. Poly	1	1	1

Notes: The table explores the robustness of the RKD results by using the post unemployment wage as a placebo forcing variable instead of the pre-unemployment highest quarter of earnings. The post unemployment wage used is the wage for the first quarter of full employment after an unemployment spell. Post unemployment wages are available only for spells starting after September 1979 in Louisiana. Post unemployment wages are correlated with lifetime earnings but are not too much correlated with the highest quarter of earnings that determines the benefit level. Therefore, this table explores to what extent the baseline results are driven by some functional dependence between earnings and unemployment duration and shows that we cannot detect any effect in these placebo specifications using post unemployment wages as a forcing variable. α is the RK estimate of the average treatment effect of benefit level on the outcome. Standard errors for the estimates of α are in parentheses. The displayed estimates are for the optimal polynomial order chosen to minimize the Aikake Information Criterion.

Non-parametric tests for the the existence and location of a kink

An important concern in the RKD is that the estimates are picking up some spurious breakpoints in the relationship between the forcing variable and the outcome of interest. Despite their usually bad small sample properties, I recommend that non-parametric or semi-parametric tests for the detection and location of structural breakpoints are always performed when running RKD estimation, following the tests existing in the time series analysis literature, like for instance Bai and Perron [2003]. The number of tests that one can implement is large, but will usually fall within one of two categories. Tests for the existence of one or several breakpoints. And tests trying to detect the location of these breakpoints. By essence, testing for the statistical significance of the RKD estimates can be seen as falling into the first category. One could nevertheless envisage testing for the existence of more than one breakpoint, in order to make sure that the RKD estimates are not driven by the existence of multiple kinks in the relationship between the outcome and the forcing variable. An example of such tests can be found in Bai and Perron [1998].

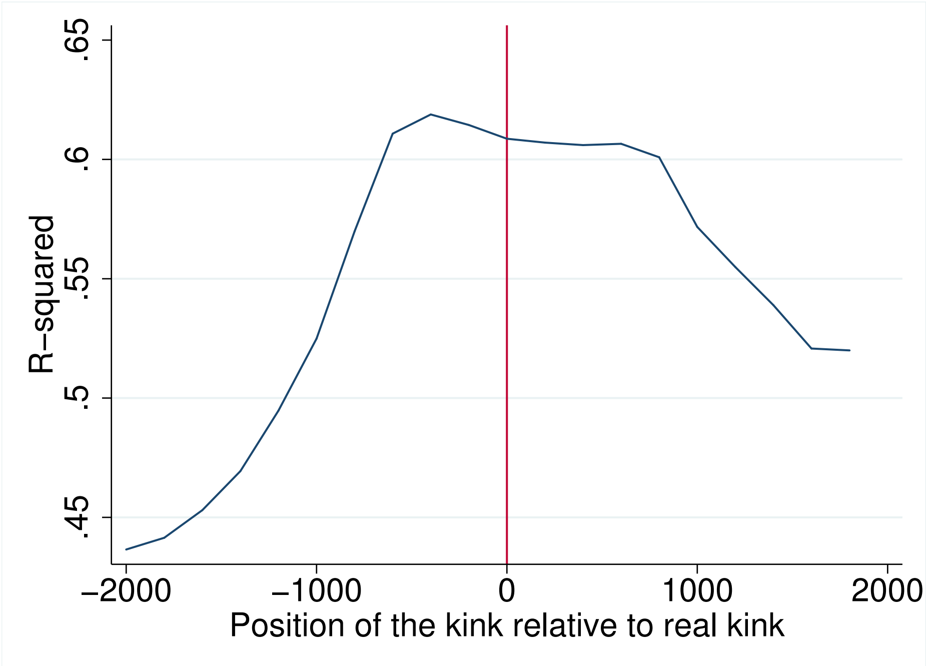
Here, I carry out a straightforward test that falls in the second category. I intend to make sure that the real location of the kink in the schedule is the location that would be detected if one were to look for the location of the kink in the data without knowing where the kink actually stands. The test simply consists in running the RKD specification of equation (5) for a large number of virtual kink points k, and then in looking at the kink point that minimizes the residual sum of squares or equivalently that maximizes the R-squared[57]. Because of the large variance of unemployment durations across individuals, I collapse the observations in bins of \$50 of the assignment variable in order to reduce the residuals sum of squares to begin with[58]. I report in figure B2 the evolution of the R-squared as I change the location of the kink point in specification (5). The evolution of the R-squared as one varies the location of the kink points provides evidence in support of the validity of the RKD design. For both periods, the R-squared increases sharply as one moves closer to the actual kink point and then decreases sharply, supportive of the existence of a kink around 0. For the first period, the kink point that maximizes the R-squared is situated \$370 to the left of the real kink point, but as one may infer from figure B2, one cannot actually reject the hypothesis that the kink point is actually at 0. For the second period, the kink point that maximizes the R-squared is situated \$200 to the right of the real kink point, but once again one cannot actually reject the hypothesis that the kink point is actually at 0. I interpret these results as strong evidence in support of the validity of the RK design.

[57] I conduct here a simple grid search but these tests can become computationally burdensome when looking for several breakpoints or for more complicated models, in which case the use of more efficient algorithms is recommended, as in Bai and Perron [2003]

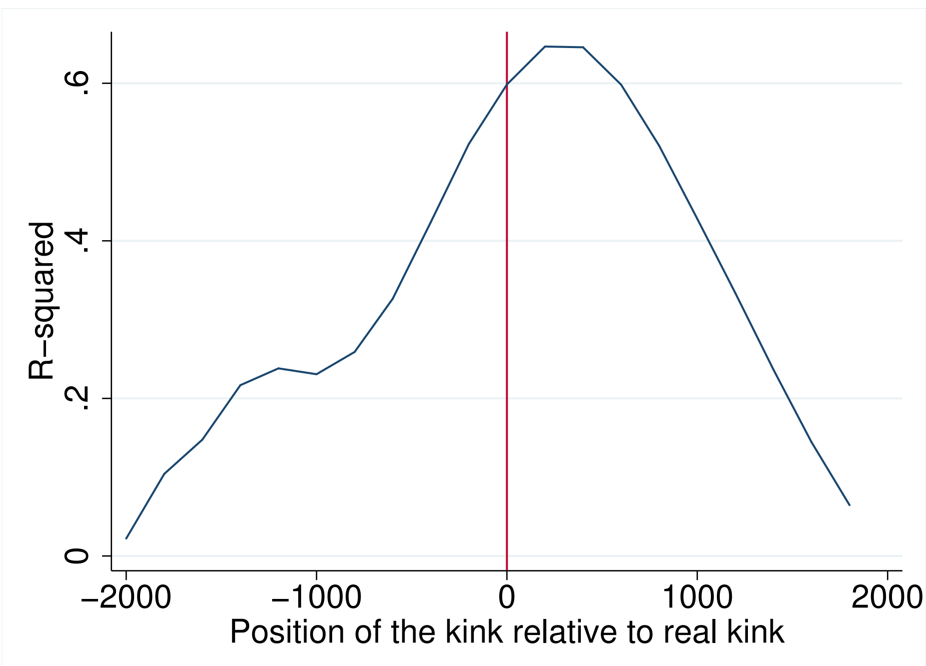
[58] This procedure increases the power of the test considerably.

Figure B2: R-SQUARED AS A FUNCTION OF THE LOCATION OF THE KINK POINT IN RKD SPECIFICATION (5), LOUISIANA

A. JAN 1979 - SEPT 1981



B. SEPT 1981 - DEC 1983



Notes: The graph shows the value of the R-squared as a function of the location of the kink point in RKD specification (5). The assignment variable is centered at the actual kink point in the benefit schedule so that virtual kink points are expressed relative to the real kink point in the schedule. Inspired by non-parametric tests for the detection of structural breakpoints in time series analysis, I conduct a grid search to look for the kink point that maximizes the R-squared. See text for details.

Proportional hazard models

To get a sense of the validity of the RK design, it is useful to compare the RKD estimates to the estimates of more standard empirical strategies widely used in the existing literature. Most empirical studies on US data use proportional hazard models. In table B3, I report the estimates of Cox proportional hazard models on the CWBH data which enables me to compare my results to the widely cited benchmark of Meyer [1990], who used a smaller sample of the same CWBH records.

This table estimates the effect of UI weekly benefits levels b on the hazard rate of leaving UI using the CWBH complete data for the 5 US states . I fit standard Cox proportional hazard models. All specifications include controls for gender, ethnicity, marital status, year of schooling, a 6-pieces exhaustion spline and state fixed effects. u denotes the state unemployment rate. log(b) denotes the log-weekly UI benefit amount. p25 and p75 denote the 25th and 75th percentile of unemployment rates (among all state$\times$quarter in our data).

Coefficient estimates for log(b) in the proportional hazard models can be interpreted as the elasticity of the hazard rate s with respect to the weekly benefit level. Under the assumption that the hazard rate is somewhat constant, these elasticities can be easily compared to the RKD elasticities of unemployment duration, since $D \approx 1/s$ so that $\varepsilon_D \approx -\varepsilon_s$.

Column (1) replicates the specification of Meyer [1990], Table VI, column (7). Note that Meyer [1990] was using a much smaller sample of the same CWBH records. The estimates show that the result of Meyer [1990], who found an elasticity of .56, can be fully replicated using his specification. The drawback of these estimates is that they do not fully address the endogeneity issue due to the joint determination of UI benefits and previous earnings. Meyer [1990] only controls for previous wages using the log of the base period earnings. Column (2) further adds non-parametric controls for previous earnings and experience. Column (3) further adds year$\times$state fixed effects. Interestingly, if one adds this richer set of non parametric controls for previous earnings to mitigate the concern of endogeneity, and fully controls for variations across labor markets by adding time fixed effects interacted with state fixed effects, the results converge to the RKD estimates and the elasticity goes down to around .3. The reason is that, as one controls more efficiently for the functional dependence between unemployment duration and previous earnings, the only identifying variation in benefit level that is left comes from the kink in the benefit schedule, and the model naturally converges to the identification strategy of the RKD. Overall, I find this evidence to be supportive of the validity of the RK design.

Columns (4) to (6) investigate the cyclicality of the partial equilibrium labor supply elasticities in the standard proportional hazard model to analyze the robustness of the results of table B4. Columns (4) and (5) add the interaction of log(UI) and high unemployment dummies (unemployment rate above the median across all US states in the same quarter in column (4) and unemployment rate above 8% in column (5)). Column (6) adds the interaction of log(b) with quartiles for the level of unemployment (quartiles defined across all state$\times$quarter cells in our sample).

Table B3: SEMI-PARAMETRIC ESTIMATES OF HAZARD RATES

	(1) Meyer [1990]	(2)	(3)	(4)	(5)	(6)
log(b)	-0.587***	-0.274***	-0.320***	-0.341***	-0.323***	
	(0.0394)	(0.0365)	(0.0368)	(0.0374)	(0.0370)	
State unemployment rate	-0.0550***	-0.0552***	-0.0207	-0.0226	-0.0251	-0.105***
	(0.00518)	(0.00519)	(0.0142)	(0.0143)	(0.0153)	(0.0209)
log(b)$\times$ (u>median)				0.0248**		
				(0.00812)		
log(b)$\times$(u> .08)					0.00527	
					(0.00685)	
log(b)$\times$(u<p25)						-0.363***
						(0.0376)
log(b)$\times$(p25<u<median)						-0.353***
						(0.0371)
log(b)$\times$(median<u<p75)						-0.292***
						(0.0371)
log(b)$\times$(u>p75)						-0.274***
						(0.0378)
Non-param controls for previous wage & experience	NO	YES	YES	YES	YES	YES
Year$\times$state F-E	NO	NO	YES	YES	YES	YES
# Spells	39852	39852	39852	39852	39852	39852
Log-likelihood	-136305.0	-136364.8	-135976.0	-135971.4	-135975.7	-135946.2

Notes: Standard errors in parentheses, * p<0.10, ** p<0.05, *** p<0.01.
This table estimates the effect of UI weekly benefits levels b on the hazard rate of leaving UI using the CWBH complete data for 5 US states from the late 1970s to early 1980s. I fit Cox proportional hazard models. All specifications include controls for gender, ethnicity, marital status, year of schooling, a 6-pieces exhaustion spline and state fixed effects. u denotes the state unemployment rate. log(b) denotes the log-weekly UI benefit amount. p25 and p75 denote the 25th and 75th percentile of unemployment rates (among all state$\times$quarter in our data). Column (1) replicates the specification of Meyer [1990], Table VI, column (7) (Meyer [1990] was using a much smaller dataset). Column (2) further adds non-parametric controls for previous earnings. Column (3) further adds year$\times$state fixed effects. Columns (4) and (5) add the interaction of log(b) and high unemployment dummies (unemployment rate above the median across all US states in the same quarter in column (4) and unemployment rate above 8% in column (5)). Column (6) adds the interaction of log(b) with quartiles for the level of unemployment (quartiles defined across all state$\times$quarter cells in our sample).

Cyclical behavior:

Following the Great Recession, a recent literature has been interested in estimating how labor supply responses to UI vary over the business cycle in order to assess the optimality of UI rules that are contingent on the state of the labor market (Schmieder et al. [2012], Kroft and Notowidigdo [2011]). I take advantage of the large variations in labor market conditions across states and over time in the CWBH data to investigate how the RKD estimates vary with indicators of (state) labor market conditions. I correlate the RKD estimates with the average monthly unemployment rate from the Current Population Survey prevailing in the state for each period[59]. Results are displayed in table B4. In all specifications, I weight the observations[60] by the inverse of the standard error (of the elasticity)[61]

Column (1) to (3) correlates the estimated elasticity with the unemployment rate for all three duration outcomes. In all three columns, the coefficient on the state unemployment rate is very small (around -.02 and not significantly different from zero), which means that a 1 percentage point increase in the unemployment rate is associated with a .02 percentage point decrease in the estimated elasticity. This result implies that elasticity varies between .38 (.09) when the state unemployment rate is at 4.5% (minimum in the CWBH data) and .25 (.10) when the unemployment rate is at 11.8% (the max in the CWBH data). This evidence is in line with the evidence of Kroft and Notowidigdo [2011] for the US, though the cyclicality of the estimates is somewhat larger in their analysis. One needs to acknowledge though that the standard errors on the estimated coefficient is rather large and the result of this type of exercise should always be interpreted with caution.

The estimates are not affected by the inclusion of state fixed effects as shown in column (4). In column (5), I add more observations by estimating the RKD model for subsets of the labor force in each state and sub-period. Here, I estimate the RKD elasticity for young (below 40) and old (above 40 years old) workers separately, but one can think of other partitions of the labor market, as long as: 1) unemployment rates can be computed for these sub-labor markets, 2) variation in unemployment rate across these sub-labor markets is large enough, and 3) each sub-labor market is large enough in order to estimate RKD elasticities with enough precision. Adding several estimates within state and sub-periods has two advantages. First, it increases the statistical power of the analysis, and more importantly, it enables me to control for the level of the policy parameters at which the elasticity is estimated. Each RKD elasticity is of course by nature endogenous to the level of the maximum benefit amount and the potential duration at which it is estimated, and these parameters vary for each state and sub-period. Results in column (5) show that partitioning the data into a larger number of sub-labor markets does not affect the result. The coefficient of the correlation between the unemployment rate in the sub-labor market and the RKD elasticity is still negative, and somewhat smaller in absolute value, though the amount of variation over time in each sub-labor market when controlling for sub-labor market fixed effects (here for age group

[59] To know to what extent variations in labor market conditions across states are a good proxy for business cycle fluctuations is another question. I tend to prefer in table B4 specifications with state fixed effects so that all variation in labor market conditions is variation over time, which mimics more clearly the concept of business cycles.

[60] Each observation is a RKD elasticity estimate of unemployment duration with respect to the UI benefit level for a state and sub period.

[61] Weighting reduces substantially the standard errors on the estimates of the correlation of the elasticity with labor market conditions, without affecting the point estimates.

fixed effects) is rather limited.

In table B3, columns (4) to (6) , I also investigate how the effect of the log benefit correlates with state unemployment conditions in the standard Cox proportional hazard model, and find similar results, with the estimated elasticity decreasing slightly as the state unemployment rate increases.

Construction of weights for the reweighted approach estimation in liquidity effects and moral hazard estimates

To make sure that our comparison of the effect of benefit level and potential duration using the two deterministic and kinked benefit schedules is not mixing heterogenous individuals, we re-weight the observations in the sample for the RKD estimates of $\left.\frac{\partial s_0}{db}\right|_B$ (sample 1) to match the distribution of observable characteristics of observations in the sample for the RKD estimates of $\frac{\partial s_0}{dB}$ (sample 2). To generate these weights, for each period, I merge observations from both samples. I then estimate a probit model of the probability that a given observation in this merged sample belongs to sample 1. The predictors in this regression are gender, age, age squared, education in years, and dummies for 5 main industries. Using predicted propensity score p, I then weight each observation in the RKD regressions with the weight $\omega = p/(1-p)$

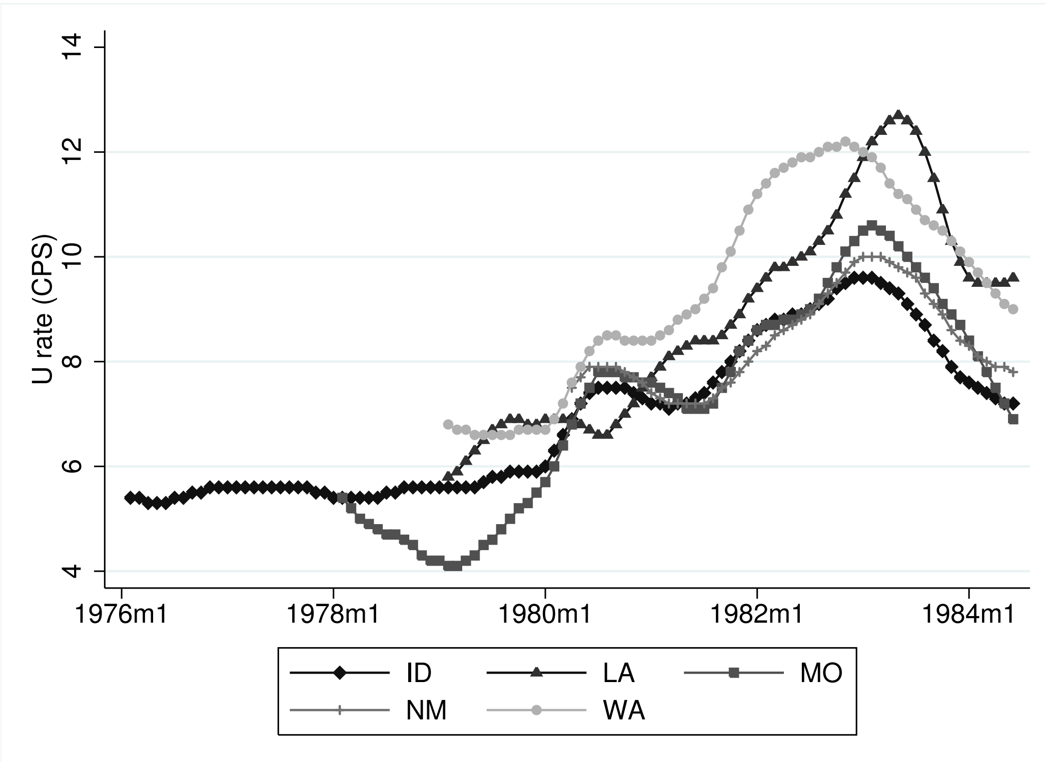
Table B4: CYCLICAL BEHAVIOR OF THE RKD ESTIMATES OF THE EFFECT OF BENEFIT LEVEL

	(1)	(2)	(3)	(4)	(5)
	Average Treatment Effects				
	ε_b	ε_b	ε_b	ε_b	ε_b
	Initial Spell	UI Paid	UI Claimed	Initial Spell	
U	-0.0195	-0.0293	-0.0259	-0.0289	-0.00576
	(0.0262)	(0.0263)	(0.0239)	(0.0303)	(0.0445)
Kink (K$2010)					-0.111
					(0.170)
Potential Duration					-0.00950
					(0.0177)
State F-E				×	×
Age Group F-E					×
Inverse s-e weights	×	×	×	×	×
N	26	26	26	26	52

Notes: Standard errors in parentheses, * p<0.10, ** p<0.05, *** p<0.01.
Each observation is a RKD estimate of the elasticity of unemployment duration with respect to the UI benefit level for a state and sub period. Initial spell refers to the elasticity of the duration of the initial unemployment spell as defined above. UI paid refers to the elasticity of the duration that UI is paid, and UI claimed refers to the elasticity of the duration of the UI claim. U is the average monthly state unemployment rate from CPS and in column (5) U is the average monthly state unemployment rate from CPS for each age group (the young, below 40, and the older workers, above 40 years old). Unemployment rates are expressed in percentage points, so that the results in column (1) for instance should be interpreted as follows: a 1 percentage point increase in the unemployment rate is associated with a .019 percentage point decrease in the estimated elasticity.

Additional Figures

Figure B3: UNEMPLOYMENT RATES IN CWBH STATES 1976-1984

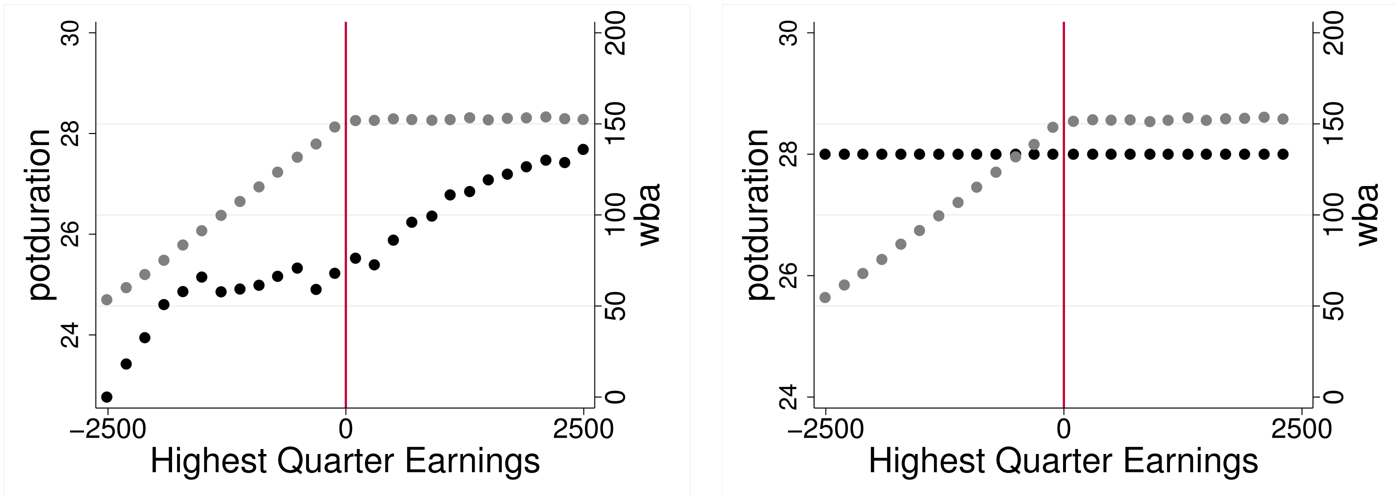


Sources: Current Population Survey
Notes: The graph shows the evolution of the monthly unemployment rate in the 5 states with the universe of unemployment spells available from the CWBH data. The CWBH data for the 5 states covers period of low unemployment as well as the two recessions of 1980 and 1981-82 with two-digit national unemployment rates, which gives the opportunity to examine the evolution of behavioral responses to UI over the business cycle.

Figure B4: UI BENEFIT SCHEDULE: WEEKLY BENEFIT AMOUNT (GREY) & POTENTIAL DURATION(BLACK), LOUISIANA

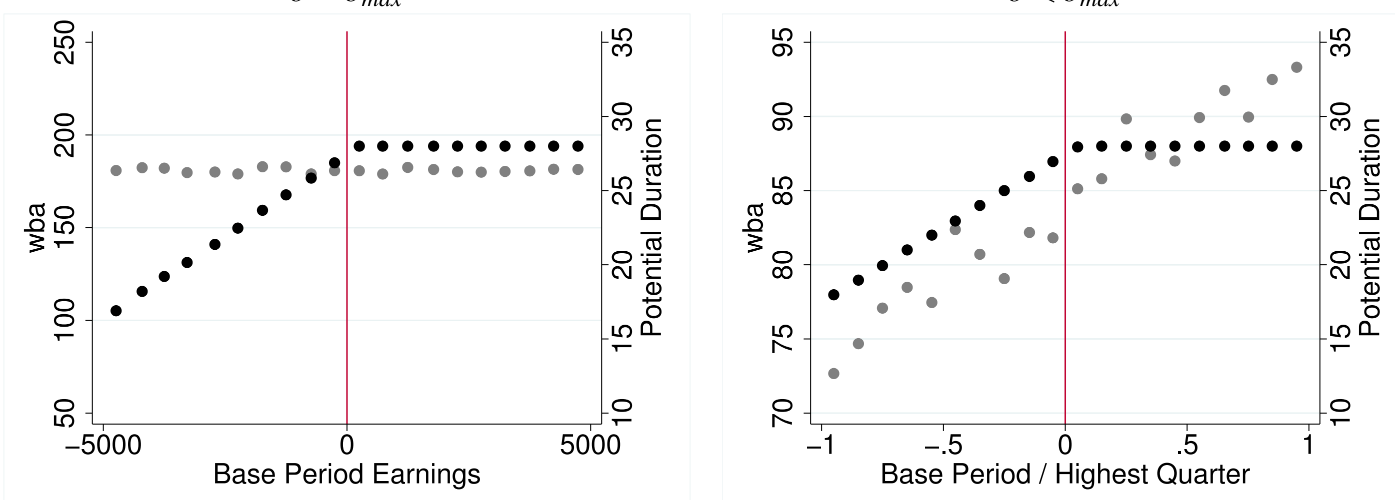
A. WBA as a kinked function of Highest Quarter Earnings

Full Sample $\qquad$ $D = D_{max}$



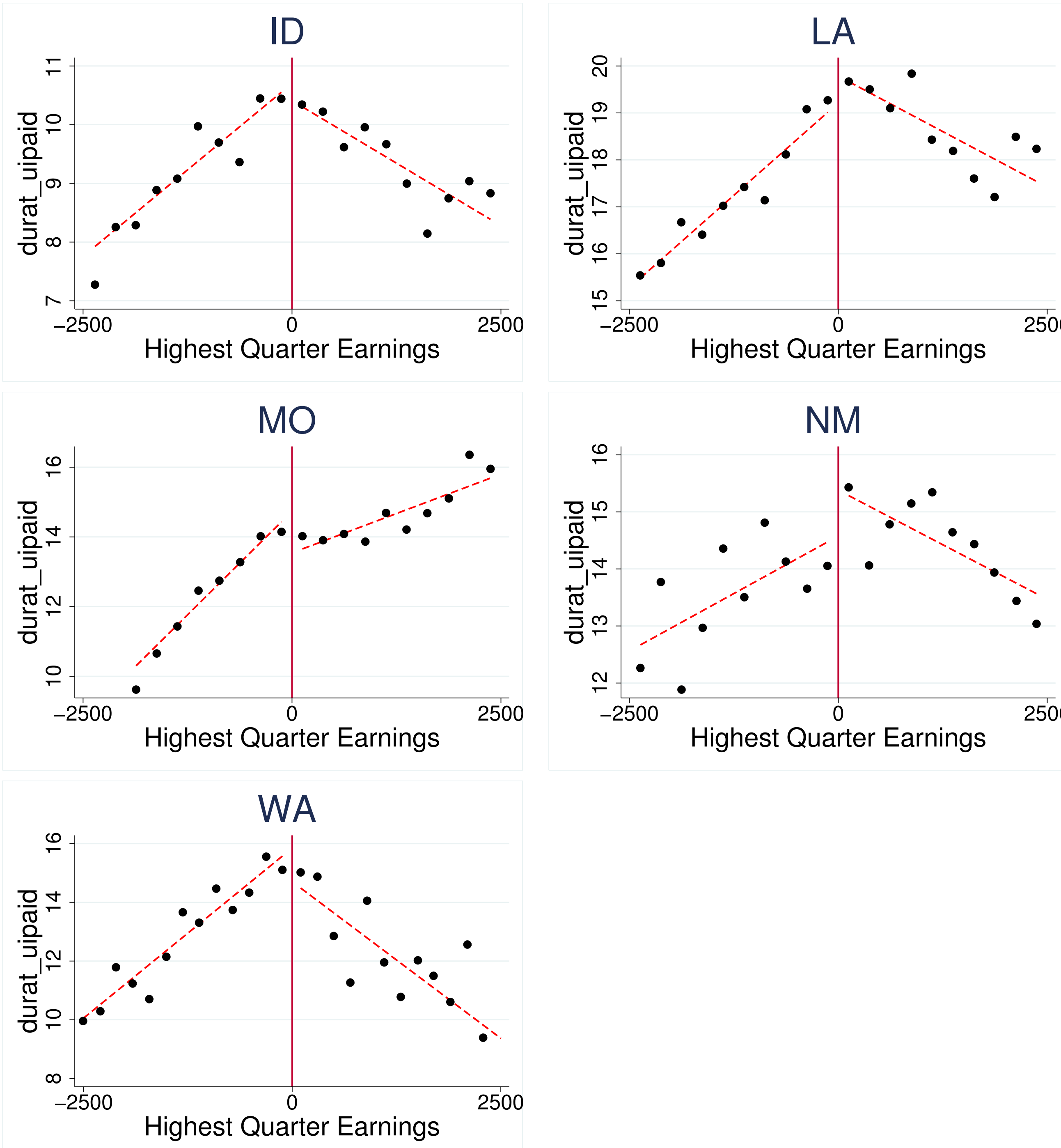
B. Potential Duration as a kinked function of Previous Earnings

$b = b_{max}$ $\qquad$ $b < b_{max}$



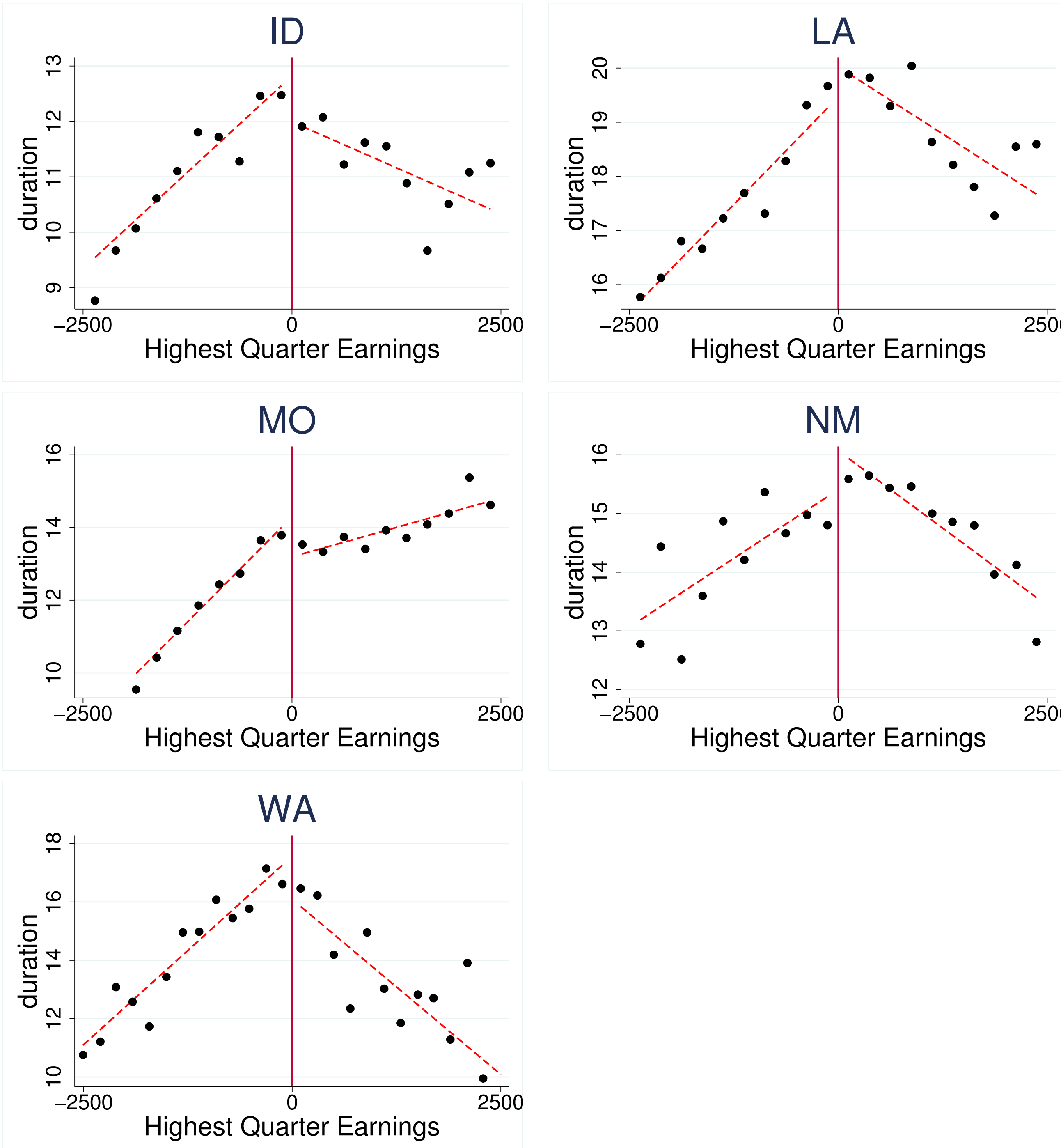
Notes: The graph shows the weekly benefit amount (wba: grey dots) and potential duration (potduration: black dots) of Tier I observed in the CWBH data for Louisiana for 1979 to 1983. Each dot is the average value in the corresponding bin of the assignment variable. Panel A shows that the weekly benefit amount is a kinked function of the highest quarter of earnings. Panel B shows that potential duration is a kinked function of the base period earnings for individuals with $b = b_{max}$ (left) and of the ratio of base period to highest quarter earnings for individuals with $b < b_{max}$ (right).

Figure B5: RKD EVIDENCE OF THE EFFECT OF BENEFIT LEVEL: DURATION UI PAID VS HIGHEST QUARTER EARNINGS FOR ALL 5 STATES



Notes: The graph shows for the first sub-period of analysis in each state the mean values of the duration of paid UI in each bin of $250 of highest quarter of earnings, which is the assignment variable in the RK design for the estimation of the effect of benefit level. The assignment variable is centered at the kink. The graph shows evidence of a kink in the evolution of the outcome at the kink. Formal estimates of the kink using polynomial regressions of the form of equation 5 are displayed in table 1. The red lines display predicted values of the regressions in the linear case allowing for a discontinuous shift at the kink.

Figure B6: RKD EVIDENCE OF THE EFFECT OF BENEFIT LEVEL: DURATION OF INITIAL UNEMPLOYMENT SPELL VS HIGHEST QUARTER EARNINGS FOR ALL 5 STATES



Notes: The graph shows for the first sub-period of analysis in each state the mean values of the duration of initial spell in each bin of $250 of highest quarter of earnings, which is the assignment variable in the RK design for the estimation of the effect of benefit level. The assignment variable is centered at the kink. The graph shows evidence of a kink in the evolution of the outcome at the kink. Formal estimates of the kink using polynomial regressions of the form of equation 5 are displayed in table 1. The red lines display predicted values of the regressions in the linear case allowing for a discontinuous shift at the kink.

Appendix C: Proofs and Results

Timing of the model: enter unemployment at period $t = 0$. At the beginning of every period, if the individual is still unemployed, she chooses search effort. Once search effort realized, she chooses consumption. The value function of finding a job at time t is:

$$V(A_t) = \max_{A_{t+1} \geq L} u(A_t - A_{t+1} + w_t - \tau) + \beta V(A_{t+1})$$

The value function of being unemployed at time t is:

$$U(A_t) = \max_{A_{t+1} \geq L} u(A_t - A_{t+1} + b_t) + \beta J(A_{t+1})$$

$$J(A_t) = \max_{s_t} s_t \cdot V(A_t) + (1 - s_t) \cdot U(A_t) - \psi(s_t)$$

s.t.

$$u(c_t^u) \geq 0$$
$$u(c_t^e) \geq 0$$

We assume that $\psi(.)$ is increasing and convex.

Optimal search:

$$\psi'(s_t) = V(A_t) - U(A_t) \tag{8}$$

Euler equations:

$$\forall t \quad u'(c_t^e) = \begin{cases} \beta u'(c_{t+1}^e) \\ u'(w - \tau) \ \text{ if } A_t = L \end{cases}$$

$$\forall t \quad u'(c_t^u) = \begin{cases} \beta[s_{t+1} u'(c_{t+1}^e) + (1 - s_{t+1}) u'(c_{t+1}^u)] \\ u'(b_t) \ \text{ if } A_t = L \end{cases}$$

Therefore, if the credit constraint is not binding at time t we have that:

$$\forall t \ \ u'(c_0^e) = \beta^t u'(c_t^e) \tag{9}$$

$$\begin{aligned} \forall t \ \ u'(c_0^u) &= \sum_{j=1}^{t} (\prod_{i=1}^{j-1} (1 - s_i) s_j) \beta^j u'(c_j^e) + \beta^t \prod_{i=1}^{t} (1 - s_i) u'(c_t^u) \\ &= \sum_{j=1}^{t} f_1(t) u'(c_0^e) + \beta^t S(t) u'(c_t^u) \\ &= F_1(t) u'(c_0^e) + \beta^t S(t) u'(c_t^u) \end{aligned} \tag{10}$$

where $f(t) = \prod_{i=0}^{t-1} (1 - s_i) s_t$ is the probability that the unemployment spell lasts exactly t periods and $f_1(t) = \prod_{i=1}^{t-1} (1 - s_i) s_t$ is the probability that the unemployment spell lasts exactly t periods conditional on being still unemployed at the beginning of period 1. Similarly, $\prod_{i=0}^{t} (1 - s_i) = S(t)$, is the survival rate at time t and $\prod_{i=1}^{t} (1 - s_i) = S_1(t)$is the survival rate conditional on being still unemployed at period 1. $F(t) = 1 - S(t) = \sum_{s=0}^{t} f(s)$ is the probability that the length of a spell is inferior or equal to t and $F_1(t)$ is the same probability conditional on being still unemployed at period 1.

Effect of benefit level at time t on optimal search:

$$\frac{\partial s_t}{\partial b_t} = -\frac{u'(c_t^u)}{\psi''(s_t)}$$

Effect of benefit level at time $t+j$ on optimal search at time t:

$$\frac{\partial s_t}{\partial b_{t+j}} = -\frac{\beta^j \prod_{i=1}^{j}(1-s_{t+i})u'(c_{t+j}^u)}{\psi''(s_t)}$$

We define the effect on any variable Z of a change in the constant benefit level b for a finite period of potential duration of UI benefits B as:

$$\left.\frac{\partial Z}{\partial b}\right|_B = \sum_{i=0}^{B-1} \frac{\partial Z}{\partial b_i}$$

Decomposition of the effect of an increase in benefit level at time t into the moral hazard and liquidity effects:
From 8, we have that:

$$\frac{\partial s}{\partial A_t} = \frac{u'(c_t^e) - u'(c_t^u)}{\psi''(s_t)}$$

$$\frac{\partial s}{\partial w_t} = \frac{u'(c_t^e)}{\psi''(s_t)}$$

so that:

$$\frac{\partial s}{\partial b_t} = \frac{\partial s}{\partial A_t} - \frac{\partial s}{\partial w_t} \tag{11}$$

which is the Chetty (2007) decomposition of the effect of benefits between the liquidity and moral hazard effect.

Similarly, the effect on search effort at time 0 of a change in the constant benefit level b for a finite period of potential duration of UI benefits B can also be written as the sum of two components, a moral hazard and a liquidity effect:

$$\left.\frac{\partial s_0}{\partial b}\right|_B = \overbrace{\left.\frac{\partial s_0}{\partial a}\right|_B}^{\text{liquidity effect}} - \underbrace{\left.\frac{\partial s_0}{\partial w}\right|_B}_{\text{moral hazard effect}} \tag{12}$$

where $\left.\frac{\partial s_0}{\partial a}\right|_B = \sum_{i=0}^{B-1} \frac{\partial s_0}{\partial a_i}$ is the effect of a change in the level of an annuity that pays \$$a$ every period and $\frac{\partial s_0}{\partial w}\Big|_B = \sum_{i=0}^{B-1} \frac{\partial s_0}{\partial w_i}$

Planner's problem:

The social planner chooses the UI benefit system to maximize expected utility subject to a balanced-budget constraint:

$$\max_{b,B,\tau} W_0 = (1-s_0)U(A_0) + s_0 V(A_0) - \psi(s_0)$$
$$\text{subject to} \quad D_B \cdot b = (T-D)\tau$$

Proof of proposition 1 in the case of benefit level b:

$$\frac{dW_0}{db} = (1-s_0)\left[\frac{\partial U_0}{\partial b}\Big|_B - \frac{\partial U_0}{\partial w}\Big|_B \frac{d\tau}{db}\right] + s_0\left[\underbrace{\frac{\partial V_0}{\partial b}\Big|_B}_{=0} - \frac{\partial V_0}{\partial w}\Big|_B \frac{d\tau}{db}\right]$$

From 8, we have that:

$$\forall y,\ \frac{\partial s_0}{\partial y}\Big|_B = \frac{1}{\psi''(s_0)}\left[\frac{\partial V_0}{\partial y}\Big|_B - \frac{\partial U_0}{\partial y}\Big|_B\right]$$

So that:

$$\frac{dW_0}{db} = -(1-s_0)\psi''(s_0)\frac{\partial s_0}{\partial b}\Big|_B - \frac{d\tau}{db}\left((1-s_0)\frac{\partial U_0}{\partial w}\Big|_B + s_0\frac{\partial V_0}{\partial w}\Big|_B\right) \tag{13}$$

We also know that: $\forall t,\quad \frac{\partial V_0}{\partial w_t} = \beta^t u'(c_t^e)$ so that :

$$\begin{aligned}\frac{\partial V_0}{\partial w}\Big|_B &= \sum_{t=0}^{B-1} \beta^t u'(c_t^e) \\ &= Bu'(c_0^e) \quad \text{if the credit constraint does not bind at time } B\end{aligned} \tag{14}$$

And, similarly: $\forall t,\quad \frac{\partial U_0}{\partial w_t} = \sum_{j=1}^{t} f_1(j)\beta^t u'(c_t^e)$ so that :

$$\begin{aligned}\frac{\partial U_0}{\partial w}\Big|_B &= \sum_{t=1}^{B-1} F_1(t)\beta^t u'(c_t^e) \\ &= \sum_{t=1}^{B-1} F_1(t)u'(c_0^e) \quad \text{if the credit constraint does not bind at time } B\end{aligned} \tag{15}$$

And therefore, if the credit constraint does not bind at time B

$$
\begin{aligned}
(1-s_0)\left.\frac{\partial U_0}{\partial w}\right|_B &= \sum_{t=1}^{B-1}(1-s_0)F_1(t)u'(c_0^e) \\
&= \sum_{t=1}^{B-1}F_0(t)u'(c_0^e) \\
&= (B-D_B-s_0)u'(c_0^e) \qquad (16)
\end{aligned}
$$

where we use the fact that $\sum_{t=0}^{B-1}S(t)=D_B$, the average duration of unemployment truncated at B.

Note that the moral hazard effect of an increase in b can also be expressed as a simple function of $u'(c_0^e)$ if the credit constraint is not binding at time B:

$$
\begin{aligned}
\left.\frac{\partial s_0}{\partial w}\right|_B &= \frac{1}{\psi''(s_0)}\Big[\left.\frac{\partial V_0}{\partial w}\right|_B - \left.\frac{\partial U_0}{\partial w}\right|_B\Big] \\
&= \frac{(D_B-s_0(B-1))u'(c_0^e)}{(1-s_0)\cdot\psi''(s_0)} \qquad (17)
\end{aligned}
$$

Using (12), (14), (16) and (17), we can rewrite (13) such that:

$$
\frac{dW_0}{db}=-(1-s_0)\psi''(s_0)\Big[\Big(\left.\frac{\partial s_0}{\partial a}\right|_B-\left.\frac{\partial s_0}{\partial w}\right|_B\Big)+\frac{d\tau}{db}\Big(\left.\frac{\partial s_0}{\partial w}\right|_B\cdot(B/(D_B-s_0(B-1))-1)\Big)\Big]
$$

We get from the government budget constraint that:

$$
\frac{d\tau}{db}=\frac{D_B}{T-D}(1+\varepsilon_{D_B}+\varepsilon_D\frac{D}{T-D})
$$

where $\varepsilon_{D_B}=\frac{b}{D_B}\frac{dD_B}{db}$ is the elasticity of the duration of paid unemployment with respect to the benefit level and $\varepsilon_D=\frac{b}{D}\frac{dD}{db}$ is the elasticity of the duration of total unemployment with respect to the benefit level.

Therefore, if the credit constraint is not yet binding at time B, the first-order condition $\frac{dW_0}{db}=0$ takes a simple form:

$$
1+\rho_1=(\frac{B}{D_B-s_0(B-1)}-1)\frac{D_B}{T-D}(1+\varepsilon_{D_B}+\varepsilon_D\frac{D}{T-D}) \qquad (18)
$$

where $\rho_1=-\frac{\left.\frac{\partial s_0}{\partial a}\right|_B}{\left.\frac{\partial s_0}{\partial w}\right|_B}$ is the liquidity to moral hazard ratio in the effect of an increase of benefit level. When the lefthand side of 18 is superior to the righthand side, it is socially desirable to increase the benefit level b, at the given level of potential duration B.

Proof of proposition 1 in the case of potential duration B:

To analyze marginal changes in B, I assume that a marginal change in the potential duration of benefits B normalized by the benefit amount b is therefore the same as a marginal change in b_B [62]. In this context, following the same logic as previously, we have that :

$$\frac{dW_0}{dB} = b.\frac{dW_0}{db_B} = b.\Big(-(1-s_0)\psi''(s_0)\Big[\big(\frac{\partial s_0}{\partial a_B} - \frac{\partial s_0}{\partial w_B}\big) + \frac{d\tau}{db}\big(\frac{\partial s_0}{\partial w_B}\cdot(1/(S(B)-s_0)-1)\big)\Big]\Big)$$

Differentiating the budget constraint of the government, we get that:

$$\frac{d\tau}{db_B} = \frac{1}{b}\cdot\frac{d\tau}{dB} = \frac{D_B}{B\cdot(T-D)}(\varepsilon_{D_B,B} + \varepsilon_{D,B}\frac{D}{T-D}) \tag{19}$$

where $\varepsilon_{D_B,B} = \frac{B}{D_B}\frac{dD_B}{dB}$ is the elasticity of the duration of paid unemployment with respect to the potential duration of UI benefits and $\varepsilon_{D,B} = \frac{B}{D}\frac{dD}{dB}$ is the elasticity of the duration of total unemployment with respect to the potential duration of UI benefits. Note of course that because $D_B = \sum_{t=0}^{B-1} S(t)$, we have that $\frac{\partial D_B}{\partial B} = \sum_{t=0}^{B-1}\frac{\partial S(t)}{\partial B} + S(B)$, which means that the effect of a change in potential duration on the actual average duration of UI benefits is the sum of the mechanical effect of truncating the distribution of spells at a later point in time $S(B)$ and a behavioral response. This point is central to the argument in Schmieder et al. [2012].

Using (19) and

$$1+\rho_2 = (\frac{1}{S(B)-s_0} - 1)\frac{D_B}{B\cdot(T-D)}(\varepsilon_{D_B,B} + \varepsilon_{D,B}\frac{D}{T-D}) \tag{20}$$

where $\rho_2 = -\frac{\frac{\partial s_0}{\partial a_B}}{\frac{\partial s_0}{\partial w_B}}$ is the liquidity to moral hazard ratio in the effect of an increase of potential duration. When the lefthand side of 20 is superior to the righthand side, it is socially desirable to increase the potential duration of benefits, at the given level of benefit level b.

Proof of corollary 1:

Consider the choice between two policies, a benefit extension and an increase in generosity that would relax the budget constraint of an equivalent amount so that $\frac{d\tau}{dB} = \frac{d\tau}{db}$. Given an equivalent relaxation of the budget constraint, the social planner will find it more desirable to increase the potential duration of benefit B if for $\frac{d\tau}{dB} = \frac{d\tau}{db}$ we have $\frac{dW_0}{dB} \geq \frac{dW_0}{db}$. The result of proposition 1 follows immediately from 2 and 1.

Proof of proposition 2:

[62] This is the case if B can potentially be increased by a fraction of period (a week in our case) and that if the potential duration B is not an integer number of periods, then, we can change b_t within a period such that the benefits in a given period is the fraction of the period that is covered time the benefit amount b.

Effect of increase in benefit level on exit rate at time 0 if potential duration=B:

$$\left.\frac{\partial s_0}{\partial b}\right|_B = \sum_{i=0}^{B-1} \frac{\partial s_0}{\partial b_i} = -\frac{u'(c_0^u)}{\psi''(s_0)} - \sum_{i=1}^{B-1} \frac{\beta^i S(i) u'(c_i^u)}{\psi''(s_0)}$$

Using Euler equation when borrowing constraint does not bind, we have that:

$$\left.\frac{\partial s_0}{\partial b}\right|_B = -\left\{ \frac{Bu'(c_0^u)}{\psi''(s_0)} - \sum_{t=1}^{B-1} \frac{F_1(t) u'(c_0^e)}{\psi''(s_0)} \right\} \tag{21}$$

Effect of an increase in potential duration scaled by the benefit level b, using Euler equation when borrowing constraint is not binding:

$$\frac{1}{b}\frac{\partial s_0}{\partial B} = \frac{\partial s_0}{\partial b_B} = -\left\{ \frac{u'(c_0^u)}{\psi''(s_0)} - F_1(B) \frac{u'(c_0^e)}{\psi''(s_0)} \right\} \tag{22}$$

Using 21 and 22, we have that:

$$\frac{1}{B}\left.\frac{\partial s_0}{\partial b}\right|_B - \frac{1}{b}\frac{\partial s_0}{\partial B} = (S(B) - \tfrac{D_B + s_0}{B}) \left\{ \frac{u'(c_0^e)}{(1-s_0)\psi''(s_0)} \right\} \tag{23}$$

The moral hazard effect of increasing benefit level b for B periods is given by (17) so that:

$$\frac{1}{B}\left.\frac{\partial s_0}{\partial b}\right|_B - \frac{1}{b}\frac{\partial s_0}{\partial B} = \Phi_1 \Theta_1 \tag{24}$$

where $\Phi_1 = \frac{S(B) - \frac{D_B + s_0}{B}}{D_B - s_0(B-1)}$